An investigation into the determinants of exchange rate volatility

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Abstract

Exchange rate volatility has significant effects on decisions made by many economic agents who participate in foreign exchange markets, most notably exporters, importers and foreign investors. The literature in the field of international macroeconomics has mainly concentrated on changes in the level of exchange rates rather than exchange rate volatility itself. Since we believe that the second moment of the exchange rate should be given more attention, we directly investigate the relationship between exchange rate volatility and macroeconomic fundamentals in developed as well as developing countries. For this reason, from the traditional exchange rate models which relate exchange rate levels to a set of fundamentals, we derive equations that can be used to examine the determinants of exchange rate volatility. We also investigate the possible impact of different variability measures and data frequencies. Our empirical results are generated from a very recently developed approach to cointegration analysis, namely, the bounds testing method of Pesaran et al., 2001. Using four industrialized countries and four less developed countries over the period 1973 to 1998, we found that the volatility of some macroeconomic fundamentals does indeed have a significant impact on the volatility of exchange rates in both groups of countries. Finally, whilst different variability proxies and data frequencies slightly affect the signs of significant variables, they do highly impact on the significance and weight that should be given to the relevant fundamental.

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Declaration

I confirm that no part of the material offered has previously been submitted by me for a degree in this or any other university. None of the work has been generated through joint work and in all cases the work of others has been acknowledged and quotations and paraphrases are suitably indicated.

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Chapter One

Introduction

In the 1950s and 1960s most international economists advocated floating exchange rates in preference to fixed exchange rates (see for example MacDonald, 1988). There were several arguments on which they relied in their advice to governments to adopt flexible exchange rate regimes. The first argument is related to the competitive position of a country in the international market. Friedman (1953) assumed that the internal price level is sticky in a downward direction. Thus, if a price level in a country rises, this makes this country uncompetitive and the balance of payments will suffer from a deficit if starting from a position of equilibrium. In order to keep equilibrium in the balance of payments, the country may use macroeconomic policy to reduce the price level, and because prices are sticky-downwards this may lead to painful adjustment and may result in welfare losses. It is better, then, to leave the exchange rate to depreciate to compensate for the rise in price level and to keep the competitive position of the country without a need to undergo such long and painful adjustment. The second argument of the proponents of flexible exchange rates is built upon the assumption that the stabilizing behaviour of speculators will make exchange rates relatively stable compared to fixed rates. Friedman (1953) for example, stated that under flexible regimes the exchange rate changes only in response to significant difficulties. That is to say, a change in the exchange rate comes after a difficulty has emerged and is delayed as long as possible and only after growing substantial pressure on the exchange rate. Thus, the direction of exchange rate change is likely to be known. For instance, if a currency depreciates from its long run value, speculators would know that the move is temporary, hence, would buy the currency since it is expected to appreciate in the future. Therefore, they stabilize the exchange rates' actual movements. The third claim by the proponents of the flexible exchange rate system is that this regime would protect a country from adverse external shocks. Under the Bretton Woods system a decrease in the demand for the exports of a country would cause a domestic contraction in this country. However, under a flexible exchange rate system the exchange rate would adjust to compensate for the shock, maintaining equilibrium in the current account and competitiveness and subsequently the level of demand. This merit of a floating exchange rate regime also gives the country the opportunity to exert an independent monetary policy. More precisely, under a fixed exchange rate an expansion in the money supply would lead to a deficit in the balance of payments which may cause a reversal result. On the other hand, under a floating exchange rate which keeps the balance of payments at equilibrium, monetary expansion is most likely to do the assumed job in the domestic economy. Moreover, as the flexible exchange rate keeps equilibrium in the balance of payments, there is no need to impose restrictions on trade by barriers and tariffs to equilibrate the balance of payments, and thus the flexible exchange rate should be associated with more liberal trade compared to fixed rates.

In the late 1960s and early 1970s, major countries suffered from fundamental current account deficits which fixed exchange rates were not able to deal with. This situation strongly pushed governments towards switching to flexible exchange rate regimes in 1973. As mentioned above, one of the justifications for a system of floating exchange rates is that they would be relatively stable compared to the Bretton Woods system. However, one of the outstanding features of the period of floating has been the high volatility of bilateral exchange rates. McKinnon (1976) stated that "*current* *movements of spot exchange rates of 20 per cent quarter-to-quarter, 5 per cent week-to-week or, even 1 per cent on an hour-to-hour basis are now not unusual, although they are very unusual by historical standards*". Macdonald (1988) has presented the minimum and maximum monthly percentage exchange rate changes for the French franc, German mark, Japanese yen, Swiss franc and UK pound against the US dollar. He has shown that the magnitude of exchange rates volatility is high and has not decreased as the experience with floating has increased.

It is believed that such large exchange rate movements create uncertainty about future proceeds/costs for all dealers who participate in international transactions, such as exporters, importers and foreign investors. Such uncertainty about future earnings makes planning for future production by companies participating in the foreign exchange market more difficult. Firms producing goods for export, for instance, often need to plan for future production relying on what they expect the exchange rate to be in the future. With low exchange rate volatility (even with high levels of exchange rates), planning becomes easier compared to planning in an environment of high volatility (even with low levels of exchange rates).

The economic theory of international trade (see for instance, Hooper and Kohlhagen, 1978; and Goture, 1985) states that exchange rate volatility creates uncertainty as to the prices importers would have to pay, or exporters would receive, in terms of their own home currencies, at some time in the future. If the participants in international trade are risk averse, they may prefer to switch to domestic activities where profits are relatively less uncertain rather than continuing trading in foreign markets where uncovered profits earned are highly uncertain as a consequence of exchange rate volatility. Alternatively, international traders may attempt to use forward foreign exchange markets in order to hedge against any possible losses. This can be successful for short term transactions and in mature forward markets. However, this would be at an extra cost and may be unavailable in most countries that have no mature markets of this kind (Hooper and Kohlhagen, 1978; and Gotur, 1985). In fact there has been a great deal of research that addresses the link between exchange rate volatility and trade flows. However, there is no unambiguous empirical evidence concerning this relationship as a result of contradictory findings.

As to the relationship between exchange rate variability and welfare, Obstfeld and Rogoff (1998) for example, pointed out that exchange rate volatility may be costly for welfare through two possible ways. Firstly, if a country's currency appreciates, this makes the goods of this country more expensive for foreign consumers, which leads to less demand and then less home production. Less production means less employment and hence less spending on consumption. Thus, appreciation reduces consumption. The opposite is true in the case of depreciation. Such fluctuations in the value of consumption do not favour the welfare of people. Secondly, higher volatility of exchange rates implies a higher risk premium which means higher prices; hence, driving consumption to be less than socially optimal. Moreover, investors who hold a substantial proportion of their wealth in terms of foreign assets will face large valuation effects on their wealth as exchange rates change. These possible valuation impacts cause international investors to devote time and spend resources in an attempt to minimize such effects on their wealth (MacDonald, 1988). It is also believed that exchange rate variability has significant impacts on many other micro and macroeconomic variables, such as external financial obligations, balance of payments, and production in non-traded sectors.

Consequently, we think that it is more relevant to examine the relationship between exchange rate volatility and macroeconomic fundamentals, rather than the determinants of exchange rate levels as most previous research has done. In fact in the field of finance a great deal of research has studied exchange rate volatility using time series models. However, in the area of international macroeconomics, little or no attention has been given to the variance of exchange rates. Therefore, this thesis aims to investigate the underlying factors determining exchange rate variability for a sample of developed and less-developed economies. Macdonald (1988) has stressed the importance of investigating whether exchange rate volatility is indicative of overshooting (by which he means exchange rate changes by more than the changes in the fundamentals), or if it is a natural response of exchange rates to the variability of the factors determining them. Such discrimination is crucial for policy makers who try to adopt the optimal policy.

In fact, most previous research about the behaviour of exchange rates has been devoted to explaining and forecasting exchange rate levels and not their volatility. Several structural models have been suggested to capture the pattern of exchange rates, such as monetary exchange rate models and portfolio balance models. However, none of these models was able to outperform a naïve random walk model in forecasting in sample exchange rate (see, for instance, Meese and Rogoff, 1983).

Chapter two of the thesis introduces one of the most commonly used structural models of exchange rates, that is the monetary approach to exchange rates. We present two versions of this approach. The first is known as the flexible price monetary model of exchange rates. This model assumes the validity of the purchasing power parity (PPP) and uncovered interest parity (UIP) hypotheses all times; thus, it is usually described as a long run model since these hypotheses are likely to hold in the long run, if at all. The second version is known as the real interest rate model which was a result of work by Dornbusch (1976) and Frankel (1979). This approach assumes that prices are sticky in the short term, and then the PPP is likely to be held only in the long run; hence, this model is usually referred to as the sticky price monetary model. The empirical research on these models was supportive in the early stages of the floating period, poor when the sample was extended to include the experience of the 1980s, and a bit more encouraging when recent cointegration analysis was applied. The main assumptions of the monetary models were blamed for their weak performance, especially during the second period of estimation. Chapter two also introduces the so-called Redux model of exchange rates based on the new wave of open economy macroeconomics pioneered by Obstfeld and Rogoff (1995). The theoretical developments in the Redux model have been quite rapid, but empirical work on this new model is limited.

Since we stress the importance of exchange rate volatility in international economics, chapter three addresses this issue. More details are given about the significance of exchange rate variability for private traders, investors and policy makers. One of the most difficult tasks in dealing with volatility is its measurement; thus, the main part of chapter three discusses this matter. A survey of the volatility proxies used in the literature is provided and their main advantages and disadvantages are highlighted. Finally, three candidate measures used in our empirical work are presented. These include a simple proxy which measures deviations from a long term average; thus, this proxy may better capture long term volatility. The second proxy is the most commonly used one in the literature; i.e. the standard deviation. This may better capture short term volatility since it measures deviations from the short term average. The third is an ARCH model-based proxy. This proxy takes account of time series properties such as heteroscedasticity and leptokurtosis, especially for financial data.

However, it is more useful for high frequency data, such as daily and weekly data, which has high degrees of noise (Abbott, 1999).

Chapter four introduces the empirical models to be used in investigating the underlying factors determining exchange rate variability. The empirical equations are simply modelled by taking the variance of both sides of the equations presented in chapter two, which are derived from the monetary approach and Redux models of exchange rates. Taking the variance of such equations assumes that exchange rate volatility is determined by the volatility of variables appearing in the right hand side. Since all the models introduced so far assume the PPP in the long run at least, we graphically examine the validity of PPP using both the absolute and relative versions for our sample countries. The results support the PPP hypothesis at least for the relative version.

Moreover, chapter four presents another model which explicitly addresses exchange rate variability. This model was introduced by Devereux and Lane (2003) and links the variability of bilateral exchange rates to a set of factors originally suggested by the optimum currency areas theory and some bilateral financial claims. In this chapter we also derive a hybrid model consisting of the transformed traditional models and Devereux-Lane model. Data sources, definitions of variables and symbols of the variables are also provided at the end of this chapter.

In chapter five of the thesis we outline the econometric methodology which can be employed to empirically test our developed models. Two cointegration approaches are outlined. The first is the Johansen multivariate cointegration method, which has the merit of accounting for the presence of more than one cointegrating relationship; however, it is only applicable for $I(1)$ variables. The second approach is the bounds testing method which was recently developed by Pesaran et al. (2001). This method can be applied irrespective of whether the order of integration of each variable is known to be $I(0)$ or $I(1)$, although it assumes a unique cointegrating relationship. It is still important to apply unit root tests to our data sets, even for the bounds method, to ensure that there is no variable whose order of integration is $I(2)$ or more. Thus, we also outline the unit root tests which are to be used with our time series data.

Chapter six provides the results of the unit root tests in which it was found that the variables involved in all models are either I(0) or I(1); hence, none of the variables appeared to be integrated of an order higher than one. These results rule out the use of the Johansen multivariate cointegration method; thus, the bounds testing procedure is proven to be a valid method in investigating the presence of cointegrating relationships in our equations. Furthermore, we followed the procedures needed to estimate a volatility proxy based on ARCH models. Using monthly data we found a few cases in which an ARCH effect exists, but almost none were found using quarterly data. Therefore, an ARCH model-based volatility proxy is estimated for the variables that contain an ARCH effect where monthly data are concerned.

The empirical results of investigating the existence of long run relationships in our models are given in chapter seven. Using the bounds testing approach the results indicate the existence of such relationships in most of the cases using monthly and quarterly observations and different volatility proxies. Thus, one can conclude that the volatility of some macroeconomic fundamental differentials are indeed long run forcing variables to exchange rates volatility, at least in our sample countries. This result implies that policy makers can to some extent affect exchange rate volatility by influencing the volatility of some economic fundamentals. Most of the cointegrating relationships are proven to be unique, which means that the bounds testing method can be used to derive the long run parameters of such relationships. Therefore, using the Delta method, the estimated long run coefficients are also provided in this chapter. In addition, using annual data for Devereux-Lane model, the presence of level relationships was only found in Algeria and Venezuela and using the augmented Devereux-Lane model we found a significant relationship only in Algeria.

In chapter eight we compare the results obtained: (1) using different models and the same data frequency and volatility proxy; (2) using different volatility proxies and the same model and data frequency; and (3) using different data frequencies and the same model and variability proxy. In other words, we examine the impact of using different models, volatility measures and data frequencies on the findings. We found some models to be most appropriate for each country, and some models to be preferred to others under some conditions for each country and for groups of countries. In general, we found that the sticky-price exchange rate monetary-based volatility equation works better than the other exchange rate model-based equations for both developed and less-developed countries. Moreover, we found that these results are sensitive to the volatility measures chosen and data frequency used. Furthermore, within the variables for which we found a significant impact on exchange rate variability, the volatility of inflation differentials was found to have a large effect on exchange rate variability compared to the other regressors, as indicated by its large estimated coefficients. The Devereux and Lane model, on the other hand, performed poorly, as it produced significant relationships in only two of the developing countries in our sample. The hybrid Devereux-Lane model generally resulted in no improvement to the results of the original Devereux-Lane model. In addition, explanations of the estimated parameters of our tested models and their policy implications are provided at the end of the chapter.

Chapter nine provides an overall summary and the conclusions of the thesis.

This thesis makes contributions in several different areas. Firstly, new models are developed in chapter four to explicitly investigate the underlying factors that determine exchange rate volatility. These models are built upon the existing models of the determination of exchange rate levels. Knowing the factors driving exchange rate variability is an important step for firms participating in foreign exchange markets and for policy makers so that they can plan for the future with more confidence. Secondly, we use different volatility measures and data frequencies with competing models to see how these affect the results, which may help in explaining the mixed results obtained in this study and in the literature which involves volatility in estimation. Thirdly, we apply our models to a sample of developed countries as well as to developing countries, which have been neglected in previous research. Finally, in our empirical work we have used quite recent methodology for testing the existence of long run relationships; namely, the bounds testing method.

Chapter two

Conventional Exchange Rate Determination Models

2.1 Introduction

The advent of the era of the generalized floating of the major international currencies in the early 1970s has produced a growing number of attempts to find a structural model capable of explaining the behaviour of exchange rates. This behaviour has been described by a high degree of volatility to the extent that it can be approximated by a random walk. Several structural models have been proposed in the literature to explain and forecast the movements of the exchange rates, but none has generated entirely satisfactory results. Virtually all of these structural models originated from the asset market theory of exchange rate determination. The asset market is commonly assumed to be efficient in this theory, which simultaneously means that expectations are rational, and thus the asset prices will reflect available information (Smith and Wickens, 1986). However, within asset theory there is no consensus about the relevant assets for exchange rate determination. Bilson (1978b) and Frankel (1979), for instance, by assuming foreign and domestic bonds to be perfect substitutes, focus on the excess supply of relative monies as determinants of exchange rates. In this narrowest case the exchange rate becomes a relative price of national monies, a case in which the emphasis is given to the capital account of the balance of payments instead of the current account. Foreign exchange, according to this theory, is considered as a financial asset and its price is determined by the demand and supply of foreign exchange stock. This view has been widely labelled the monetary model of exchange rates. Another asset market theory-based approach is the portfolio balance approach which is an extension of Tobin's financial framework to accommodate an open economy. In this model the assumption of perfect substitutability is relaxed allowing for current account imbalances to have a feedback effect on wealth and, subsequently, on long run equilibrium. Although the latter approach may be considered as a richer view of exchange rate determination, researchers have concentrated on the monetary approach in their empirical investigations of the asset view (MacDonald, 1984) due apparently to its empirical appeal and simplicity.

Therefore, this chapter outlines the theoretical foundations and the practical formulations of the monetary approach to exchange rates and its two versions, the flexible and sticky-prices models. A much newer model developed in the 1990s and labelled as the Redux model, is also explained.

2.2 The flexible-price monetary model

Since the exchange rate is the price of one country's money in terms of that of another, it is reasonable to analyze the determinants of that price in terms of the stock of and demand for the two monies. This rationale constitutes the main argument of the monetary approach to exchange rates (see for example, Frankel, 1976; and Mussa, 1976).

The flexible-price monetary model of exchange rate determination assumes a strong relationship between the nominal exchange rate and a simple set of monetary fundamental and relies on certain assumptions which are outlined as follows:

1-Purchasing power parity (PPP) holds continuously. The absolute version of the PPP hypothesis presumes that goods prices are completely flexible at home and abroad and the transaction costs are negligible. Thus, the general level of prices, when converted to a common currency, will be the same in every country. That is,

$$s_t = p_t - p_t^* \tag{2.1}$$

where s is the log of spot exchange rate measured as the price of a unit of the foreign currency in terms of the domestic currency units[1], p is the price level in log form, t denotes time and an asterisk denotes a foreign magnitude .

2-Stable money demand functions at home and abroad. The money demand is a function of real income and interest rate, and prices are determined by monetary equilibrium in both countries.

$$m_t - p_t = \alpha_1 y_t - \alpha_2 r_t \tag{2.2}$$

$$m_t^* - p_t^* = \alpha_1^* y_t^* - \alpha_2^* r_t^* \tag{2.3}$$

where m is the nominal money stock, y is the real income and r is the nominal interest rate. All variables, except r, are in logarithm form, therefore, α_1, α_2 are the income elasticity and interest rate semi-elasticity of money demand.

In this equation, which is based on the familiar Cagan-style money demand function, it is assumed that money is homogeneous of degree 1 in prices.

[1] We have chosen this definition since it is widely used in the literature. It implies that an increase in the exchange rate represents depreciation, and a decrease in the exchange rate represents an appreciation.

3-Uncovered interest parity (UIP) holds in the exchange market. This assumption implies that domestic and foreign assets are perfect substitutes, there is perfect capital mobility and economic agents are risk neutral (or risk is entirely avoidable) and rationally form their expectations of future exchange rates. Moreover, the interest rate is assumed to be determined in the world markets in the long-run. This hypothesis states that in the case of certainty or risk neutrality the expected change in exchange rate equals the interest rate differential:

$$\Delta s^e = r - r^*.$$

From (2.2), (2.3) we can get:

$$p_t = m_t - \alpha_1 y_t + \alpha_2 r_t \qquad (2.4),$$

$$p_t^* = m_t^* - \alpha_1^* y_t^* + \alpha_2^* r_t^* \qquad (2.5)$$

p^* is determined by the world money supply, thus, it is exogenous for the domestic country. The domestic price p is determined at home by the domestic money supply. Hence the exchange rate is determined by relative money supplies. Therefore by substituting 2.4, 2.5 into (2.1) yields:

$$s_t = m_t - m_t^* - \alpha_1 y_t + \alpha_1^* y_t^* + \alpha_2 r_t - \alpha_2^* r_t^* \qquad (2.6)$$

It is commonly assumed in the monetary model that the money demand parameters are identical at home and abroad, which imposes the following restrictions on equation (2.6); $\alpha_1 = \alpha_1^*$ and $\alpha_2 = \alpha_2^*$. Therefore, Eq (2.6) can be written as:

$$s_t = (m_t - m_t^*) - \alpha_1 (y_t - y_t^*) + \alpha_2 (r_t - r_t^*) \qquad (2.7)$$

Imposing such restrictions can be justified on the basis of the presence of multicollinearity problem, in the case of which the efficiency of coefficient estimates will be increased. Nevertheless, this may lead to biased estimates and sign reversal (see for example, Odedokun, 1997; and MacDonald and Taylor, 1992). In fact, the validity of these constraints is an empirical question. The other restriction imposed on Eq (2.7) is that the coefficient on (*m-m**) is unity (neutrality of money), which means that there is a proportional relationship between nominal exchange rate and relative money supply. That is, an increase of 1% in domestic money supply with constant foreign money stock is followed by a rise of 1% in the exchange rate. This comes from the assumption that money demand is homogenous of degree 1 in prices.

Some econometric researchers relax these imposed constraints, which yields the following form:

$$s_t = \gamma_1 m_t - \gamma_2 m_t^* - \gamma_3 y_t + \gamma_4 y_t^* + \gamma_5 r_t - \gamma_6 r_t^* \qquad (2.7a)$$

Equation (2.7) is sometimes simplified by excluding the interest rate differential term (as in Rapach and Wohar, 2002, 2004).

Expression (2.7) represents the flexible price monetary model view, and it implies that a rise in home money supply relative to the foreign one induces domestic economic agents to get rid of the excess money by spending more on goods and services, which, in turn, drives prices up and, through the PPP condition, leads to a depreciation in the domestic currency (s increases), whereas a rise in the domestic real income relative to its foreign counterpart raises the demand for the domestic money stock which, in turn, reduces expenditure. Therefore, the home price level falls until the money market is cleared. The fall in the domestic price level with constant foreign price level implies an appreciation in the home currency (s decreases) according to PPP theory. Finally, an increase in the domestic interest rate relative to the foreign one reduces the demand for domestic money and then causes a depreciation of the domestic currency, which is an increase of the exchange rate (s).

Given that the three assumptions at the core of the flexible monetary model-namely stable money demand functions at home and abroad, purchasing power parity, and uncovered interest parity-are unlikely to hold continuously, this model should be viewed as a long run or steady state model of exchange rate determination (MacDonald, 1984).

2.3 The sticky-price real interest monetary model

Under continuous PPP, the real exchange rate does not change. In the real world however, the real exchange rate is highly changeable. Dornbusch (1976) therefore, evolved a sticky-price model which allows for a considerable overshooting of both nominal and real exchange rates. Dornbusch assumes that exchange rates and interest rates jump to compensate for the stickiness of other prices, mainly goods prices. The intuition behind Dornbusch's idea is as follows. In the short-run goods prices are sticky; therefore, a fall in the money supply implies a fall in the real money supply which, in turn, raises the interest rate to keep the monetary market in a position of equilibrium. The increase in the domestic interest rate encourages capital inflows which lead to an appreciation in the nominal exchange rate. Given sticky prices, this means an appreciation of the real exchange rate as well. Foreign investors know that they are artificially forcing up the exchange rate and they may be subject to a loss in foreign exchange when they convert their earnings into their domestic currency. Hence, they expect the exchange rate to depreciate. However, investors will continue to buy that country's assets as long as their capital gain (the interest differential) is higher than their foreign exchange loss (the expected rate of depreciation). When equality between the interest rate differential and the expected rate of depreciation holds, short-run equilibrium is achieved; that is, uncovered interest parity holds. Thus, for a nonzero interest differential, the expected rate of depreciation must be nonzero. Consequently, the exchange rate must have overshot its long-run equilibrium level (PPP).

In the longer term and as a response to the money supply decrease, the domestic prices start to fall which results in a decline in the domestic interest rate. Then, the exchange rate depreciates slowly towards its long-run level.

Dornbusch (1976) has emphasized the role of expectations in determining exchange rate behaviour, and formulate this in the following expectations mechanism:

$$\Delta s^e = \Phi(\bar{s} - s) \qquad \Phi > 0 \qquad (2.8)$$

where Δs^e is the expected change in the exchange rate, $\bar{s}$ is the log of long-run exchange rate and Φ reflects the sensitivity of market expectations to the (proportionate) over- or undervaluation of the currency relative to equilibrium or the speed at which the gap between the spot exchange rate and its long-run equilibrium counterpart is expected to close.

Frankel (1979) has assumed that the expected change in the exchange rate is a function of the gap between the current spot rate and the equilibrium rate, and of the expected long run inflation differential between the domestic and foreign countries. This fundamental assumption can be written as follows:

$$\Delta s^e = \Phi(\bar{s} - s) + (\pi^e - \pi^{e*}) \qquad (2.9)$$

where π^e , and π^{e*} are the current rates of expected long run inflation rates at home and abroad respectively.

The log of the equilibrium exchange rate, $\bar{s}$, is defined to rise at the rate $\pi^e - \pi^{e^*}$, in the absence of new shocks. More precisely, equation (2.9) tells us that in the short run the exchange rate is anticipated to revert to its long run value at a rate which is proportional to the current gap, and in the long run when $s = \bar{s}$ it is expected to change at the long run rate $\pi^e - \pi^{e*}$. Frankel (1979) justified equation (2.9) to be a reasonable form for expectations taken by participants in an inflationary world.

This shows that when the exchange rate is at its equilibrium level, it is not necessary to stay constant, but it is expected to depreciate by the difference between the expected domestic and foreign inflation rates.

From uncovered interest parity (UIP), which states that the expected change in exchange rate equals the interest rate differential; $\Delta s^e = r - r^*$, equation (2.9) becomes:

$$r - r^* = \Phi(\bar{s} - s) + (\pi^e - \pi^{e*})$$

$$\bar{s} - s = \frac{1}{\Phi}[(r - r^*) - (\pi^e - \pi^{e*})]$$

Note that in the long run when $s = \bar{s}$, we must have $r - r^* = \pi^e - \pi^{e*}$ which results from the fact that both UIP, $\Delta s^e = r - r^*$ and relative PPP, $\Delta s^e = \pi - \pi^*$ hold in the long run. Now by solving for the long run exchange rate, we get:

$$\bar{s} = s + \frac{1}{\Phi}\left[(r - r^*) - (\pi^e - \pi^{e*})\right] \quad (2.10)$$

Following Dornbusch in assuming that the monetary model determines only the equilibrium, and not the actual exchange rate, equation (2.7) in the long run can be written as follows:

$$\bar{s} = (m - m^*) - \alpha(y - y^*) + \beta(\pi^e - \pi^{e*}) \quad (2.11)$$

In the absence of PPP, the real interest rates must diverge. Therefore, the inflation rate (differential) is reflected in the long-term interest rate (differential), but not necessarily in rates in the short term. By combining (2.10) and (2.11) we obtain:

$$s = (m - m^*) - \alpha(y - y^*) + \beta(\pi^e - \pi^{e*}) - \frac{1}{\Phi}\left[(r - r^*) - (\pi^e - \pi^{e*})\right] \quad (2.12)$$

or, alternatively, $s=(m-m^*)-\alpha(y-y^*)+[(\beta+\frac{1}{\Phi})(\pi^e-\pi^{e*})]-\frac{1}{\Phi}(r-r^*)$, then

$$s=(m-m^*)-\alpha(y-y^*)+\delta(\pi^e-\pi^{e*})-\frac{1}{\Phi}(r-r^*) \qquad (2.13)$$

where $\delta=\beta+\frac{1}{\Phi}$

Equations (2.12) and (2.13) represent the sticky-price monetary (Dornbusch-Frankel) model or the real interest differential model, which allow for a slow adjustment in domestic prices, and hence deviations from PPP. This formulation shows the role of expectations and the real interest rate, the term between brackets in (2.12), in determining exchange rate changes (Copeland, 2000). It indicates that a rise in home money supply, a decrease in domestic real income, a fall in home real interest rate or an increase in domestic expected inflation rate, *ceteris paribus*, will lead to depreciation in the home currency (Sarmas, 1996).

It is worth noting the difference in the sign of the nominal interest rate differential between the flexible-price and sticky-price equations. In order to highlight this issue, we should clear up the source of difference. Within the asset approach to the exchange rate there is a conflict regarding the relationship between the exchange rate and the interest rate. The first view might be called the Chicago theory because it assumes prices are perfectly flexible as stated by Frankel (1979) or it may be attributed to the orthodox view based on the popular Mundel-Fleming model as reported by MacDonald and Taylor (1992). As a result of the assumption of flexible prices, movements in the nominal interest rate are caused by changes in the expected inflation rate. Consequently, an increase in the domestic interest rate relative to the foreign one is due to the expected loss in the value of domestic currency resulting from inflation and depreciation. Thus, the demand for domestic currency decreases causing depreciation in the home currency (a rise in the exchange rate). This is a positive relationship between interest rate differential and exchange rate.

The second view, in which prices are assumed to be fixed in the short run, can be labelled as Keynesian theory due to its sticky-price assumption. According to this view, changes in the nominal interest rate can be attributed to the tightness of monetary policy. A fall in the home money supply relative to domestic money demand leads to a rise in the home interest rate relative to the foreign rate, with constant prices. A higher domestic interest rate than its foreign counterpart attracts capital inflows, which result in an appreciation of the domestic currency. Hence, a negative relationship between exchange rate and nominal interest rate differential exists (see Frankel, 1979).

The first perspective is a more realistic description in the case of high variation in inflation differentials as in the German hyperinflation of the 1920s. The second theory is a more realistic prescription in the case of small variations in the inflation differential as in the Canadian-US exchange rate during the 1950s (Copeland, 2000).

The Frankel-Dornbusch sticky-price model combines the two views and emphasizes the role of expectations and rapid adjustments in capital markets. This model concludes that the exchange rate differs from its equilibrium value by an amount which is proportional to the real interest rate differential; i.e. nominal interest rate differential minus the expected inflation rate differential. A high nominal interest differential as a result of tight money results in a fall in the exchange rate below its long run value. A high nominal interest differential as a consequence of a high expected inflation rate differential leads to equality between the exchange rate and its equilibrium value, which rises at the rate of the inflation differential over time.

This model, therefore, yields equation (2.13), in which the sign of the nominal interest differential is hypothesized to be negative and the sign of the expected inflation differential is hypothesized to be positive (Frankel, 1979).

From equation (2.12) it can be stated that the underlying monetary model is a special case of the Dornbusch-Frankel model. That is, when the exchange rate expectations elasticity, Φ, is infinite, the coefficient of real interest rates will be zero.

2.4 The empirical performance of the monetary approach

An enormous number of studies have been devoted to testing the empirical validity of the monetary model of exchange rates using different frequencies and spans of data and a variety of methods of estimation. Such studies have produced mixed results with comparison to a naïve random walk model. The empirical evidence on the monetary model of exchange rates generally can be divided into three periods. The first period covers studies conducted in the early regime of floating rates in the 1970s until around 1978. This work was largely supportive of the monetary model. The second episode covers studies which extended the data beyond 1978 and produced poor results for the monetary model. The third period includes recent studies which concentrate on long run cointegration properties in the monetary model. These studies have revived hope in the performance of the monetary model.

2.4.1 First period of investigation

One of the first tests of the flexible-price exchange rate monetary model was conducted by Frenkel (1976) for the German mark-US dollar exchange rate for the period 1920-1923 (a period of hyperinflation). The Frenkel's results of were supportive of the monetary model. Bilson (1978a) tested the flexible-price equation for the exchange rate of the deutschmark-pound sterling over the period January 1972 through April 1976. Incorporating dynamics and using Bayesian estimation methods, the results were in agreement with the expectations of the monetary model. Bilson (1978b) added a time trend to capture the secular decline in the demand for the pound relative to the mark for a longer period 1970-1977. He found that the results offered considerable support to the monetary approach. Hodrick (1978) also tested the flexible-price monetary model for the exchange rate of US dollar-deutschmark and of the pound sterling-US dollar using monthly data for the period July 1972 through June 1975. His findings were in line with the predictions of the model. Putnam and Woodbury (1980) used the UK pound-US dollar exchange rate over the period 1972-1974. They found that all of the coefficients were significant and correctly signed as suggested by the monetary model. Dornbusch (1979) using monthly data from March 1973 to May 1978 for the mark-dollar exchange rate and involving the long term interest rate differential, found results highly supportive of the flexible-price monetary model. Frankel (1979) assumed equalized long term real interest rates, and using a long bond interest differential as a proxy for the expected inflation term, tested his real interest differential model for the mark-US dollar exchange rate for the period July 1974- February 1978. Frankel's results were in favour of his model.

2.4.2 Second period of investigation

The early optimistic impression about the empirical validity of the monetary model of exchange rates was dramatically reversed once the sample period was extended beyond 1978. For example, Dornbusch (1980) and Haynes and Stone (1981) estimated the real interest rate differential model, and found that it yielded poor performance with respect to the signs and significance of the estimated coefficients as well as the in-sample predictive power, not to mention the weak out-of sample forecasts. The results also showed poor explanatory power and the existence of serial autocorrelation problems.

In fact the most-cited work in this period is the seminal paper by Meese and Rogoff (1983). Meese and Rogoff compared the out of sample forecasting accuracy of a set of time series and structural models of exchange rates. The competing models were used to predict at one to twelve month horizons for the exchange rates of the dollar-pound, dollar-mark, dollar-yen and trade-weighted dollar. Among the structural models was the monetary model with its two versions, the flexible and sticky price real interest differentials. Their model was specified as follows:

$$s = a_0 + a_1(m - m^*) + a_2(y - y^*) + a_3(r_s - r_s^*) + a_4(\pi^e - \pi^{e*}) + a_5\overline{TB} + a_6\overline{TB}^* + u$$

where s is the logarithm of the dollar price of foreign currency, $m - m^*$ the logarithm of the ratio of the US money supply to the foreign money supply, $y - y^*$ the logarithm of the ratio of US to foreign real income, $r_s - r_s^*$ is the short term interest rate differential, $\pi^e - \pi^{e^*}$ is the expected long run inflation differential, $\overline{TB}$ and $\overline{TB}^*$ represent the cumulated US and foreign trade balances, and u is a disturbance term. The flexible price monetary model imposes the following restriction: $a_4 = a_5 = a_6 = 0$, and the sticky price monetary model constraints $a_5 = a_6 = 0$.

The authors used the ordinary least squares, generalised least squares and Fair's (1970) instrumental variables techniques as well as specifications incorporating lagged adjustments in estimating the models. The monthly data series used in their regression started in March 1973 through June 1981. In order to generate exchange rate forecasts, Meese and Rogoff used actual realized values of their respective explanatory variables. The out of sample accuracy was measured by the mean error, mean absolute error and root mean square error. They concluded that the monetary model did not outperform the random walk model, despite the fact that their forecasts were based on realized values of the determinants. Adjusting their specifications by allowing for separate coefficients on money supplies and real incomes and adding domestic and foreign price levels as explanatory variables yielded no gain in forecasting accuracy. They suggested that structural change due to oil price shocks and changes in macroeconomic policy regimes, as well as the failure of the models to appropriately incorporate other real disturbances, may have affected the results.

These results, in effect, were built on the assumption of the presence of cointegration between the exchange rate and its fundamental determinants, resulting in inconsistent coefficients which were thus meaningless for forecasting.

The results of Meese and Rogoff's study spawned an enormous amount of research that used a variety of econometric techniques and data sets to predict exchange rate movements. Backus (1984) used US-Canadian data for the period 1971-80, and confirmed the findings of Meese and Rogoff. Wolff (1987) applied time-varying coefficients to compensate for instability in the model, and also supported Meese and Rogoff's results. Several other studies following the development of the two-step procedure provided by Engle and Granger (1987) for testing the cointegration relationships, failed to establish a cointegration between exchange rates and their fundamental determinants. For instance, Meese and Rogoff (1988) could not find strong evidence of a cointegrating relationship between exchange rates and real interest rate differentials. Moreover, other studies failing to find cointegration included Baillie and Selover (1987) and Kearney and MacDonald (1990). Meese and Rogoff's results of 1983, therefore, remained sound.

2.4.3 Potential reasons for the failure of the monetary model of exchange rate

Research on the validity of the monetary approach to exchange rates in explaining exchange rate behaviour produced mixed results and disappointing findings, particularly in the second period of investigation. This poor performance of the monetary model has been attributed to several aspects, which are mainly related to the assumptions which underpin the monetary model. Potential sources of the failure of the monetary approach can be outlined as follows:

1- The assumption of purchasing power parity (PPP). Under continuous PPP, real exchange rate, which is the exchange rate adjusted for differences in national price levels, must not diverge. However, it witnessed a high degree of swings since the advent of the floating exchange rate regime (see for example Dornbusch, 1987). Such fluctuations certainly cast doubt on the validity of PPP. Therefore, empirical work was directed towards investigating the hypothesis of PPP in both short and long run horizons. The practical research found evidence in favour of PPP in the long run, although it is unlikely to hold in the short run. Examples of such studies are Frankel and Rose (1996), Oh (1996), Wu (1996), Papell (1997) and Taylor and Sarno (1998). The absence of cointegrating relationship between exchange rates and price levels in the short run as proposed by the PPP may partly explain the failure of the flexible-price monetary model in the short run.

2- The assumption of uncovered interest parity (UIP). In international economics it is widely accepted that the hypothesis of UIP and rational expectations are at best poor and often perverse predictors of future exchange rate movements. Froot and Thaler (1990) conducted a survey of 75 published estimates and reported few cases in which the signs of the coefficient of interest rate differential in the exchange rate equation are consistent with the above hypothesis and not a single case where it exceeds the theoretical value of unity. However, almost all of these studies tested the hypotheses using relatively short maturities financial instruments. Chinn and Meredith (2005) on the other hand, used long-horizon interest rates on longer maturity bonds for the G-7 countries. Their results were much more positive for the hypotheses and were consistent with theoretical predictions.

3- The assumption of stable money demand functions in the two countries. A further explanation for the invalidity of the monetary approach can be attributed to the relative instability of the underlying money demand equations. Indeed some studies of money demand functions have shown shifts in the velocity of money (see Artis and Lewis, 1981). Meese and Rogoff (1983) stated that possible shifts in the underlying parameters of money equations may have happened as a result of the two oil shocks in the 1970s and changes in policy regimes. Such changes in structural parameters can be a possible source of the failure of the structural exchange rate models. Smith and Wickens (1986) analyzed possible reasons why the monetary model fails, and tested a random walk hypothesis for the exchange rate. They employed bilateral sterling-US dollar and German mark-US dollar exchange rates using quarterly data for the period 1973:3-1982:3. Their results showed that the breakdown of the PPP assumption and the misspecification of money demand functions were the main causes of the failure of the monetary approach. If the sources of misspecification were included, this substantially improved the explanatory power of the monetary model.

Other potential reasons for the poor performance of the monetary approach include the assumption of identical money demand coefficients in the home and foreign countries. (This assumption may be justified on the grounds of the existence of multicollinearity problems in which case the efficiency of the estimated coefficients increases. However, Haynes and Stone (1981a,b) showed that the subtractive constraints used in the equations were particularly dangerous as they may cause biased estimates and also sign reversals. Other reasons are: the assumption of the exogeneity of the money supply; the possibility of swings in expectations about future values of the exchange rate, that is, bubbles may detach the exchange rate from fundamental values in the short run; and using a single equation estimation method rather than a system estimation method to capture exchange rate dynamics.

2.4.4 Third period of investigation

After the disappointing performance of the monetary approach to exchange rates in the second period of testing, a new method of examination was adopted, that is, the long horizon and panel estimation of long run relationships. This new wave of examination relied on the investigation of the existence of cointegrating relationships between exchange rate changes and a set of explanatory variables provided by the monetary approach. For example, MacDonald and Taylor (1991) tested an unrestricted equation for the deutschmark, pound sterling and yen exchange rates against the US dollar over the period January 1976 through December 1990 using the multivariate cointegration techniques proposed by Johansen (1988) and Johansen and Juselius (1990) to examine the validity of the monetary model. They found a significant long run equilibrium relationship for these exchange rates and accepted all the coefficients restrictions implied by the monetary model.

MacDonald and Taylor (1994) used monthly data on the dollar-franc exchange rate over the period January 1976 through December 1990 to re-examine the flexible-price monetary model. Applying the multivariate cointegration technique, their findings supported the validity of the static monetary approach to exchange rates when considered as a long run equilibrium condition. However, MacDonald and Taylor rejected the monetary model as a short run explanation of exchange rate changes, since the data reject the full set of restrictions imposed by the forward-looking monetary model.

Moreover, Moosa (1994) examined an unrestricted version of the sticky-price monetary equation using monthly data covering the period 1975-86 for exchange rates of the pound sterling, the mark and the yen against the US dollar. He applied the Johansen multivariate technique of cointegration to a model allowing a distinction between traded and nontraded goods. Moosa found strong evidence in favour of the existence of a cointegrating relationship between the nominal exchange rate and a vector of explanatory fundamentals provided by the monetary approach.

Outstandingly, Mark (1995) found some evidence of the capability of the monetary model to predict exchange rate changes. Mark studied the end of quarter exchange rates of the Canadian dollar, the deutsche mark, the Swiss franc and the Japanese yen against the US dollar over the period 1973:2-1991:4. He estimated projections of 1, 4, 8, 12 and 16 quarterly changes in the logarithmic exchange rate on the deviation of the current log exchange rate from its fundamental value. The simplest monetary model that relates exchange rate changes to money supply differential and real income differential is considered to contain such fundamental value. Mark used the following function:

$$e_t = \frac{\delta}{\phi} E_t \left(\sum_{j=1}^{\infty} \delta^j f_{t+j} \right) + c$$

where $f_t \equiv (m_t - m_t^*) - \lambda(y_t - y_t^*)$; and where e is the log of domestic-currency price of one unit of foreign exchange, λ the common money demand income elasticity, ϕ the common money demand interest semi-elasticity, $\delta = \frac{\phi}{1+\phi}$, c is a constant, t refers to time and stars denote foreign quantities.

Mark (1995) presented evidence that there is an economically significant predictable component in long horizon changes in log exchange rates. This is because in three out of four exchange rates studied, he found the out of sample forecasts of the regression to outperform those of the random walk without drift at longer horizons. Although, short horizon movements tend to be dominated by noise, this noise is averaged out over time, therefore revealing systematic exchange rate changes that are determined by economic fundamentals; that is, relative money supply and relative real income. Mark's study has been criticized by Berkowitz and Giorgianni (2001) insofar as his conclusion critically depends on the assumption of the existence of a stable long run relationship between the nominal exchange rates and their fundamental determinants. Given that a number of studies (see for example, McNown and Wallace, 1989; and Sarantis, 1994) have found little evidence of cointegration among nominal exchange rates and monetary fundamentals in the post Bretton Woods float era, a relatively short span of data is usually blamed for the results of no cointegration. Using data in terms of a quarterly or monthly basis does not help, as the power of unit root and cointegration tests depend on data spans rather than frequency.

Mark and Sul (2001) examined the existence of long-run relationships between nominal exchange rates and monetary fundamentals in a quarterly panel of 19 countries for the period 1973:1 to 1997.1. Their simple formula was as follows:

$$x_{it} = f_{it} - s_{it}$$

where $f_{it} = m_{it} - m_{0t} - \lambda(y_{it} - y_{0t})$; s_{it} is the time-t log nominal exchange rate between country i=1,2,....,N and the numeraire country, labelled as 0, m, y are log nominal money stock and log real income respectively, and x is the deviation of the exchange rate from its monetary fundamental value (f).

Mark and Sul also re-examined the forecasting power of the monetary fundamentals for future exchange rate changes. The authors tried to improve on the imprecise univariate estimates and forecasts by taking advantage of the available cross-sectional data in a panel data set and assuming modest homogeneity restrictions in estimation. Using the panel dynamic OLS estimator, they found that there is a cointegrating relationship between exchange rates and the long–run determinants proposed by the economic theory. Their panel method also supports the significance of the forecasting power of the monetary fundamentals of exchange rates. Other examples of recent studies that find support to the monetary model are those which use long span data. Rapach and Wohar (2002), for instance, tested the long-run monetary exchange rate model for 14 developed countries from the late nineteenth century to the twentieth century. Rapach and Wohar used the following simple form of the monetary exchange rate model:

$$e_t = \beta_0 + \beta_1(m_t^* - m_t) + \beta_2(y_t^* - y_t)$$

where e is the nominal exchange rate measured as the number of foreign currency units per a domestic currency unit, m is the money supply, y is real income, t refers to time and asterisks denote a foreign quantity.

Their results support the simple form of the long-run monetary model for over half of their sample. However, they failed to find evidence in favour of the long run monetary model in six countries in their sample using long spans of data. Furthermore, Rapach and Wohar compared the out of sample exchange rate forecasts from a naïve random walk model with those grounded on monetary fundamentals. They found a significant forecasting power only for Belgium, Italy and Switzerland.

In addition, Francis et al. (2001) and Ahn and Oh (2001) employed the panel cointegration approach to the exchange rates of seven developed nations vis-à-vis the US dollar for the period 1973:1-1997:2 and found supportive results to the presence of cointegration from a simple monetary model.

Francis et al. (2001), Makrydakis (1998) and Miyakoshi (2000) are other examples of studies that have found evidence supportive of the validity of the monetary model of exchange rates in the long run.

Consequently, we can conclude that recent work has rekindled hope in the capability of the monetary model in explaining and forecasting exchange rate movements.

2.4.5 How well the monetary approach fits developing countries

To the best of the knowledge of the author, studies on developing economies are very scarce in the context of examining the validity of the monetary approach to exchange rates. An example of such work is the study by Odedokun (1997) which tested the monetary approach to floating exchange rates for five sub-Saharan African countries, namely, Gambia, Ghana, Nigeria, South Africa and Zaire. He used monthly data over the span 1986 through 1992. Estimating regression equations for both the flexible-price and sticky-price variants of the monetary model, Odedokun found strong support for the monetary approach. Another example is the research conducted by

Diamandis and Kouretas (1996) for four Greek drachma bilateral exchange rates for the recent experience of flexible exchange rates from April 1975-Febrauary 1994. Using the Johansen-Juselius procedure, they found that an unrestricted monetary model is a valid method for analysing the long run equilibrium relationship of the exchange rate. Moreover, Edwards (1983) analyzed the Peruvian experience with floating exchange rates during the early 1950s. He estimated a short run version of the simple flexible-price monetary model using monthly data for the period 1950-54. Despite the institutional and economic characteristics of less developed economies, such as the absence of certain markets, Edwards found that the monetary approach provides a useful benchmark for analyzing the process of exchange rate determination in these countries.

As we have seen, the monetary approach has given mixed results, especially in the short run, when it was applied to developed economies. Empirical investigations of underdeveloped economies are very rare due to, perhaps, data availability and the small number of countries that have had floating exchange rate experience. This raises the question of the capability of the monetary approach to exchange rates in explaining the behaviour of exchange rates in less developed countries. The following section gives our expectations of the performance of the monetary model if it was applied to data from developing countries.

Because developing economies are subject to more external and internal shocks, as well as policy changes, than developed economies, their structural parameters are likely to experience more movement compared to those of industrialized countries. Thus, such structural breaks, and in particular in money demand functions, are likely to reduce the capability of the monetary approach to exchange rates in less developed economies.

Given that most developing countries either have no financial markets or they are still immature, the UIP is unlikely to hold which, in turn, restricts the power of the monetary model. In addition, interest rates that are determined by the monetary authorities, on which data is available, are usually low and constant.

It may be more pertinent that the monetary approach to exchange rates was mainly formed to explain the behaviour of exchange rates that are determined by market forces, that is, demand and supply, whereas exchange rates in developing economies are fixed either to a single currency or a composite of currencies of developed economies or floated within a managed float regime[2]. Therefore, the behaviour of exchange rates in underdeveloped countries is more related to bilateral or multilateral factors, such as trade and financial linkages, than to certain sets of monetary fundamentals as assumed by the monetary model. Consequently, we anticipate the monetary approach to exchange rates to perform badly for developing countries compared to for developed countries.

2.5 Redux model

Although the traditional framework developed by Mundell (1961, 1963) and Fleming (1962) and further elaborated by Dornbusch (1976) provides an undeniable time-tested appeal of the traditional sticky prices methods, it suffers from severe limitations. In particular, the lack of explicit microfoundations introduces problems at several levels. The model provides an ad hoc specification of the price determination process and ignores the current account in exchange rate determination (Isard, 1995).

[2] Although some developing countries have actually chosen to float their currencies, this experiment has only been adopted recently. However, the monetary approach can be applied to the black market exchange rates in developing countries in which exchange rates are determined by market forces, see for example, Odedokun, M. O. (1996) Monetary Model of Black Market Exchange Rate Determination: Evidence from African Countries *Journal of Economic Studies*, **23**, 4 31-49. and Bhawnani, V. and Kadiyala, K. R. (1997) Forecasting Foreign Exchange Rates in Developing Economies *Applied Economics*, **29**, 1 51-62..

Moreover, it disregards the intertemporal budget constraints needed to give a coherent description of the current account and fiscal policy; it provides no clear description of how monetary policy affects production decisions; and it has no meaningful welfare criteria, a fact which may yield misleading policy prescriptions (Obstfeld and Rogoff, 1995). Such drawbacks stimulated researchers to develop an approach introducing nominal rigidities and market imperfection into a dynamic general equilibrium model with well-specified microfoundations. This has produced a new wave of research, called the new open economy macroeconomics and sometimes labelled the "Redux model" mainly launched by Obstfeld and Rogoff (1995) which is built on the earlier work of Svensson and Wijnbergen (1989).

Imperfect competition either in products or factor markets is an essential element in the new models. In contrast to perfect competition, where individuals are price-takers, monopoly power hypothesis assumes individuals are price setters, and therefore it explicitly allows one to analyse pricing decisions. Furthermore, setting prices above marginal cost justifies the Keynesian hypothesis of demand-determined output in the short term. Moreover, monopolistic power implies a distortion of production levels being under the social optimum, which can be corrected by monetary policy intervention (Lane, 2001).

Obstfeld and Rogoff have built a link between the rigour of the intertemporal approach (as in Sachs, 1981; Obstfeld, 1982; and Frenkel and Razin, 1996) and the descriptive plausibility of the classic contributions of Fleming (1962), Mundell (1963, 1964), and Dornbusch (1976). Their model includes the main ingredients of the intertemporal method alongside short-run nominal price stickiness and explicit microfoundations of aggregate supply. Obstfeld and Rogoff (1995) developed a sticky-price, intertemporal general equilibrium model appropriate for examining issues of open economy macroeconomics, including modelling real and nominal exchange rates. Obstfeld and Rogoff assume that there are two countries (home and foreign) which produce differentiated goods. Individuals in both countries have identical preferences and goods flow freely between the two nations. This means that the law of one price will prevail for each good and the PPP holds (Obstfeld and Rogoff, 1995). Prices are assumed to be set one period in advance, that is, prices are sticky in the short-run. For the sake of simplicity the assumption of stickiness of prices is supposed without referring to the underlying source of stickiness. First they solve the model for a steady state and then investigate the short-run effects of a monetary shock by taking a log linear approximation around this steady state. Our focus with this model will mainly be on exchange rate determination. They start with the standard consumption Euler equations, (2.14), (2.15) and money market equilibrium conditions, (2.16), (2.17), which are as follows:

$$C_{t+1} = g(1+i_t)C_t, \tag{2.14}$$

$$C_{t+1}^{*} = g(1+i_t)C_t^{*}, \tag{2.15}$$

$$\frac{M_t}{P_t} = \left[\chi^{C_t}\left(\frac{1+r_t}{r_t}\right)\right]^{1/\varepsilon}, \tag{2.16}$$

$$\frac{M_t^{*}}{P_t^{*}} = \left[\chi^{C_t^{*}}\left(\frac{1+r_t^{*}}{r_t^{*}}\right)\right]^{1/\varepsilon} \tag{2.17}$$

where C is per capita consumption index, i is the real interest rate, M is the domestic money supply, P is the price index, r is the nominal interest rate, 0<g>1, ε>0 and asterisks denote foreign variables.

The equations (2.16) and (2.17) equate the marginal rate of substitution of composite consumption for the services of real money balances to the consumption opportunity cost of holdings real balances.

Introducing consumption instead of income into the money demand function has two major reasons. Firstly, if permanent income is the relevant variable in the money demand equation, consumption is an ideal proxy because it is proportional to this unobservable variable. Secondly, since money demand is based on a transaction motivation and since all components of GNP generate transactions, consumption probably generates more money demand than the other components (Mankiw and Summers, 1986).

The consumption equations (2.14) and (2.15) take the log-linear forms:

$$\hat{C}_{t+1} = \hat{C}_t + (1-g)\hat{i}_t, \tag{2.18}$$

$$\hat{C}^*_{t+1} = \hat{C}^*_t + (1-g)\hat{i}_t, \tag{2.19}$$

near the initial steady-state path. The money demand equations (2.16) and (2.17) become

$$\hat{M}_t - \hat{P}_t = \frac{1}{\varepsilon}\hat{C}_t - \frac{g}{\varepsilon}\left(\hat{i}_t + \frac{\hat{P}_{t+1} - \hat{P}_t}{1-g}\right), \tag{2.20}$$

$$\hat{M}^*_t - \hat{P}^*_t = \frac{1}{\varepsilon}\hat{C}^*_t - \frac{g}{\varepsilon}\left(\hat{i}_t + \frac{\hat{P}^*_{t+1} - \hat{P}^*_t}{1-g}\right) \tag{2.21}$$

where ˆ denotes deviations from a symmetric steady state path, that is, the percentage change from the initial steady state value.

Considering either temporary or permanent changes from the baseline, Obstfeld and Rogoff stated that the world economy reaches its new steady state after one period. Thus, all $(t+1)$ subscripted variables in the equations (2.18)-(2.21) can be replaced with steady state changes, and all (t) subscripted variables in these equations are now interpreted as short run values.

By subtracting the foreign consumption Euler equation (2.19) from its domestic counterpart (2.18), Obstfeld and Rogoff obtained:

$$\hat{C}-\hat{C}^{*}=\hat{\bar{C}}-\hat{\bar{C}}^{*} \qquad (2.22)$$

where the hatted variables without t subscripts now represent short-run values and the hatted variables with overbars without $t+1$ subscripts represent steady state changes.

This equation, given that PPP holds, means that the real interest rate is identical home and abroad. Obstfeld and Rogoff presume that PPP holds in the short and long runs,

i.e. $\hat{s}=\hat{P}-\hat{P}^{*}$ and $\hat{\bar{s}}=\hat{\bar{P}}-\hat{\bar{P}}^{*}$ (2.23)

where s is the nominal exchange rate. Equation (2.22) states that disturbances have permanent effects on the difference between personal consumption at home and abroad.

A similar operation for money demand equations (2.21) and (2.20) gives:

$$\left(\hat{M}-\hat{M}^{*}\right)-\hat{s}=\frac{1}{\varepsilon}\left(\hat{C}-\hat{C}^{*}\right)-\frac{g}{(1-g)\varepsilon}\left(\hat{\bar{s}}-\hat{s}\right) \qquad (2.24)$$

The only vital difference between equation (2.24) and the central equation of the flexible price monetary model of exchange rates is that, unlike the flexible price model, relative money demand in Eq (2.24) depends on consumption differences rather than output differences. Here the decision to hold money involves an opportunity cost which hinges on the marginal utility of consumption.

Going back to equation (2.24) and leading it by one period, we obtain a steady state equation which is:

$$\hat{\bar{s}}=\left(\hat{\bar{M}}-\hat{\bar{M}}^{*}\right)-\frac{1}{\varepsilon}\left(\hat{\bar{C}}-\hat{\bar{C}}^{*}\right) \qquad (2.25)$$

Since the money supply shock is permanent, the short run change in relative domestic real balances must equal the long-run change, or:

$$(\hat{M} - \hat{M}^*) = (\hat{\bar{M}} - \hat{\bar{M}}^*) \tag{2.26}$$

From (2.22) and (2.24) we obtain:

$$\hat{s} = \left(\hat{M} - \hat{M}^*\right) - \frac{1}{\varepsilon}\left(\hat{C} - \hat{C}^*\right) \tag{2.27}$$

hence $\hat{s} = \hat{\bar{s}}$. This means that the exchange rate jumps immediately to its long-run level in spite of the stickiness of prices. Therefore, there is no overshooting phenomenon in the Redux model. The reason for this appears from equation (2.24). If consumption differentials and money differentials are anticipated to be constant, agents must also anticipate constant exchange rates (Obstfeld and Rogoff, 1995).

Since Redux model has been launched, a number of developments and modifications have been introduced into the original new open economy macroeconomics model, which represent different variants of it. The following sections, therefore, provide an overview of such extensions.

2.5.1 Small country with nontradables

In the appendix of their paper, Obstfeld and Rogoff (1995) extended their model by sketching a simple small open economy model with nontraded goods in which there is a possibility of exchange rate overshooting. The nontraded sector is subject to monopolistic competition and its prices are set one period in advance, whereas the traded sector is characterized by a single homogenous tradable good that sells for the same price all over the world, with perfect competition and totally flexible prices. In this appendix the equation of exchange rate response is as follows:

$$\hat{s} = \hat{P}_T = \frac{\beta + (1-\beta)\varepsilon}{\beta + (1-\beta)(1-\gamma+\gamma\varepsilon)} \hat{M} \tag{2.28}$$

where $\hat{P}_T$ is the price of traded goods which here changes proportionally to the exchange rate as a result of the dominance of the law of one price in the traded sector. A permanent money supply shock does not generate imbalance in the current account. Because tradables output is fixed, current account behaviour is determined by the time path for tradables consumption, which under log-separable preferences and a discount rate equal to the world interest rate implies that the time path of tradables consumption is perfectly flat. Therefore, the current account remains balanced. Since in this environment the monetary shock does not cause imbalance in the current account, money is neutral and the nominal exchange rate rises in proportion to the money shock in the long-run. Unlike in the two country setup, however, the exchange rate may show overshooting behaviour in the short run and the consumption elasticity of money demand determines the response of the money demand. If the elasticity $(\frac{1}{\varepsilon})$ is less than one, it will only increase by a small amount after the disturbance and vice versa. Since non-tradable goods prices are fixed, the tradable goods price, by depreciation, has to carry the burden of achieving money market equilibrium. The intuition behind the overshooting behaviour essentially parallels that of Dornbusch (1976).

2.5.2 Pricing to market behaviour

In the Redux model it is assumed that the law of one price always holds. However, recent research has pointed out that market segmentation could violate this assumption. Market segmentation means that some firms are able to charge different prices for the same good in different countries (Lane, 2001). One of the forms of market segmentation is the third degree of price discrimination which is well-known as pricing to market (PTM). Betts and Devereux (1996) formed the response of price indices to exchange rate movements as follows:

$$\hat{P} = (1\text{-n})\,(1\text{-}f)\,\hat{s}, \tag{2.29}$$

$$\hat{P}^* = \text{-n}\,(1\text{-}f)\,\hat{s} \tag{2.30}$$

where n is the home produced goods, (f) is a fraction of firms that can charge different prices in domestic and foreign markets. For any variable x let $\hat{x} = (x - \bar{x})/\bar{x}$; where $\bar{x}$ is the initial equilibrium value. Under sticky prices, as f approaches 1, P and P* are completely unaffected by exchange rate changes. Using the money market equilibrium equations and equations (2.29) and (2.30), Betts and Devereux wrote the movements of exchange rate as follows:

$$\hat{s}\,(1\text{-}f) = (\hat{M} - \hat{M}^*) - \frac{1}{\varepsilon}\,(\hat{C} - \hat{C}^*) \tag{2.31}$$

From this equation the exchange rate will depreciate (or appreciate) in response to relative national money growth (relative to national real consumption growth). The existence of the term (1-f) in the right-hand side of equation (2.31) is a consequence of the fact that the size of f determines the magnitude of the departure from PPP.

Using a linear approximation for national budget constraints combined with the equations of goods markets clearing, home and foreign non-PTM firms, and of home and foreign PTM firms, they obtained:

$$\hat{s} = (\hat{C}\text{-}\hat{C}^*)\,/\,(1\text{-}f)(\rho\text{-}1)+f \tag{2.32}$$

where $\rho > 1$

According to this equation, if f=0 (there is no PTM), a depreciation in the exchange rate leads to an increase in the foreign goods relative price which, in turn, shifts world demand towards the home goods away from foreign ones. Therefore, there will be a rise in home output, income and consumption compared to the foreign country's.

On the other hand if f=1 (full PTM), a depreciation in the exchange rate does not affect the relative prices, that is, there is no pass-through from exchange rate to relative prices. Thus, there is no expenditure-switching effect. However, as a result of the depreciation, the home currency earnings of domestic firms will increase and the foreign currency earnings of foreign firms will decrease. Hence, income will be redistributed in favour of the home country. Therefore, domestic consumption will rise for both foreign and domestic goods.

By combining (2.31) and (2.32) Betts and Devereux obtained:

$$\hat{s} = \varepsilon\,(\hat{M} - \hat{M}^{*}) \Big/ (1-f)(\varepsilon + \rho - 1) + f \qquad (2.33)$$

From (2.33) it can be said that, if $\varepsilon > 2 - \rho$, the response of the exchange rate will be to rise as f increases. Since empirical studies suggest that $\rho > 2$, the existence of PTM results in volatility in the exchange rate for any value of ε (the inverse of the consumption elasticity of money demand). Moreover, if $\varepsilon > 1$ and $f > 0$, PTM causes exchange rate overshooting as a reaction to the money shock. Thus overshooting happens when $f > 0$ and the consumption elasticity of money demand is less than unity. Since UIP must hold, nominal exchange rate overshooting occurs if monetary expansion decreases the home nominal interest rate relative to the foreign one, which is possible only if the exchange rate is anticipated to appreciate. That is, there must be a "liquidity effect" in relative short term interest rates (Betts and Devereux, 1996, 2000).

Since under PTM the expenditure-switching effect of exchange rate movements disappears, the impact of exchange rate movements is limited to consumption. Therefore, restoring monetary equilibrium requires larger exchange rate changes, which, in turn, raise the possibility of exchange rate overshooting (Lane, 2001).

In the traditional model a monetary shock poses relative price movement. Therefore, world demand will be reallocated which, in turn, reduces relative price change, which mitigates the exchange rate response. However, in the PTM model there is no change in the relative price, hence there is no demand move away from foreign goods. The real exchange rate response can be seen from the following expression:

$$\hat{P}^* + \hat{s} - \hat{P} = f\,\hat{s}.$$

If f=0 (entire pass-through), there is no volatility in real exchange rates and the law of one price holds entirely. If f = 1, PPP fails completely and the structural monetary equations play no role in determining exchange rates. Thus, the response of real and nominal exchange rates becomes identical.

The degree of impact of PTM on exchange rate volatility depends on the magnitude of three parameters: the elasticity of demand for consumer goods, ρ (positive link); the consumption elasticity of money demand, $1/\varepsilon$ (negative link); and finally, the share of goods subject to PTM, f (positive link). To show how much PTM affects exchange rate volatility, Betts and Devereux compared the contribution of PTM to the exchange rate variability with an economy where the law of one price holds continuously. They relied on the estimated values of the parameters, $\rho, \frac{1}{\varepsilon}$ and f from the previous literature. Betts and Devereux found that the variance of the exchange rate is almost three times higher than that of an economy where the law of one price entirely holds. Therefore, they concluded that the existence of PTM has a significant impact on exchange rate volatility.

2.5.3 Nominal rigidities

Hau (2000) examined the role of factor price (wages) rigidities and the presence of nontradable goods in affecting the international monetary transmission mechanism.

Since every household supplies a differentiated labour input, the product and labour markets are monopolistic. These monopolistic firms set prices as a constant markup over the factor cost, because they face a constant elasticity of demand. Thus, product prices remain unchanged as a result of the stickiness of wages in the short run. Hence, factor price rigidities provide the same international transmission effects as the product price rigidities-based Redux model does. Hau (2000) reported that an unexpected positive monetary shock raises domestic total demand, given predetermined wages. An increase in the price level is required to clear the real balances market, since the domestic demand expansion does not fully explain the money stock increase. Because product prices are sticky in the short run, import prices can take part in the price level increase as foreign exporters pass through any exchange rate depreciation to the home market. A large amount of nontradables means less impact of import prices on the home price level. Therefore, a larger home depreciation and more import price inflations are required to restore equilibrium into the home money market. Hence, nontradables create an exchange rate magnification effect for any given gap between relative money supply and relative consumption. This effect may account for the higher exchange rate volatility relative to the price level volatility.

2.5.4 Uncertainty and exchange rate

So far, all variants of the Redux model discussed rely on the assumption of certainty. While certainty equivalence allows one to derive the exact equilibrium relationships, it precludes a serious analysis of welfare changes that affect the variance of endogenous variables. Obstfeld and Rogoff (1998) again set up a model within a stochastic environment moving away from the analysis of only unanticipated shocks.

They introduced monetary uncertainty by assuming that home and foreign money disturbances follow a log-normal stochastic process. In this setting risk has an impact on asset prices, short term interest rates, and the price-setting decisions of firms, and hence on expected output and international trade flows. This new version of the Redux model has several important implications. The model allowed Obstfeld and Rogoff to compare the welfare costs of alternative exchange rate regimes and to challenge the conventional idea that small economies do better in fixing their exchange rates. Furthermore, the model contributes to attempts to explain the forward premium puzzle. However, most importantly, the model introduces exchange rate risk as a further source of exchange rate variability. Obstfeld and Rogoff (1998) assumed that domestic and foreign countries have equal trend inflation rates, which are equal to the long run nominal interest rates. Assuming valid PPP and using the traditional log-linearization, they obtained an equation for exchange rate determination. This equation essentially matches that of the Redux model except for a time-varying risk premium term. The level risk premium enters the exchange rate equation under the assumption of no bubbles. This term missing from the conventional monetary models may play a major role in explaining their failure.

These sequential developments have created a number of debates regarding, among other things, the choice of currency in which prices are sticky and whether stickiness is better assumed for goods prices or for factor prices and the effects of monetary shocks on nominal and real variables.

2.5.5 Empirical studies to the variants of Redux model

The theoretical literature on the new open economy macroeconomics is developing very rapidly; however, there has been little effort yet to test the predictions of the new models. Theorists working in this field should specify exactly which empirical exchange rate equations they would empirically estimate (Sarno, 2001).

In an attempt to gain a sense of the quantitative importance of some of the mechanisms emphasized in the theoretical models, Bergin (2003) estimated a structural general equilibrium model of a semi-small open economy using the maximum likelihood method. The fit of the model was compared to that of an unrestricted model using the likelihood ratios. The data set used in the model includes the nominal exchange rate, the current account, output, money, home and foreign price levels and the world real interest rate for three economies, namely, Australia, Canada and the United Kingdom. Bergin (2003) analysed the effects of several structural shocks of which money supply disturbance. The results of Bergin's study, in terms of the forecasting power of individual variables, were mixed. While the model predicted better for price level and output, it could not beat the forecasting of a random walk model for movements in the exchange rate or the current account for any of the three countries.

Betts and Devereux (1996) employed the VAR model to study the impact of monetary shocks on real exchange rates and included the trade balance in the system. They concluded that a calibrated PTM model explained movements in the data well, and outperformed the PPP-based Redux model which cannot generate real exchange rate movements. Hau (2002) tested his hypothesis that a monetary shock has a larger impact on the real exchange rate when the relative size of nontradables is larger. Hau found that real effective exchange rate volatility and economic openness were indeed inversely correlated in a sample of forty eight countries. That is, differences in trade openness explain a large part of the cross-country variation in the variability of real exchange rates. Other studies on the empirical side of the open economy macroeconomics models have investigated the effects of monetary disturbances on real variables, such as output and the current account, which is beyond the scope of this work.

Since the Redux model assumes the validity of PPP and links exchange rate changes to a certain set of fundamental macroeconomic variables, that is, money supply differentials and consumption differentials, just as the monetary model does, we may call them "conventional or traditional models".

So far we have discussed the theoretical foundations for the traditional fundamental exchange rate models which focus on the first moment of exchange rate behaviour, and the results of their empirical performance. Since our interest is associated with the second moment of exchange rate behaviour, the next chapter concentrates on issues of exchange rate variability, such as its importance and measures, and chapter four reformulates the conventional exchange rate models in terms of volatility.

Chapter three

Exchange Rate Volatility

3.1 Introduction

The advent of flexible exchange rate regime in 1973 caused enormous real and nominal exchange rate changes for most internationally accepted currencies. It is believed that such large exchange rate movements create uncertainty about future proceeds/costs for all dealers who participate in international transactions, such as exporters, importers and foreign investors. It is also believed that exchange rate variability has a significant impact on many micro and macroeconomic variables, such as employment, external financial obligations, the balance of payments, and the production of non-traded sector. Therefore, exchange rate volatility can be considered as an important issue for both policy makers and private economic agents.

Such effects of exchange rate fluctuations apply to all countries without exception. However, developing countries may be more concerned about exchange rate uncertainty due to specific-country characteristics. For example, most primary commodities which are the main part of these countries' exports are invoiced in terms of a major currency, often the US dollar. Forward markets in developing economies are either incomplete or do not exist, which results in a lack of hedging possibilities against exchange rate risks in such nations. In addition, economies that have a high level of openness are subject to the problem of imported inflation. It is believed that the present work is original since it is the first study in the international macroeconomic analysis that investigates the relationship between exchange rates volatility and the volatility of some macroeconomic fundamentals.

Due to the central importance of exchange rate uncertainty, this chapter discusses this issue in some detail. Thus, the following sections define the meaning of volatility, and explain the importance of variability in exchange rates. The trickiest issue in exchange rate volatility, namely, measuring variability, is then discussed followed by investigating the merits and drawbacks of a variety of exchange rate volatility proxies, before presenting the variables used in this study.

3.2 Definition of volatility

The volatility of a series in macroeconomics in general, and in financial economics in particular, has attracted an increasing deal of attention over recent decades. Before discussing volatility in more depth we should first define what we mean by volatility. Volatility is defined as the likelihood of a variable of fluctuating over time by swinging around a mean or trend or following a random walk. It is worth noting that most economic factors are volatile; however, some are much more volatile than others. For instance, prices of financial assets, and primary commodity and exchange rates tend to be much more volatile compared to manufactured goods prices and wage rates. Such volatility creates ambiguity regarding the future for economic agents. This ambiguous state is sometimes called risk and sometimes uncertainty which are proxied by different exchange rate volatility measures. However, some researchers distinguish between risk and uncertainty depending on the information economic agents have when calculating the probabilities of the occurrence of a specific event. Knight (1921) was the first to distinguish between risk and uncertainty, arguing that when the probability of the occurrence of a particular outcome can be calculated, risk exists, whereas uncertainty refers to events when it is not possible to calculate probabilities. LeRoy and Singell (1987) explained Knight's distinction by the presence of insurance markets. In the case of risk, insurance markets exist to provide cover against risk because the probabilities of the possible outcomes can be computed. On the other hand, in the case of uncertainty, each individual recognizes his own information and accordingly formulate his own subjective probabilities. Therefore, in the context of international exchange markets, a decision of risk is present when forward markets are available for hedging. A decision under uncertainty conditions exists when forward markets are not available or are incomplete (Abbott, 1999). Despite these differences the terms of exchange rate risk and uncertainty which are created by volatility, are used interchangeably here.

3.3 Importance of volatility

Some researchers argue that asset prices, including exchange rates, follow a random walk (see, for example, Koop, 2005). This makes the empirical study of behaviour of asset prices by financial economists highly difficult as a consequence of the unpredictability of change in such prices. Thus, economists try to find out whether or not the volatility of asset prices changes in a predictable way. But why are economists interested in the volatility of a variable rather than the level of the variable itself? Volatility plays an important role in a variety of areas of economics, particularly in the area of financial time series and macroeconomics. For example, early research in macroeconomics was interested in the volatility of inflation rate over time. Koop (2005) reported that some rational expectations theories of macroeconomics argue that the level of inflation itself may be not important, but its variability matters. Even if the inflation rate level is high, economic agents and decision makers can plan for the future with higher degrees of confidence if the variance of inflation is low. By contrast, in a situation where the volatility of inflation is high, it is hard for economic agents and decision makers to reliably predict what the inflation rate will be next period, and hence it is hard for them to make an appropriate plan for the future. Similarly, in the field of international economics, for the importers, exporters and foreign exchange market participants, exchange rate variability means enormous amount of losses or profits. Firms producing goods for exporting, for instance, often need to plan for future production relying on what they expect the exchange rate will be in the future. With low exchange rate volatility, planning becomes easier compared to planning under a high volatility environment. Thus, it may be more important to examine the exchange rate volatility determinants rather than the exchange rate level determinants. Consequently, we think that it is more important to examine the determinants of exchange rate volatility rather than the determinants of exchange rate level. In fact it is very important to investigate whether exchange rate volatility is indicative of overshooting (by which we mean exchange rate changes by more than the changes in the fundamental as in Dornbusch's hypothesis); or whether it is a natural response of exchange rate to the variability of the factors determining exchange rates; or if it cannot be explained by only movements in the fundamentals. Such discrimination is crucial for policy makers who wish to adopt the optimal policy. For instance, if the variability was a result of overshooting, then it may be necessary for the authorities to intervene in the foreign exchange markets and minimize the potential painful effects of the movements. If the volatility was a result of fundamentals movements, then it may be pointless for the governments to intervene in the foreign exchange markets without dealing with the root of the problem, namely the instability of the underlying variables (MacDonald, 1988).

To shed more light on the importance of studying exchange rate variability, Obstfeld and Rogoff (1998) pointed out that exchange rate volatility may be costly for welfare through two possible ways. The first and direct way is based on the assumption that people prefer a constant value of consumption to an uncertain value which fluctuates over time. To illustrate this, take for example a case of a domestic company that sets a price in terms of the local currency for products that it sells abroad. When the home currency appreciates, these products would be more expensive in terms of the foreign currency. This means lower foreign demand on the goods of this firm leading it to hire less labour. Such a reduction in employment causes lower wages, in turn, less spending on consumption. The opposite happens when the home currency depreciates. People in this situation are considered to be less happy overall, because they do not like fluctuations in their consumption and leisure.

The second and indirect channel, through which exchange rate volatility can result in welfare loss, is related to the risk resulting from exchange rate variability. If firms are risk averse, they will attempt to hedge against the risk of future exchange rate movements. These firms will put a risk premium as an extra mark-up to cover the costs of movements when setting prices for their goods. Such higher prices exert a negative effect on demand, production and, hence, consumption, taking them to levels which are less than optimal for society (Obstfeld and Rogoff, 1998, Bergin, 2004). Moreover, investors who hold a substantial proportion of their wealth in terms of foreign assets will face large valuation effects on their wealth as exchange rates change. For instance, a UK investor who holds a significant proportion of his wealth in the US will find the pound value of his wealth decreases as the pound-dollar exchange rate depreciates. These possible valuation impacts cause international investors to devote time and spend resources in an attempt to minimize such effects on their wealth (MacDonald, 1988).

Economic theory (see, for instance, Hooper and Kohlhagen, 1978; and Goture, 1985) also states that exchange rate volatility creates uncertainty as to the prices importers would have to pay, or exporters would receive, in terms of their own local currencies, at some time in the future. Relying on the assumption of risk aversion, economists assume that participants in international trade may prefer to engage in domestic activity, where profits are relatively more certain, rather than continuing trading in foreign markets where the uncovered profits earned are subject to exchange rate volatility. Alternatively, international traders may try to hedge against this uncertainty by means of using forward foreign exchange markets but at an extra cost (Hooper and Kohlhagen, 1978, Gotur, 1985). In fact a great deal of research has addressed the link between exchange rate volatility and trade flows. However, there is no clear-cut empirical evidence concerning this relationship, as a result of contradictory findings.

Moreover, exchange rate volatility affects the selection of exchange rate regime. Koop (2005) pointed out that the negative effect created by exchange rate uncertainty may partially account for the adoption of a fixed exchange rate or a common currency between groups of countries, such as the European Monetary Union. It also accounts for the emergence of forward exchange markets, which can be used to hedge against the risks resulting from currency fluctuations. The series of exchange rate volatility effects on economic variables and policies goes on; however, we just wanted to name few.

3.4 Measuring volatility

One of the most difficult tasks in investigating the volatility of a series is the specification of the appropriate measure of that volatility. Specifically in exchange rate economics a variety of exchange rate variability (uncertainty) measures have been used in the literature. However, there is no consensus between researchers about the appropriateness of one measure relative to another or the conditions under which certain measures are relevant. Thus, the choice of a volatility proxy is a difficult issue a researcher has to deal with, using his own knowledge and judgement.

Although there is no a comprehensive survey of the different measures and their merits and drawbacks in international monetary economics, McKenzie (1999) has surveyed some of the measurements of volatility which have been used in the literature which was mainly devoted to addressing the effects of exchange rate variability on trade flows.

The purpose of this section is to review the different proxies used in the literature and to discuss their relevance to the uncertainty/risk provided by exchange rate variability. The majority of empirical studies have used measures of exchange rate variability as a proxy for the uncertainty that importers and exporters and foreign exchange markets dealers face about future exchange rate movements. It should be pointed out that Akhtar and Hilton (1984) argued that measures of exchange rate volatility tend to understate the degree of exchange rate uncertainty. Their hypothesis is dependent upon the argument that changes in exchange rate contain some predictable and some unpredictable factors. Proxies of exchange rate volatility measure the actual dispersion of exchange rate changes, whereas exchange rate uncertainty depicts unpredictable future movements in exchange rates. Therefore, low levels of actual variability may be accompanied by high exchange rate uncertainty, if the timing and magnitude of exchange rate changes are very unpredictable. On the other hand, if exchange rate variability is high but the timing and magnitude of exchange rate movements are relatively predictable, measure of exchange rate volatility tend to overstate the degree of exchange rate uncertainty.

3.4.1 Measures of exchange rate volatility

In their attempts to approximate exchange rate uncertainty researchers have used various exchange rate volatility measures. Hooper and Kohlhagen (1978) used the average absolute difference between the previous forward and the current spot rate. This measure approximates exchange rate uncertainty by the average absolute forward forecast error and can be written as follows:

$$vs_t = \sum_{i=1}^{n} |f_{t-1} - s_t| \Big/ n$$

where vs is the volatility of the spot exchange rate and t refers to time, f is the forward rate, n is the sample size and i is the observation number.

The average absolute difference between the previous forward and the current spot exchange rate measures unanticipated exchange rate deviations and provides information including expectations of the spot rate set at the beginning of a sample period rather than information over the whole sample period. Nevertheless, such a measure assumes that hedging is a possible alternative for traders. It is well known that the currency of invoices, in particular for developing countries, is that of their partners or a third currency. In addition, forward exchange markets either do not exist or are not sufficiently sophisticated to provide absolute cover in most economies. For industrialized countries, short term risk may be hedged easily in forward exchange markets. Hedging against exchange risk over longer horizons, however, is much more difficult as forward contracts are usually offered for relatively short horizons. In addition, the usage of forward rate is derived from the efficient market hypothesis which should be rejected according to most of the evidence. Furthermore, over a period of time forward rate changes, this in turn will be unable to measure deviations of exchange rate from trend over that period of time (Medhora, 1990). Cote (1994) stated that there might be a strong link between the forward spread and the actual movement in exchange rate. Therefore, this measure may reflect changes in competitiveness rather than risk; hence it would be inappropriate to use in measuring exchange rate uncertainty.

Thursby and Thursby (1985) and Bailey, et al. (1986) used the absolute percentage change of the exchange rate. This measure of volatility can be written as follows:

$$vs_t = \left| \frac{(s_t - s_{t-1})}{s_{t-1}} \right|$$

where *vs* is the volatility spot exchange rate and *t* refers to time.

This proxy measures the realized exchange rate volatility and depends upon the assumption of adaptive expectations in which economic agents predict future exchange rate changes using only past information. In fact current trading behaviour at least partly affects future movements in exchange rate (Wang and Barrett, 2002). As a result of this inconsistency and inefficiency, McKenzie (1999) argued that volatility measures such as the absolute percentage change of the exchange rate is likely to generate the measurement error problem. Therefore, this proxy is useless in measuring exchange rate risk.

The relative advantages of the scale and the Gini mean difference (GMD) measures against the standard deviation were a topic of debate between Rana (1981, 1984) and Brodsky (1984). Rana (1981) Rana (1984) and Fama and Roll (1971) stated that if exchange rate movements are non-normally distributed, the sample standard deviation would be unstable and would not converge to a normal distribution as the sample size increases. Thus such a proxy, according to Rana, is an erratic and misleading measure of variability. The mathematical formulations of the GMD and the standard deviation can be written as follows:

$$GMD = \frac{\sum_{i}\sum_{j}\left|des(j) - des(i)\right|}{[(n-1)(n-2)]}$$

where *des* is the percentage change in the effective exchange rate index.

$$SD = \sqrt{n\sum(des)^2(\sum des)^2 \Big/ n(n-1)}$$

where SD is the standard deviation. Other symbols as previously defined.

Empirical evidence (see, for example Rana, 1981; and Westerfield, 1977) in fact suggests that distributions of exchange rates, generally, are fat tailed, or leptokurtic. Rana (1981) therefore advocated the use of the scale measure and the GMD to measure exchange rate variability, and applied these to data from eight developing countries. Rana divided the data into two periods, the fixed exchange rate regime period (July 1967-August 1971) and the floating rates period (March 1973-May 1977). Both measures showed an increase in exchange rate variability from the fixed period to floating rates period for nominal and real exchange rates for all countries. The standard deviation, on the other hand, indicated a reduction in exchange rate variability in real terms for three countries. The explanation given for these findings is that since, under the pegged rates period there were a large number of extreme observations in the tails of the distribution, the standard deviation would depict more volatility during this period of time (Rana, 1981). Brodsky (1984) on the other hand argued that the choice of a suitable measure of instability cannot be made depending upon a mathematical basis alone. A more practical point of view must take into account not only the question of the underlying normality of the distributions of exchange rate movements. A non-normal distribution has too many observations in the tails, thus these extreme observations should be given greater weight under considerations of risk aversion and common sense. The choice of the appropriate measure must be grounded on one's subjective value judgments concerning the nature of instability. In particular, if economic agents are risk averse, as commonly assumed, then the erratic and misleading results given by the standard deviation are entirely reasonable. Brodsky (1984) criticized the use of scale measure to calculate exchange rate variability, since it excludes the lower 28 percent and the upper 28 percent of the distribution, which may be important in characterising risk, thus making it economically meaningless. As a result of this criticism, i.e. excluding 56 percent of the data, we rule out the use of this proxy as a measure of exchange rate risk. Brodsky also reported that the GMD gives equal weight to all observations, as does the standard deviation. However, the former uses absolute differences and the latter uses squared differences, making it more relevant under the assumption of risk aversion. Moreover, the GMD pairs all observations with each other, rather than with an intraperiod mean as in the standard deviation which at least implies some intelligence in the trader (Brodsky, 1984). Since exchange rate variability measures are used to approximate uncertainty, averaging all possible pairs of differences by the GMD makes any and all exchange rate movements uncertain, whereas the standard deviation grounds all movements to a base, that is the intraperiod mean which at least implies some intelligence of the trader (Medhora, 1990). Therefore, we believe that the standard deviation is preferred to the GMD in approximating exchange rate uncertainty.

Some researchers have attempted to avoid the problem of extreme observations using a moving average of volatility measure to smooth out the data series (Chowdhury, 1993, Koray and Lastrapes, 1989, Cushman, 1988, Gotur, 1985, Klein, 1990, Asafu-Adjaye, 1999, Bouoiyour and Rey, 2005).

The moving average of the standard deviation of the exchange rate (or of the percentage change of exchange rate) can be formed as follows:

$$vs_t = \left[\left(\frac{1}{m} \right) \sum_{i=1}^{n} (z_{t+i-1} - z_{t+i-2})^2 \right]^{1/2}$$

where z is the log relative price of foreign consumer goods in terms of US consumer goods and m is the moving average order.

This approach considers lags in the uncertainty measure ranging from four to eight quarters as in Chowdhury (1993). However, the standard deviation of the exchange rate usually does not show a long pattern of autocorrelation to justify incorporating lags up to eight quarters (Hsieh, 1988). The decision concerning the length of the moving average injects an element of arbitrariness in the calculation. Also, the moving average may understate the actual costs of exchange rate changes by smoothing the data out. Amuedo-Dorantes and Pozo (2001) argued that the rolling (moving) variance of a series assumes that economic agents do not necessarily exploit patterns in the data when forecasting uncertainty. They went further and stated that this measure calculate fluctuations of exchange rate, but not necessarily the uncertainty in the exchange rate. Hence, this rolling variance proxy may be considered inappropriate as a proxy for exchange rate risk. To understand such measures more closely, consider Figure 3-1, which depicts both the standard deviation and the four quarter moving average standard deviation of the percentage changes of the Libyan dinar/US dollar exchange rate based on monthly data for each quarter for the period 1986q1-1998q4. From the plot one can clearly see that the moving average has very low extreme fluctuations and gives a significantly smoother series compared with that of the standard deviation. Since the moving average was computed for four quarters, the standard deviation fluctuations were spread out for four time periods. For instance, in the fourth quarter of 1994 there was a remarkable rise in the standard deviation which lasted for only one period[3]. The moving average, however, reached its peak in this quarter and remained about this level for the next four quarters. This means that the standard deviation takes into account devaluation effects and other movements far better than the moving average; because the actual sudden change happened once but does not last for four time points. This is better captured by the standard deviation compared to the moving average. Hence, the moving average standard deviation is excluded from our preferred proxies.

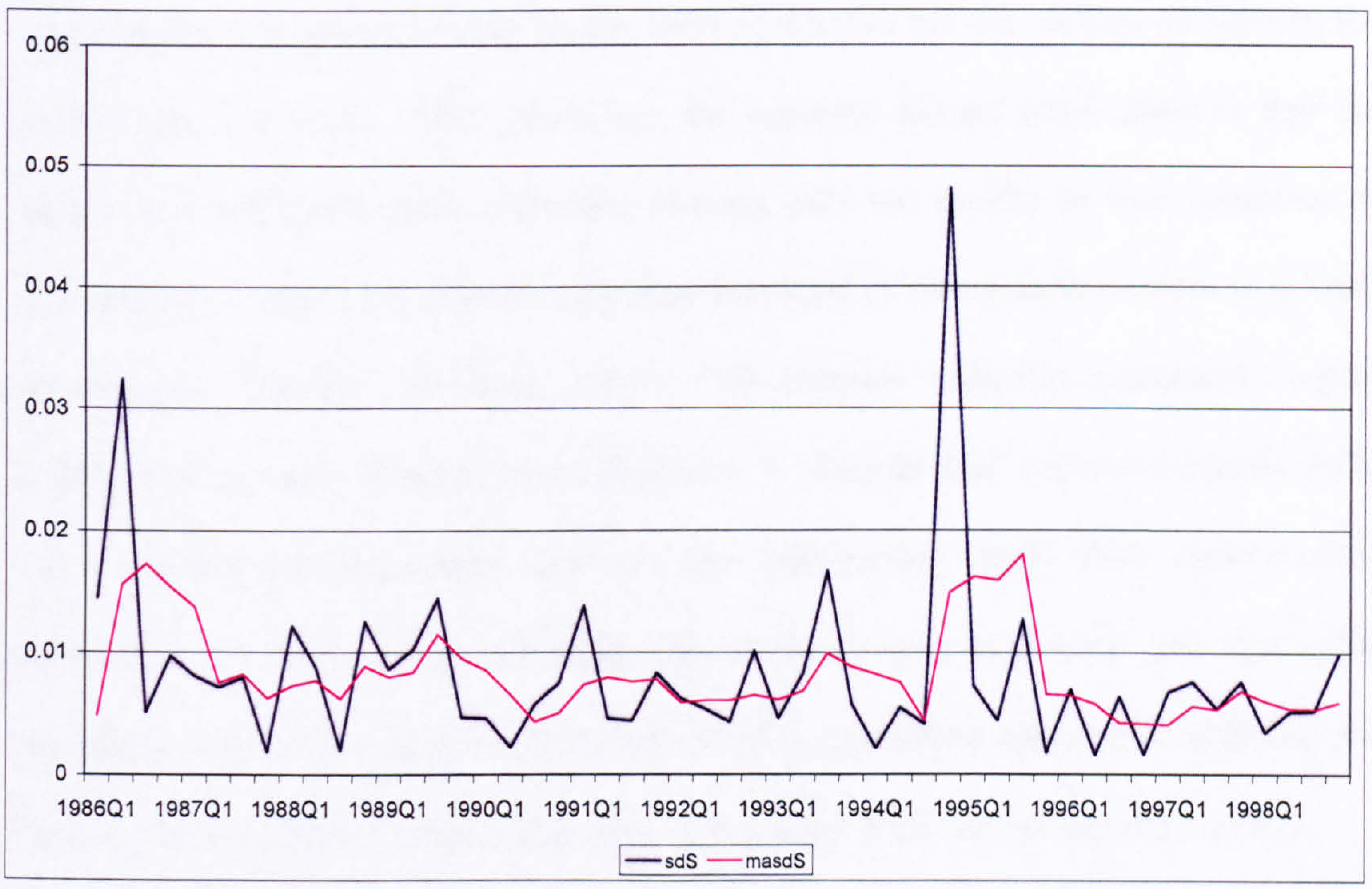


Figure 3-1: The standard deviation of the percentage changes of the Libyan dinar-US dollar exchange rate and the four quarter moving average standard deviation, 1986q1-1998q4.

[3] In the fourth quarter of 1994 the Libyan authorities devalued the dinar against the SDR from 2.24 SDR per 1 dinar to 1.9 SDR per 1 dinar which, in turn, affected the Libyan dinar/US dollar exchange rate.

Thursby and Thursby (1987) and Kenen and Rodrik (1986) used the variance of the spot exchange rate around its trend, which can be predicted from the following equation:

$$\ln s_t = \psi_0 + \psi_1 t + \psi_2 t^2 + \varepsilon_t$$

where ln stands for logarithm, ψ parameters and ε is the error term.

The use of deviations from trend, on one hand, takes into account long run misalignments of exchange rates. Traders who engage in long term transactions form expectations of future exchange rate changes on past trends, so the deviation of the spot rate movements from their underlying trend gives an indication of risk. On the other hand, using deviations from a trend of the level of the rate implies that expectations are grounded only on the level of the rate but not on past changes in the rate (Lanyi and Suss, 1982). Moreover, the variance around trend assumes that the trend is indeed predictable, therefore leaving only the misfits as true measures of uncertainty. There is no reason to attribute that kind of foresight to a trader in a small developing economy (Medhora, 1990). This criterion measures calculated ex-post rather than ex-ante forecast error, therefore it assumes that economic agents build their expectations depending only on past information rather than using current behaviour as well. Thus, although this measure can be useful for measuring misalignments in the long run, it is not useful in measuring short run variability and ex ante forecast error, which leads us to ruling it out from our preferred measures.

The standard deviation of the residuals from a first order autoregressive (AR (1)) model of the exchange rate has been used by Kenen and Rodrik (1986). In this method, under the assumption that the lagged spot rate represents the expected future spot rate, the exchange rate forecast errors are proxied by the residuals from the AR(1) model. Thus, the variability of the forecast errors acts as a proxy for exchange rate uncertainty. In similar work, Asseery and Peel (1991) used the squared residuals from the autoregressive integrated moving average (ARIMA) process fitted to the logarithm of the real exchange rate. Jansen (1989) stated that proxies of uncertainty constructed from the residuals of a model fall in an internal inconsistency. More specifically, this method estimates a model of exchange rates under the assumption of homoscedasticity and then estimates a proxy for the time-varying conditional variance from the residuals. McKenzie (1999), therefore, argued that volatility measures such as the ARIMA model residuals are likely to generate the measurement error problem, since the parameter estimates are inconsistent and inefficient. Thus, this proxy is excluded from measuring exchange rate variability in this study.

As mentioned above, the standard deviation assumes that volatility does not change over the period of time it is being calculated for. However, the volatility of a series at a specific time might be different from that at a previous or the next time. In fact, it has been increasingly recognized that asset prices and foreign exchange markets have the property of time varying variance (see, for instance, Mandelbrot, 1997). Moreover, using the variance as a measure of variability consumes the degrees of freedom which may be improper, especially in small samples. For example, one should use monthly or quarterly data to obtain volatility on an annual frequency. Accordingly, some researchers (see, for instance, Koop, 2005 and Gujarati, 2003) used another measure to act as a proxy for volatility, which can give one value of volatility using one value of the level of the variable of interest rates at each point of time, and which also shows how volatility changes over time. We will call this measure the simple measure. This measure can be illustrated as follows:

Let Y_t = exchange rate level

$Y_t^* = \log \text{ of } Y_t$

$dY_t^* = Y_t^* - Y_{t-1}^* =$ Relative change in the exchange rate

$d\overline{Y}_t^* =$ mean of dY^*

$X_t = dY_t^* - d\overline{Y}_t^*$

X_t is the mean-adjusted relative change in the exchange rate. Now we can use X_t^2 as a measure of exchange rate volatility. Being a squared quantity, a high value of volatility (X_t^2) is associated with big changes, either in a positive or a negative direction. Therefore, large increases or large falls in the exchange rate will imply that X_t^2 is positive and large. In contrast, in stable times; i.e. small or moderate movements in the exchange rate, the X_t^2 will be positive and small. To make a comparison between the standard deviation and the simple measure consider Figure 3-2, which shows the Libyan dinar-US dollar exchange rate for the period 1986-1998. The plot shows the variability of the Libyan dinar-US dollar exchange rate using the standard deviation and the simple measure. A similar pattern for the two measures can be seen, although there are some differences. For example at the beginning of the sample both proxies reveal very similar large increases in exchange rate volatility in the fourth quarter of 1994. However, in terms of scale, the standard deviation shows far larger volatility relative to that shown by the simple measure. For example the peak of the standard deviation at the end of 1994 is 0.048, whereas the peak of the simple measure at the same point of time is 0.005. This can be attributed to the different ways of measuring variability. More precisely, the standard deviation calculates the dispersion of observations of every three months from their own mean. On the other hand, the simple proxy represents the deviation of each observation in each quarter from the mean value of the whole period, as explained above. In other words, there is one crucial difference between the two measures, in that while the simple measure calculates the deviation of each observation from the sample mean value, the standard deviation calculates the deviation of each month observation of a three months period from the mean of that period. Thus, the standard deviation may capture the movement from one period to the next; whereas the simple measure captures the divergence from the equilibrium value. Consequently, one may conclude that the standard deviation is a measure of short run volatility and the simple measure captures long run fluctuations.

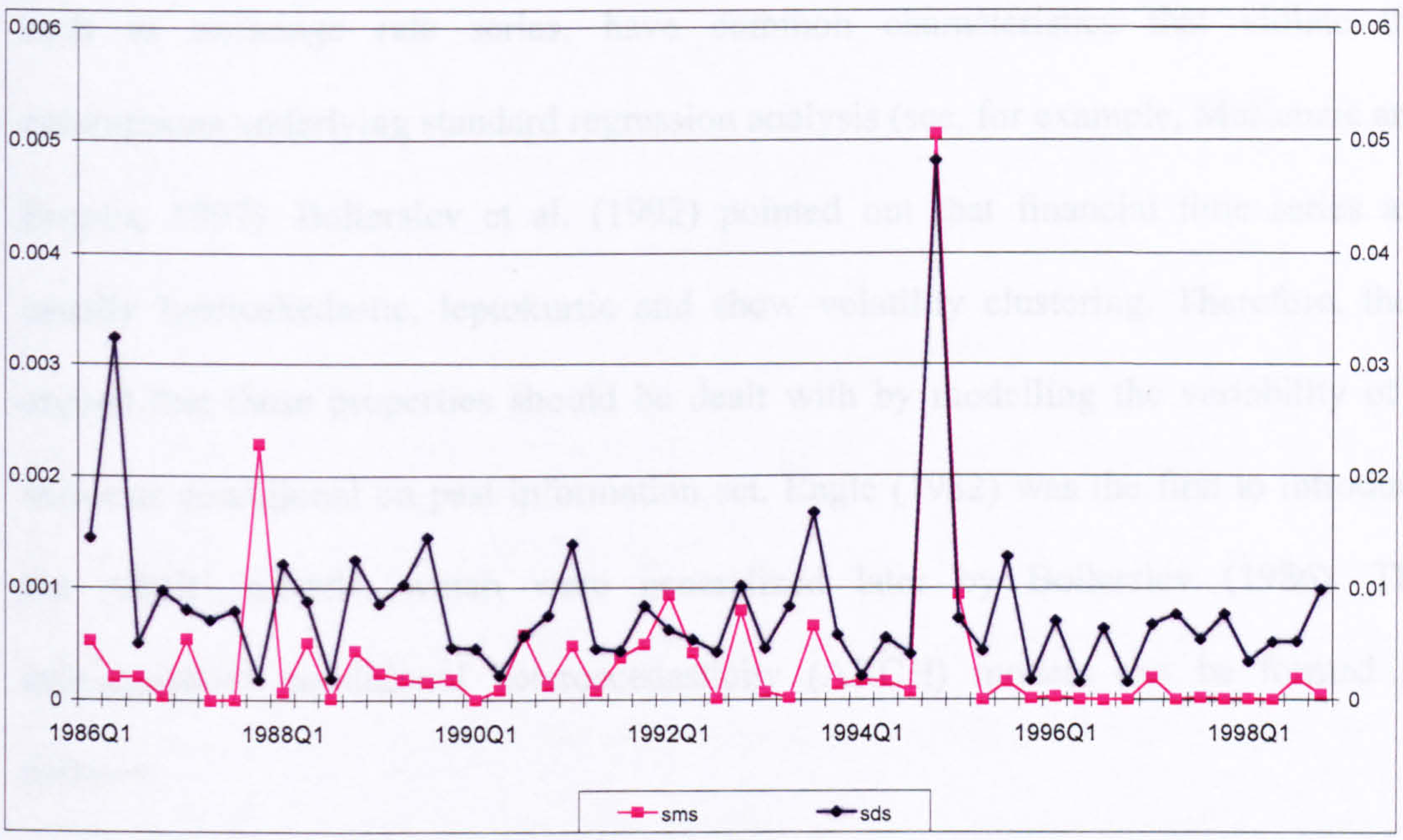


Figure 3-2: The standard deviation of the percentage changes in the Libyan dinar/ US dollar exchange rate and the simple measure of volatility, 1986q1-1998q4.

To more formally investigate the pattern of volatility, one can specify an AR(p) model for the simple measure and find out whether volatility clustering exists. Consider for example the AR(1) model:

$$X_t^2 = a_0 + a_1 X_{t-1}^2 + u_t$$

This model assumes that variability in the current period is a function of its value in the previous period plus a white noise error term. If a_1 is positive, this suggests that volatility clustering is present. In other words, if volatility was large in the previous period, it will remain large in the current period and vice versa. On the other hand, if a_1 is zero, this indicates that volatility clustering does not exist.

This model is an example of an autoregressive conditional heteroscedasticity (ARCH) model. ARCH models are increasingly popular and relatively simple tools to use, especially with financial time series data. It has been realized that asset price data, such as exchange rate series, have common characteristics that violate the assumptions underlying standard regression analysis (see, for example, McKenzie and Brooks, 1997). Bollerslev et al. (1992) pointed out that financial time series are usually heteroskedastic, leptokurtic and show volatility clustering. Therefore, they argued that these properties should be dealt with by modelling the variability of a series as conditional on past information set. Engle (1982) was the first to introduce the ARCH models, which were generalized later by Bollerslev (1986). The autoregressive conditional heteroscedasticity (ARCH) models can be formed as follows:

$$\Delta ls = x_t'\phi + \varepsilon_t \ , \ \varepsilon_t / \psi_{t-1} \sim N(0, h_t)$$

$$h_t = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + ... + \alpha_p \varepsilon_{t-p}^2$$

where x' is set of regressors, ε is an error term, ψ_{t-1} is a set of past information, h is the conditional variance, Δ is first difference operator and ls is the log of the stationary exchange rate. In order to ensure that the conditional variance is never negative, it is necessary to restrict both α_0 and α_i (where i=1, 2,…,p) to be positive. Furthermore, it is essential to assume that 0< $\sum_{i=1}^{p} \alpha_i$ <1 to ensure the stability of the autoregression process. The closer the sum of α_i to unity, the longer the persistence; i.e. any large shock in the error term in the variance equation will be associated with a persistently large variance in the ε sequence.

Bollerslev (1986) extended the ARCH model to allow for a more flexible lag structure. He added the lags of conditional variance as regressors in the conditional variance model, which produced what is known as the generalized ARCH (GARCH) model:

$$\Delta ls = x_t'\phi + \varepsilon_t \ , \ \varepsilon_t / \psi_{t-1} \sim N(0, h_t)$$

where $\varepsilon_t \sim N(0, h_t)$ and

$$h_t = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \ldots + \alpha_p \varepsilon_{t-p}^2 + \tau_1 h_{t-1} + \ldots + \tau_q h_{t-q} + u_t$$

Any high order ARCH model can be approximated by a GARCH model, and in practice it is usually the case that a GARCH(p,q) model with low values of p and q will provide a better fit to the data than an ARCH(q) model with a high value of q. Again, the sum of the coefficients α_i and τ_j must be less than unity for the model to be stationary (Harris and Sollis, 2003).

Other versions of ARCH models have been proposed in the literature. For example Engle et al. (1987) introduced the ARCH-M model in which the conditional mean of a series depends on its own conditional variance as well as that of the other regressors. The intuition behind is that risk-averse agents require compensation for holding a risky asset. Given that the riskiness of an asset can be approximated by the variance of returns, the risk premium will be an increasing function of the conditional variance of returns (Enders, 1995).

Nelson (1991) introduced the exponential GARCH (EGARCH) model, in which the natural logarithm of the conditional variance is allowed to vary over time as a function of the lagged error terms (rather than lagged squared errors). This model can capture a feature of many financial time series which is known as the asymmetry or leverage effect. In financial time series analysis, an unexpected drop tends to increase volatility more than an equivalent unexpected increase; that is, bad news increases volatility more than good news. This is true particularly for equity returns. However, in general there is no evidence that this is the case for exchange rates (Engle and Patton, 2001).

The ARCH (GARCH) model has been used as a proxy for exchange rate volatility (for example by Pozo, 1992a; Pozo, 1992b; Doroodian, 1999; Hook and Boon, 2000; and Rahmatsyah et al., 2002).

There are formal methods to test the presence of ARCH effects in a regression model that is based on time series data. The ARCH and generalized ARCH (GARCH) measure risk as a conditional variance process. They measure exchange rate variability from the squared residuals of a defined model of the exchange rate, which are assumed to be dependent on lagged squared residuals and perhaps the lagged values of the conditional variance. The merit of such approaches is that they take into account leptokurtosis in the exchange rate distribution and volatility clustering; that is periods in which large changes in a variable are likely to be followed by further large changes, and small movements are likely to be followed by further small levels of variability. In addition, these measures capture unexpected volatility compared to other measures that deal with expected volatility (Doroodian, 1999). However, the ARCH and GARCH tend to smooth out the volatility series, which may result in an understatement of the degree of exchange rate variability (Abbott, 1999).

3.4.2 The adopted measures of exchange rate volatility

Against this background, although the standard deviation of exchange rate movements as a measure for uncertainty in the international markets, mathematically may be seen as an inappropriate measure, economically it seems to be more relevant than others. Furthermore, when the standard deviation is used to calculate exchange rate variability over a floating period, the distortion it causes may be less pronounced. The simple measure appears to be similar to the standard deviation, but it may focus more on long run movements. These two volatility proxies measure the realised variability in a series. Furthermore, since financial time series such as exchange rates often exhibit the phenomena of volatility clustering and non-normal distributions the ARCH and GARCH may be appropriate measures of exchange rate volatility. This latter proxy measures the unanticipated variability in a series. Therefore, for the purpose of comparison, this study uses the standard deviation, the simple measure and the ARCH (GARCH) as proxies for volatility and their results are compared in the light of investigating the impact of fundamentals variability on exchange rate volatility.

To the best of the author's knowledge, most of the previous literature concerning exchange rate determination has been devoted to examining the determinants of exchange rate levels; i.e. they study the first moment of exchange rate changes. Our aim, in contrast, is to address the determinants of exchange rate volatility instead of level; i.e. to study the second moment of exchange rate behaviour. This is because it is believed that variability plays a crucial role in decisions made by the agents involved in exchange rate markets. We will begin the specification of volatility by using the conventional models discussed in the previous chapter. The next chapter therefore reformulates these traditional equations in terms of volatility. In addition, a model which explicitly tests exchange rate variability relating it to certain bilateral factors is presented. This model was established by Devereux and Lane (2003).

Chapter four

Modelling Exchange Rate Volatility

4.1 *Introduction*

Most of the previous literature concerning exchange rate determination has looked at the forces that drive exchange rate levels in industrialised countries which generally follow an exchange rate regime that allows currencies to freely float. In spite of its importance, exchange rate volatility has gained little or no attention from exchange rate researchers. Moreover, even within the work on the determination of exchange rate levels, underdeveloped economies have received little or no attention in terms of their exchange rate levels determination of either official or black market exchange rates. Since we strongly believe that exchange rate variability has a significant influential impact in forming the decisions of economic agents, we shall empirically investigate this matter in examining the factors that drive exchange rate fluctuations. We will rely on the most commonly used traditional exchange rate determination models in forming our volatility equations. In particular, we intend to reformulate the conventional models discussed in chapter two in order to capture exchange rate variability instead of level. In fact there has recently been some work conducted on exchange rate volatility, mainly urged by the theory of optimum currency areas (OCA). This work includes studies by Bayoumi and Eichengreen (1997a, 1997b and 1998) which was later extended by Devereux and Lane (2003). Devereux and Lane intended to examine the possible relationship between exchange rate instability and some factors derived from optimum currency areas theory and other bilateral economic criteria. This model is more about explaining the choice of exchange rate regime by a country than about the fundamentals that drive exchange rate variability. However, since the choice of an exchange rate regime can help in explaining exchange rate volatility, we will apply this model to our sample countries using annual time series observations alongside our converted traditional models. Moreover, our sample includes four developed and four less developed economies to see how these different models (traditional-based and Devereux-Lane models) perform for both sets of countries and to cover part of the gap in the previous literature resulting from ignoring less developed countries. Therefore, this chapter reformulates the conventional models, and tests for the PPP theory. The OCA theory-based model proposed by Devereux and Lane is then introduced, after which these two sets of models are compared before an augmented model is presented. Finally the data sources and variable proxies are explained.

4.2 *Reformulating the conventional models*

The exchange rate monetary models and the Redux model discussed in the second chapter were formed in terms of levels to show the effects of fundamentals' changes on the exchange rate level but not on its volatility. Since our interest is to investigate the determinants of the second moment of exchange rates, not its first moment, we need to reformulate the models introduced in the second chapter in terms of volatility. We do this by means of taking the variance of both sides of equations (2.7), (2.13) and (2.27) from chapter 2.

Firstly, the terms in differential form that appeared in equations (2.7), (2.13) and (2.27) in chapter two are replaced with tilded terms as follows:

$m - m^* = \hat{M} - \hat{M}^* = \tilde{m},$

$y - y^* = \tilde{y},$

$r - r^* = \tilde{r},$

$\pi^e - \pi^{e*} = \tilde{\pi}^e$ and

$\hat{C} - \hat{C}^* = \tilde{c}$.

Therefore, equations (2.7), (2.13) and (2.27) in chapter 2 can be rewritten respectively as below:

$$s = \tilde{m} - \alpha_1 \tilde{y} + \alpha_2 \tilde{r} \qquad \text{(Flexible-price monetary model)} \qquad (4.1)$$

$$s = \tilde{m} - \alpha \tilde{y} + \delta \tilde{\pi}^e - \omega \tilde{r} \quad \text{where } \omega = \frac{1}{\Phi} \quad \text{(Sticky-price monetary model)} \quad (4.2)$$

$$s = \tilde{m} - \gamma \tilde{c} \qquad \text{where } \gamma = \frac{1}{\varepsilon} \qquad \text{(Redux model)} \qquad (4.3)$$

Secondly, assuming that exchange rate volatility is determined by the volatility of its fundamental determinants, we may take the variance of both sides of equations (4.1)-(4.3) to put them in terms of volatility forms rather than level forms. In other words, we assume that the variance (the second moment) of exchange rates is determined by the variance of the regressors appearing in the traditional exchange rate models. Put differently, since the conventional models assume that the first moment of the fundamentals determine the first moment of exchange rates, we further assume that the second moment of these fundamentals determine the second moment of exchange rates. Accordingly, we obtain exchange rate volatility models which are based on some of the conventional exchange rate models introduced in chapter 2.

4.2.1 Exchange rate volatility model based on the flexible price exchange rate monetary model

By taking the variance of equation (4.1) we get the following function:

$$Vs = V\tilde{m} + \alpha_1^2 V\tilde{y} + \alpha_2^2 V\tilde{r} + 2\alpha_1 Cov(\tilde{m},\tilde{y}) + 2\alpha_2 Cov(\tilde{m},\tilde{r}) + 2\alpha_1\alpha_2 Cov(\tilde{y},\tilde{r}) \tag{4.4}$$

where V refers to the variance (volatility) of the series, and Cov refers to the covariance between two variables.

According to equation (4.1), a rise in the domestic money stock relative to the foreign money supply induces a depreciation of the domestic currency that is an increase in the nominal exchange rate. In other words there is a positive relationship between money supply differential and exchange rate with a coefficient equal to one. However, in equations involving variances we expect all coefficients to be positive. More precisely, in an equation such as (4.1), a change (positive or negative) in a regressor leads to a change (positive or negative) in the regressand, i.e; the relationship is either positive or negative, means that more variance (volatility) in the regressor should lead to more variance (volatility) in the regressand. Thus, all variables (in the form of variance) in the right hand side of equation (4.1) should have positive signs. These expected signs are mathematically supported; i.e. if we have two correlated random variables (X and Y), then;

$$VAR(X+Y) = VAR(X) + VAR(Y) + 2COV(X,Y),$$

$$VAR(X-Y) = VAR(X) + VAR(Y) - 2COV(X,Y) \text{ and}$$

If, however, X and Y are independent, $COV(X,Y)$ is zero (Gujarati, 2003).

This applies to all variance equations below. Therefore,

$\frac{\partial Vs}{\partial Vm} > 0$, $a_1^{\,2}>0$ and $a_2^{\,2}>0$.

Note that $\frac{\partial Vs}{\partial Vm}$ is expected to be 1 as in the differential equation.

Retaining the coefficients of the covariances would be decided in the light of economic theory.

According to money demand equations (2.2) and (2.3) in the second chapter, money demand depends on real income and the interest rate. Therefore, the covariances between $\tilde{m}$ and $\tilde{y}$ on the one hand, and between $\tilde{m}$ and $\tilde{r}$ on the other hand, are expected to be nonzero. Nevertheless, such relationships are anticipated to hold in terms of levels but not in the form of differentials between domestic and foreign quantities. Therefore, we will investigate whether such relationships in terms of differentials exist. If significant relationships are found between $\tilde{m}$ and $\tilde{y}$ on the one hand and between $\tilde{m}$ and $\tilde{r}$ on the other hand, their covariances will be kept in equation (4.4). However the effects of these variables (the covariances) on exchange rate volatility are unknown and will be left to empirical tests.

On the other hand, that $\tilde{y}$ and $\tilde{r}$ are not expected to be related to each other, and thus we expect $Cov(\tilde{y},\tilde{r}) = 0$. Accordingly, equation (4.4) can be rewritten as follows:

$$Vs = V\tilde{m} + \alpha_1^2 V\tilde{y} + \alpha_2^2 V\tilde{r} + 2\alpha_1 Cov(\tilde{m},\tilde{y}) + 2\alpha_2 Cov(\tilde{m},\tilde{r}) \qquad (4.5)$$

4.2.2 Exchange rate volatility model based on the sticky price real interest exchange rate monetary model

An operation to equation (4.2) similar to that conducted for equation (4.1) yields:

$$\begin{aligned} Vs = {} & V\tilde{m} + \alpha^2 V\tilde{y} + \delta^2 V\tilde{\pi}^e + \beta^2 V\tilde{r} + 2\alpha Cov(\tilde{m},\tilde{y}) + 2\delta Cov(\tilde{m},\tilde{\pi}^e) \\ & + 2\beta Cov(\tilde{m},\tilde{r}) + 2\alpha\delta Cov(\tilde{y},\tilde{\pi}^e) + 2\alpha\beta Cov(\tilde{y},\tilde{r}) + 2\delta\beta Cov(\tilde{\pi}^e,\tilde{r}) \end{aligned} \qquad (4.6)$$

Since the inflation rate represents the cost of holding money, it is anticipated that a negative relationship between money supply and inflation rate will exist. Thus, the relationship between $\tilde{m}$ and $\tilde{\pi}^e$ will be explored, and the covariance will be included in equation (4.6) if it is significant.

In the economic literature there has been a long debate about the relationship between interest rates and inflation in the context of the so-called Fisher Effect. Fisher (1930) claimed a one-to-one relationship should exist between inflation and interest rates in a world of perfect foresight, with real interest rates being entirely determined by real factors and not related to the expected inflation rate (see, for instance, Cooray, 2003; and Atkins and Coe, 2002). There have been mixed results in empirical tests of the Fisher Effect. Thus, we cannot confirm or deny such relationship; hence, this alleged relationship will be examined. However, we do not anticipate the presence of a relationship between inflation rate differential and real income differential, namely $Cov(\tilde{y},\tilde{\pi}^e)=0$. The rest of coefficients' expected signs are as mentioned above in equation (4.4). Thus, equation (4.6) can be rewritten as follows:

$$\begin{aligned}Vs = V\tilde{m} + \alpha^2 V\tilde{y} + \delta^2 V\tilde{\pi}^e + \beta^2 V\tilde{r} + 2\alpha Cov(\tilde{m},\tilde{y}) + 2\delta Cov(\tilde{m},\tilde{\pi}^e)\\ + 2\beta Cov(\tilde{m},\tilde{r}) + 2\delta\beta Cov(\tilde{\pi}^e,\tilde{r})\end{aligned} \tag{4.7}$$

4.2.3 Exchange rate volatility model based on Redux exchange rate model

Similarly, taking the variances of both sides of equation (4.3) will produce the following function:

$$Vs = V\tilde{m} + \gamma^2 V\tilde{c} + 2\gamma Cov(\tilde{m},\tilde{c}) \tag{4.8}$$

The level of consumption can replace real income in the money demand function (see, for example, Mankiw and Summers, 1986) which means that a link between money supply and consumption may be found. If so the covariance between $\tilde{m}$ and $\tilde{c}$ will be kept in equation (4.8).

Therefore, we now have three equations which specify the relationships between exchange rate volatility (the dependent variable) and the volatilities of some suggested fundamental differentials (the independent variables).

Since our volatility models are derived from the conventional fundamentals-based exchange rate models, which in turn substantially depend on the hypothesis of PPP as discussed in chapter two, it is essential to investigate the validity of PPP in our sample. The following section deals with this issue.

4.3 *The validity of the PPP hypothesis*

Since most structural exchange rate models critically depend on the assumption of PPP, it is important to deal with this issue in more detail.

The purchasing power parity (PPP) hypothesis is a theory of exchange rate and price determination. It states that the exchange rate, defined as the number of domestic currency units required to purchase one unit of the foreign currency, should be equal to the price ratio of the domestic to the foreign country. The key notion underlying PPP is that deviations from PPP imply profitable commodity arbitrage opportunities, which, if exploited, will tend to drive the exchange rate towards PPP. This is well-known as the strict or absolute version of PPP, which relies on the law of one price. The law of one price states that the price of a single product when converted to a common currency will be the same all over the world. However, the equilibrium price of a given good usually does not fulfil this law. Transport costs, trade barriers, technology differences, imperfect competition and so forth could lead to divergences from the law of one price. Therefore, the absolute version of PPP can hardly be expected to hold. An alternative, less restrictive version of PPP, is the so-called relative PPP which relates the relative change in exchange rate to the percentage change in relative price (that is the inflation rate differential) between two countries. While the former focuses on the relationship at a particular point in time, the latter concentrates on the relationship between two points in time.

Investigating the validity of PPP theory is an essential issue, because it represents a key assumption in international monetary economics. For example, the flexible price monetary model assumes that PPP continuously holds. The sticky-price models allow for exchange rate deviations from PPP in the short run, but they retain it as a long run equilibrium condition. Because of its importance as a long run determinant of exchange rates, PPP has been subject to a growing number of empirical tests for the G-7 countries. The practical work on the validity of PPP has produced mixed findings. For example, studies by Corbae and Ouliaris (1988) and Enders (1988) failed to find evidence of the validity of PPP, while it was supported in studies by Kim (1990) and Fisher and Park (1991). For developing countries, Bahmani-Oskooee and Mirzai (2000), for example, applied the Kwiatkowski et al. (1992) (KPSS) test to real effective exchange rate data, and stated that PPP holds in the majority of the developing countries considered.

If PPP is proven to be valid, the traditional exchange rate models will avoid a substantial criticism, as the PPP is one of their central assumptions. Therefore, we test the validity of the PPP theory for our sample. The sample under consideration here consists of four developed countries (Canada, Japan, Germany and the UK) and four developing oil-exporting countries (Algeria, Kuwait, Libya and Venezuela) and the numeraire country is the US. The developed country group represent the largest economies in the world after the US. The underdeveloped country group contains two small and two relatively big oil-producing countries. In order to test the validity of

PPP for our sample we use three methods. More precisely, we test the relative version of PPP theory by plotting the long run behaviour of percentage change in exchange rates and inflation differentials of our sample countries. Moreover, the PPP hypothesis implies that the real exchange rate (defined as nominal exchange rate multiplied by the foreign to domestic price ratio) is constant over time. Thus, we also plot the real exchange rates to see whether they are stationary over time. The price index used for computing the real exchange rates is the wholesale price index for all countries, except for Algeria and Libya where the consumer price index is used as a result of data availability. Finally, ADF unit root tests are conducted for real exchange rates to find out whether they are stationary as a further examination of the graphical test. However, before using the results of these tests are shown, an overview is given of the economic policies and events for our sample economies which may have affected the paths of their exchange rates.

4.3.1 Macroeconomic policies and data of the countries considered

This section reviews the macroeconomic and exchange rate policies conducted in our sample countries and the important economic events which took place during the period under consideration, 1973-1998.

The developed economies have generally floated their exchange rates after the collapse of the Bretton Woods's system. Moreover, we can say that the two oil price shocks have caused a major impact on macroeconomic policies in most countries, including our sample countries.

Canada

Until 1982 the Canadian monetary authorities kept the money supply (M1) as a policy target. Since that date the authorities have monitored a number of nominal indicators such as nominal demand, monetary aggregates and prices. All of this information is assessed by the Bank of Canada with respect to its implications for the achievement of its medium-term objective of stable prices. In the late 1980s the Bank of Canada has placed a greater emphasis on the broader monetary aggregates, M2, M3 etc. and since 1991 the Canadian central bank has clearly announced that inflation is the main target of its monetary policy.

The proxies for the variables used in the applied work later in the thesis for this country are as follows: exchange rate is approximated by nominal exchange rate defined as the number of domestic currency units per one US dollar; money supply is approximated by seasonally adjusted M1; income is approximated by the index of industrial production; interest rate is approximated by the Treasury bill rate; inflation rate is approximated by the first difference of the log consumer price index; and consumption is approximated by per capita real consumption.

Germany

The German public is well known as being inflation averse; therefore, the main aim of the monetary authorities has targeted inflation. After the first oil price shock there was a tightening of monetary policy. In 1979 the European exchange rate mechanism (ERM) was set up and the German mark played a central role therein. During the 1980s the central bank of Germany tightened monetary policy gradually to maintain competitiveness and intervened in the foreign exchange market to prevent further depreciation. The reunification of Germany in 1990 caused significant inflationary pressure for which the central bank of Germany (the Bundesbank) responded by increasing interest rates (contractionary monetary policy). Generally, when a conflict emerges between the internal and external stability, the choice of the Bundesbank has been to limit inflation.

The proxies for the variables used in the applied work later in the thesis for this country are as follows: exchange rate is approximated by nominal exchange rate defined as the number of domestic currency units per one US dollar; money supply is approximated by seasonally adjusted M1; income is approximated by the index of industrial production; interest rate is approximated by the Treasury bill rate; inflation rate is approximated by the first difference of log consumer price index; and consumption is approximated by per capita real consumption.

Japan

The main goal for the Japanese monetary authority after the collapse of Bretton Woods was the stability of the exchange rate for the sake of retaining the competitiveness of the Japanese economy. Thus, the exchange rate regime in Japan can be described as a managed flexible exchange rate. Following the oil price shocks the Bank of Japan tightened the monetary policy and then eased it by reducing interest rates in the 1980s, which led to an explosion in financial and estate asset prices (the bubble years). In 1993 and after the interest rates reached record lows, and the Japanese authorities conducted a stimulus fiscal policy. The Bank of Japan progressively eased short-term interest rates during 1991-1995.

The proxies for the variables used in the applied work later in the thesis for this country are as follows: exchange rate is approximated by nominal exchange rate defined as the number of domestic currency units per one US dollar; MONEY supply is approximated by seasonally adjusted M1; income is approximated by the index of industrial production; interest rate is approximated by the call money rate; inflation rate is approximated by the first difference of log consumer price index; and consumption is approximated by per capita real consumption.

United Kingdom

After 1976 the British government introduced the policy of monetary targets and reinforced this in the early 1980s. The authorities conducted a tightening fiscal and monetary policy in the late 1980s and early 1990s. However, these measures did not prevent the departure of the British sterling from the European exchange rate mechanism (ERM) in 1992. Shortly after the abandonment of the ERM, the British authorities in 1992 set up an inflation targeting policy which lasted until 1997 when the Bank of England gained its independence in conducting monetary policies.

The proxies for the variables used in the applied work later in the thesis for this country are as follows: exchange rate is approximated by nominal exchange rate defined as the number of domestic currency units per one US dollar; money supply is approximated by the narrow definition of money, M0, because there is no data available about M1. Moreover, data on M2, money plus quasi money, does not exist for the whole period on a monthly basis for this country; income is approximated by the index of industrial production; interest rate is approximated by the Treasury bill rate; inflation rate is approximated by the first difference of log consumer price index; and consumption is approximated by per capita real consumption.

For the developing country sample there has unfortunately been neither clear policy targets no transparency, accountability nor credibility. This may be as a result of the nature of their political regimes. However, there are definitely some events which may have exerted some sort of impact on their exchange rates. Since these countries are oil-exporting, we believe that the two oil price shocks have had a big influence on these economies.

Algeria

The central bank of Algeria kept the Algerian dinar (AD) linked to a composite of currencies until October 1994 and then deserted this regime in November 1994 when the AD was floated in a managed fashion. Nonetheless, the official exchange rate of the AD is still fixed to a basket of currencies (IMF: IFS). In the late 1980s and early 1990s the country witnessed high level of political instability which resulted in armed domestic conflict.

The proxies for the variables used in the applied work later in the thesis for this country are as follows: exchange rate is approximated by nominal exchange rate defined as the number of domestic currency units per one US dollar; money supply is approximated by seasonally adjusted M1; income is approximated by the index of crude petroleum production; interest rate is approximated by the discount rate; inflation rate is approximated by the first difference of log consumer price index; and consumption is approximated by per capita real consumption.

Kuwait

For the sake of enhancing the relative stability of the Kuwaiti dinar (KD) exchange rate against other currencies, especially the US dollar, and protecting the domestic economy from imported inflation, the Kuwaiti authorities have pegged the KD to a weighted basket of currencies since 18th March 1975. This composite basket consists of the currencies of the countries which have significant trade and financial relationships with Kuwait. The dollar has the lion's share in the basket, which is regularly adjustable (Central Bank of Kuwait, 2003). The country suffered from Iraqi invasion in August 1990, which lasted for six months.

The proxies for the variables used in the applied work later in the thesis for this country are as follows: exchange rate is approximated by nominal exchange rate defined as the number of domestic currency units per one US dollar; money supply is approximated by seasonally adjusted M1; income is approximated by the index of crude petroleum production; interest rate is approximated by the discount rate; inflation rate is approximated by the first difference of log consumer price index; and consumption is approximated by per capita real consumption.

Libya

Since 1986 the Libyan dinar has been fixed to the SDR basket (IMF: IFS). The Central Bank in Libya kept the interest rate fixed until 1993 when it started to reduce it gradually. As this bank is not independent, it funded the deficit in the budget by issuing money during the 1970s and 1980s. The country suffered from sanctions imposed by the UN from 1993 to 1999 when they were suspended and entirely lifted in 2003.

The proxies for the variables used in the applied work later in the thesis for this country are as follows: exchange rate is approximated by nominal exchange rate defined as the number of domestic currency units per one US dollar; money supply is approximated by seasonally adjusted M1; income is approximated by the index of crude petroleum production; interest rate is approximated by the discount rate;

inflation rate is approximated by the first difference of log consumer price index; and consumption is approximated by per capita real consumption.

Venezuela

The Venezuelan Bolivar (VB) was traditionally pegged to the US dollar until 1983 when the Bolivar was devalued. During the 1980s the exchange rate regime was characterized by a dirty float where the central bank sets a daily nominal exchange rate according to a number of factors including inflationary differentials (Edwards, 1995). The VB was independently floated in March 1989. In July 1994 the VB was fixed to the US dollar until April 1996 when a managed floating system was again chosen (IMF: IFS).

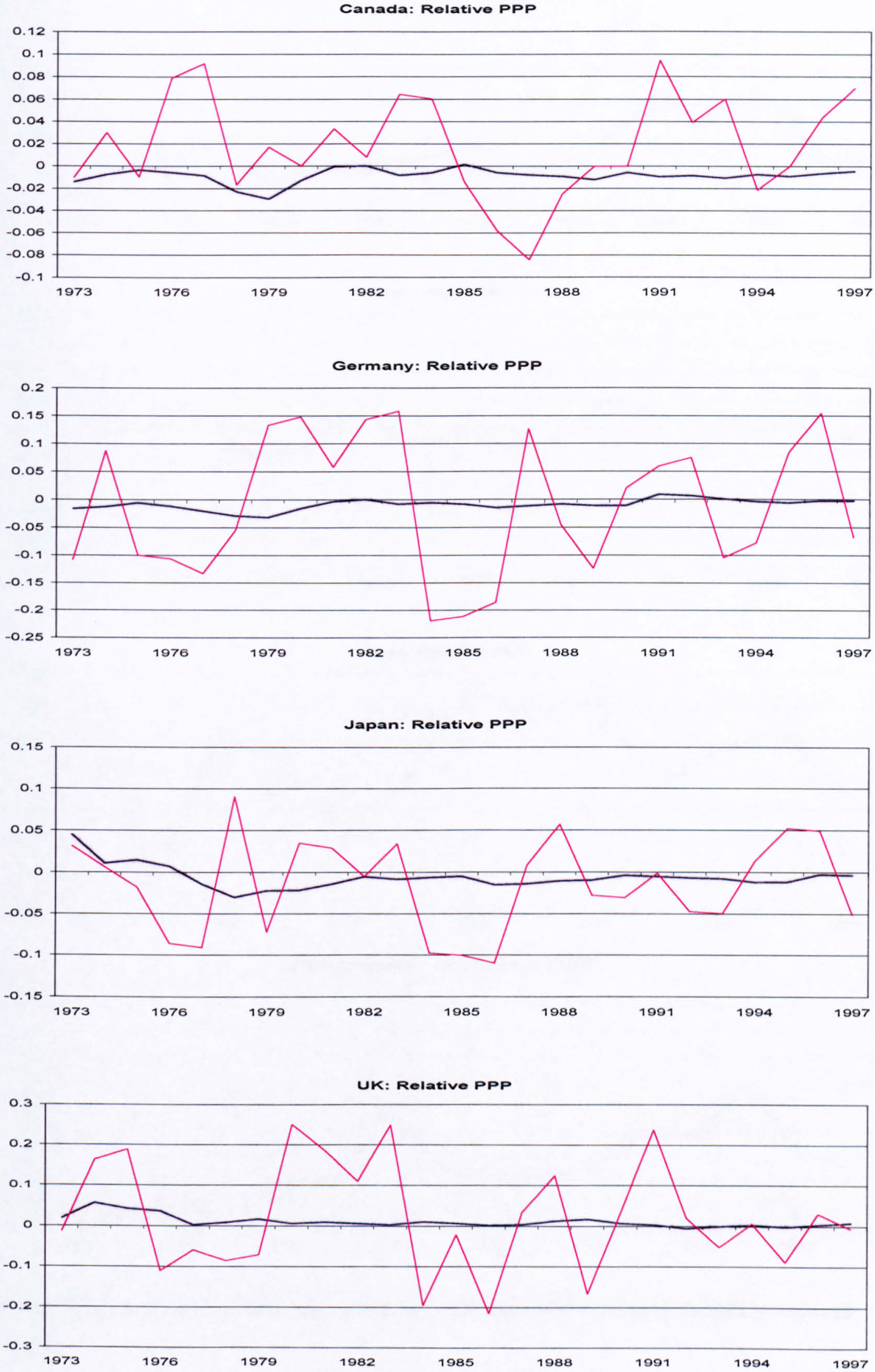
The proxies for the variables used in the applied work later in the thesis for this country are as follows: exchange rate is approximated by nominal exchange rate defined as the number of domestic currency units per one US dollar; money supply is approximated by seasonally adjusted M1; income is approximated by the index of crude petroleum production; interest rate is approximated by the discount rate; inflation rate is approximated by the first difference of log consumer price index; and consumption is approximated by per capita real consumption.

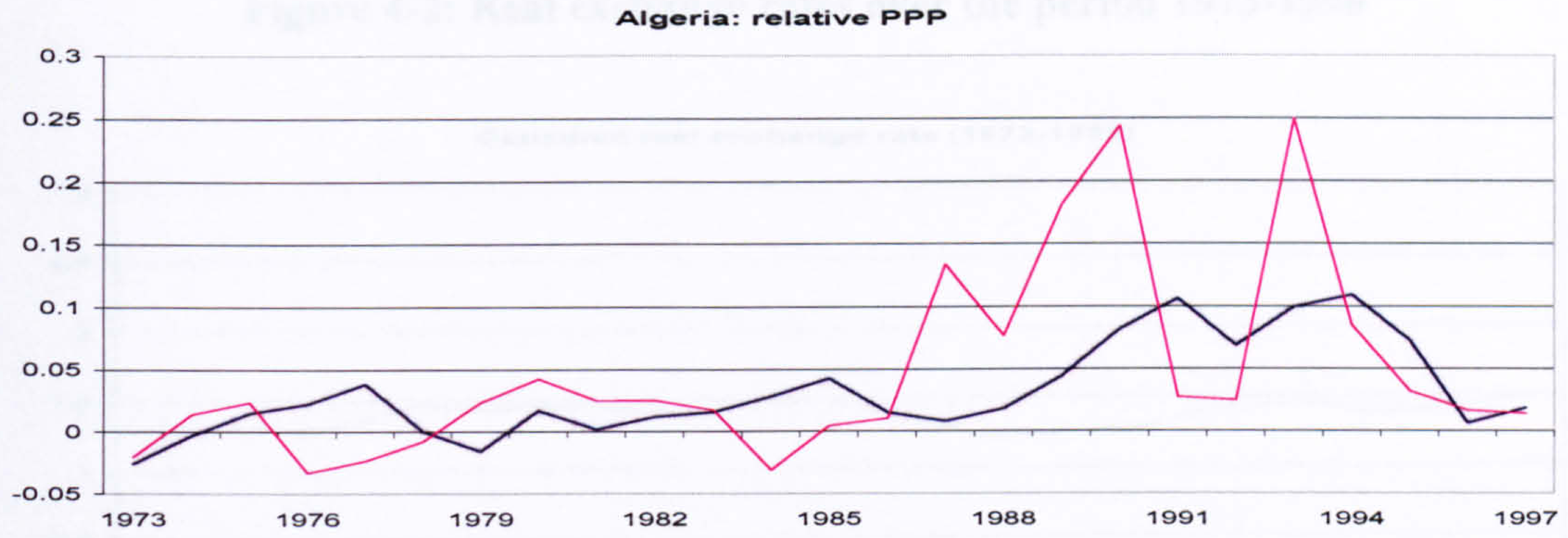
4.3.2 Examining PPP theory for the sample considered

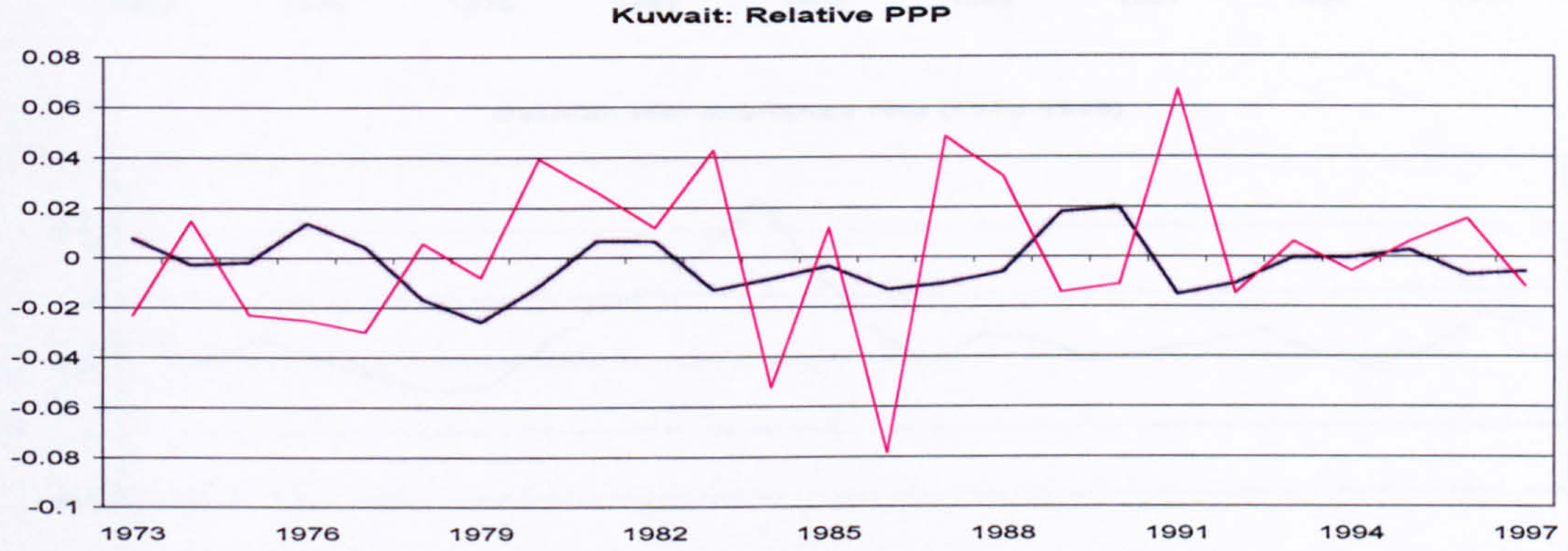
In order to investigate the validity of relative PPP, we plotted the long run behaviour of percentage change in exchange rates and inflation differentials of our sample countries. Figure 4-1 shows generally supportive results for the hypothesis of PPP. The plots in Figure 4-1 reveal a clear-cut short run invalidity of the PPP hypothesis, as exchange rate behaviour is far more volatile than relative inflation rates in all cases (except the Libyan case). However, these percentage changes in exchange rates eventually converge to some long run values. This long run value of the currency reflects inflation in the home country relative to that in the foreign country. This argues against the exchange rate models that assume that PPP continuously holds, such as the original Redux model. However, in the long run, relative PPP seems to hold as the exchange rate trend swings around the inflation differential lines. Therefore, we can state that the key assumption of most traditional exchange rate models in international economics is valid in the long run in our sample set.

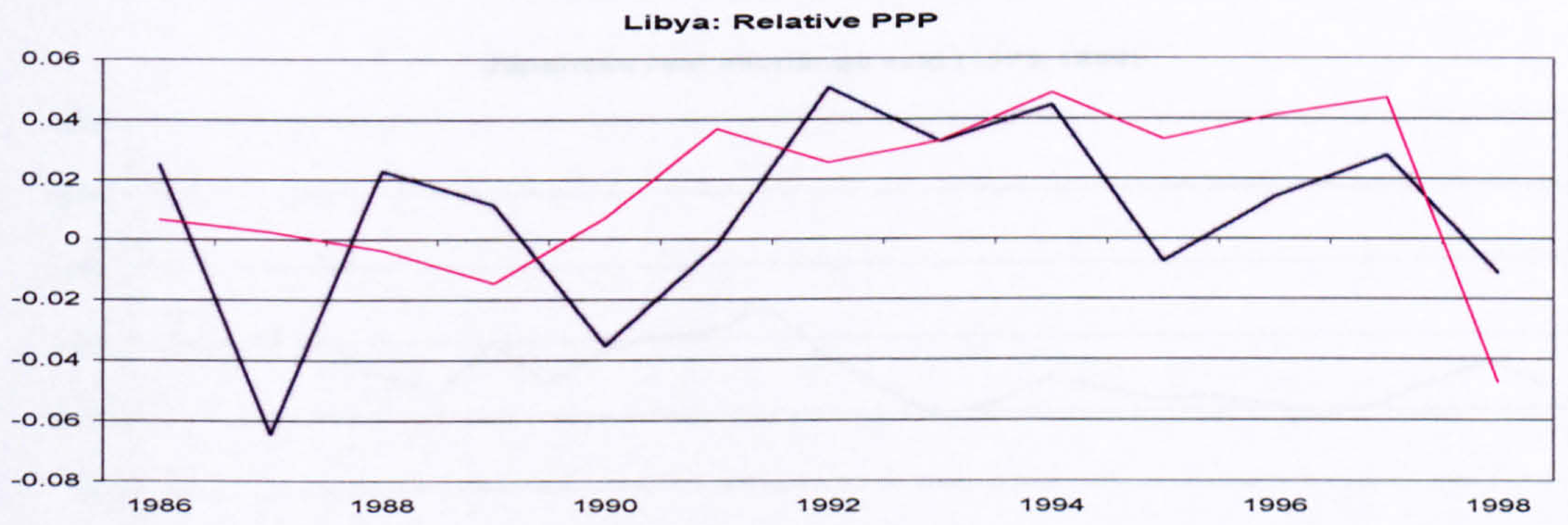
The second test for the validity of PPP in the long run is represented in the graphical plots of real exchange rates using annual data. Figure 4-2 clearly shows that during the 1980s there was a departure from the long run means, especially in Germany, Japan and the UK, which then become mean reverting. In Canada, Kuwait and Libya we can see quite stable real exchange rates. The Algerian case seems to be odd relative to the other countries in that from 1987 there was a clear increasing trend in the real exchange rate (this may be as a result of using the consumer price index instead of the wholesale price index in computing the real exchange rate or as a result of political instability). For the Venezuelan case notice that during the 1980s there was some instability compared to the 1990s. Accordingly, we may state that, except for Algeria, generally there are relatively stable and mean reverting real exchange rates. This supports the results reached above using relative PPP.

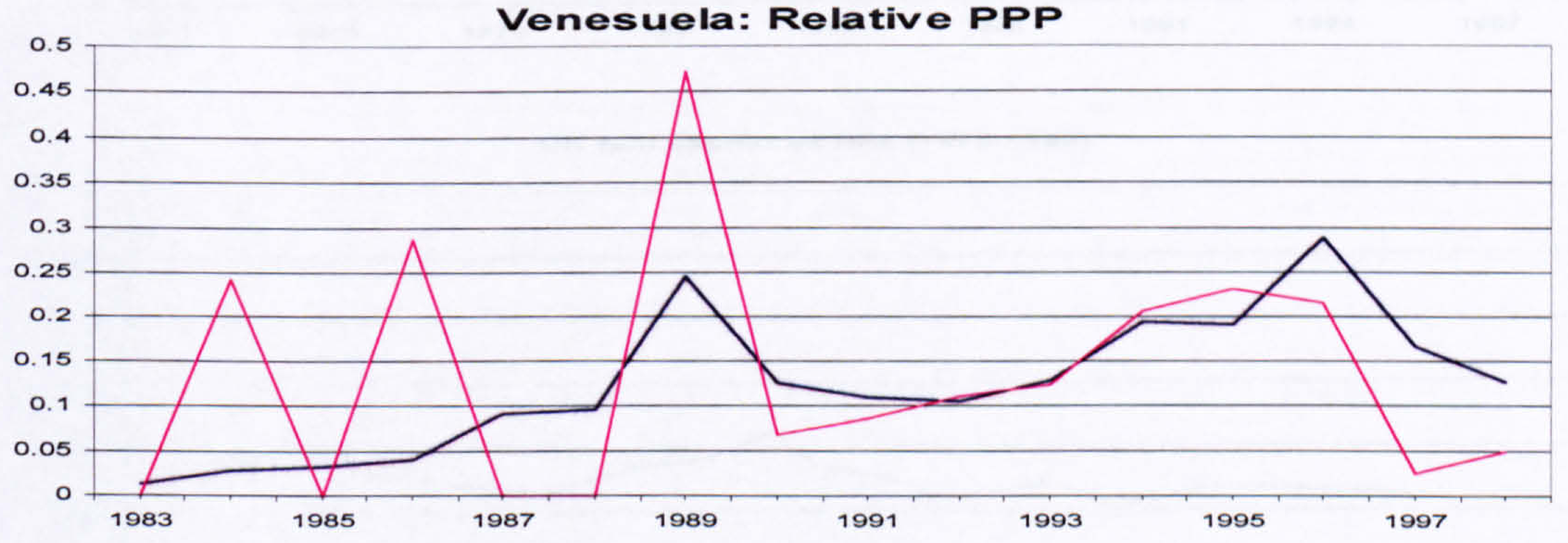
Figure 4-1: Exchange rates percentage change and inflation rate differentials, 1973-1998





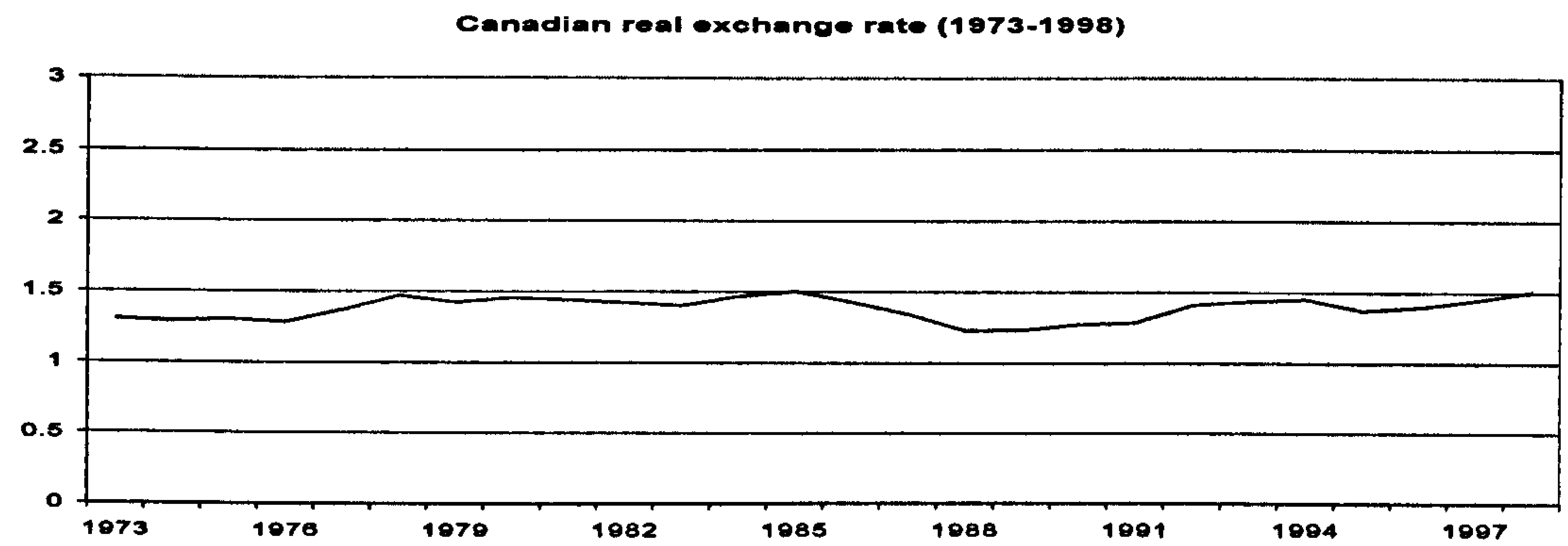


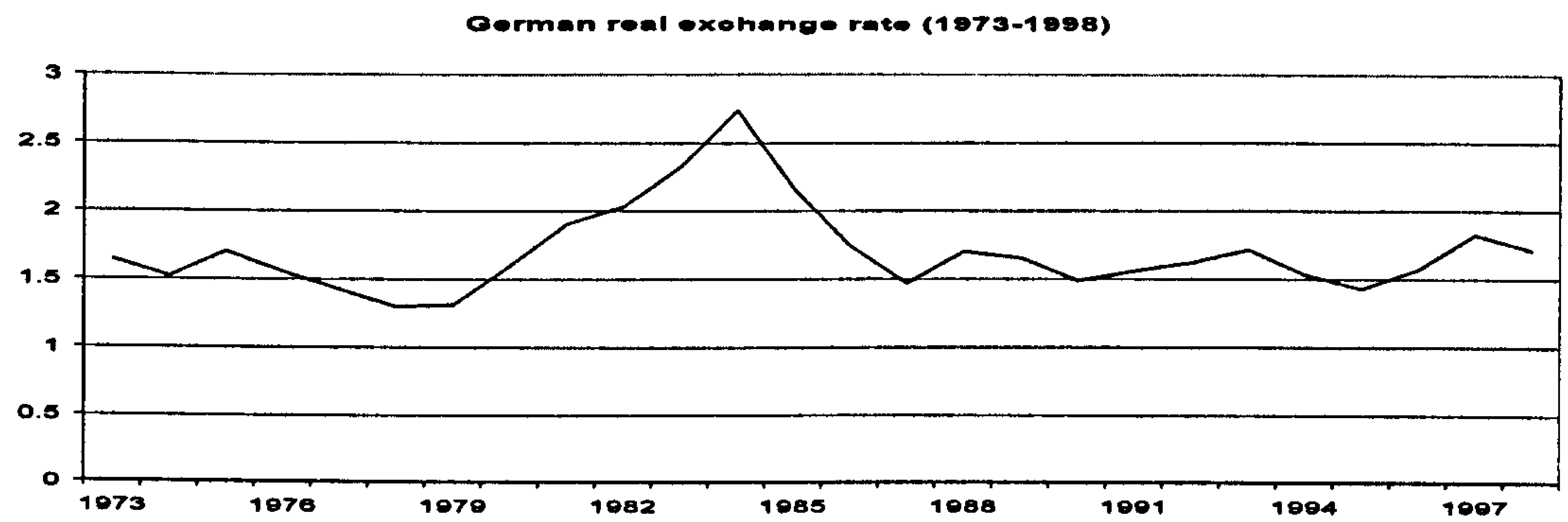


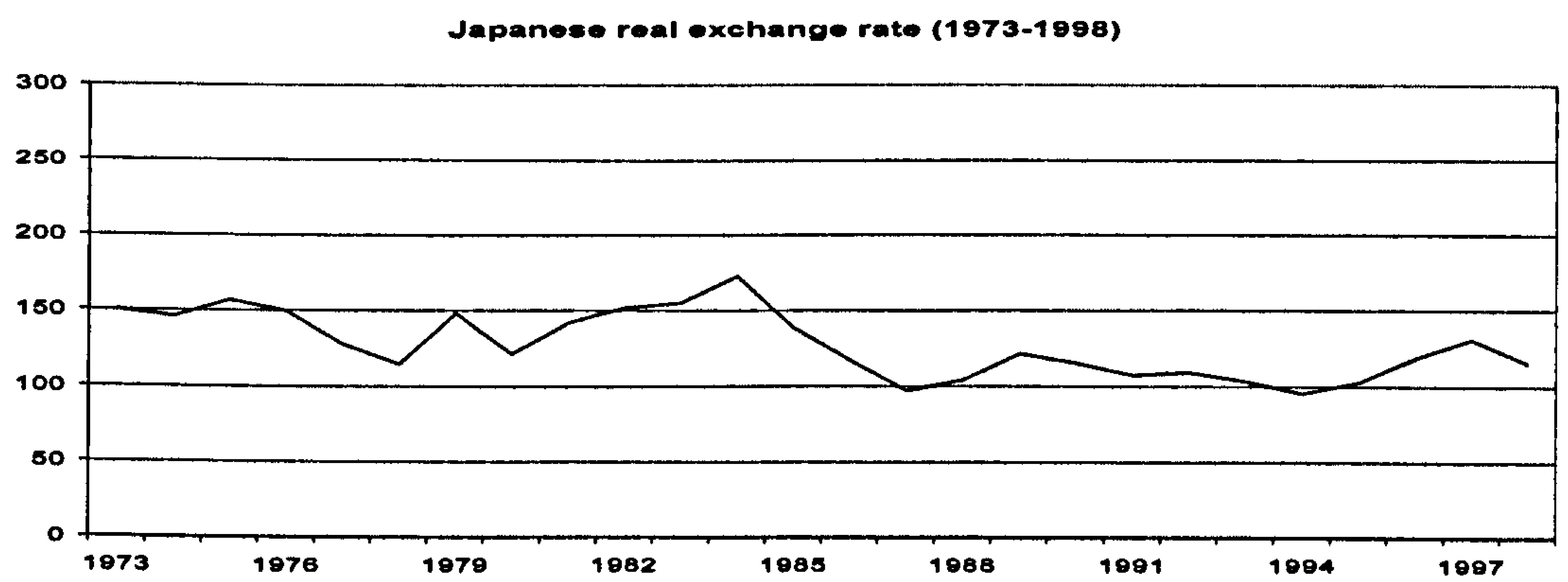


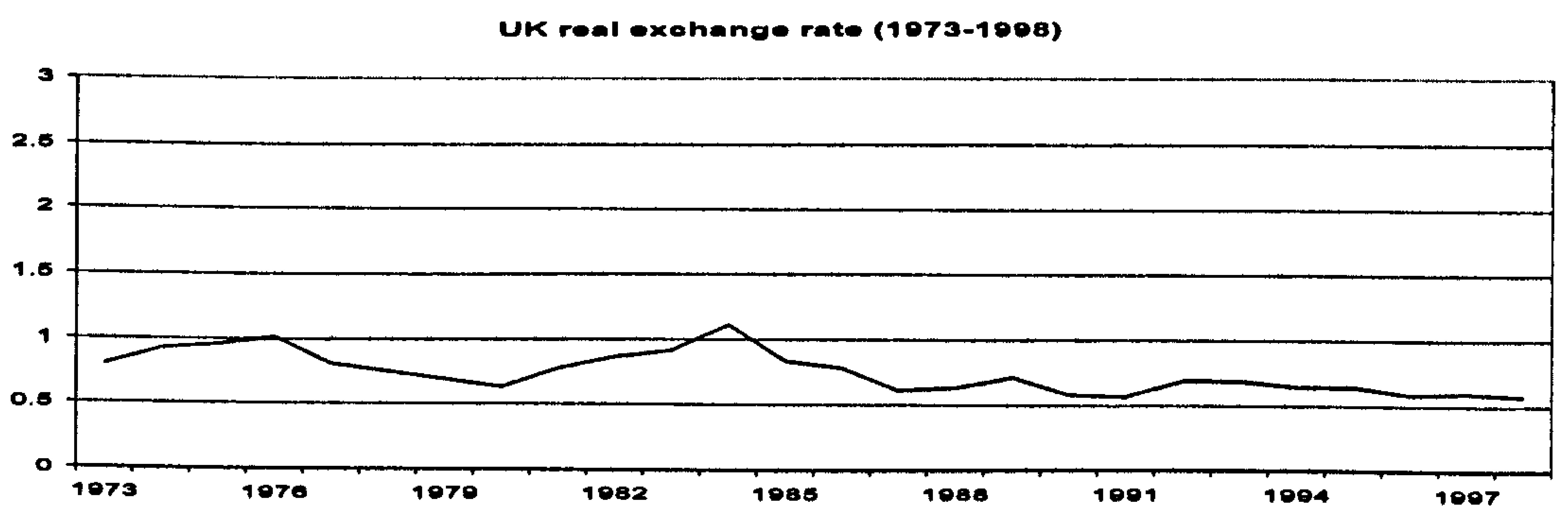
Note: blue lines refer to inflation differential and pink lines refer to relative changes in exchange rates.

Figure 4-2: Real exchange rates over the period 1973-1998

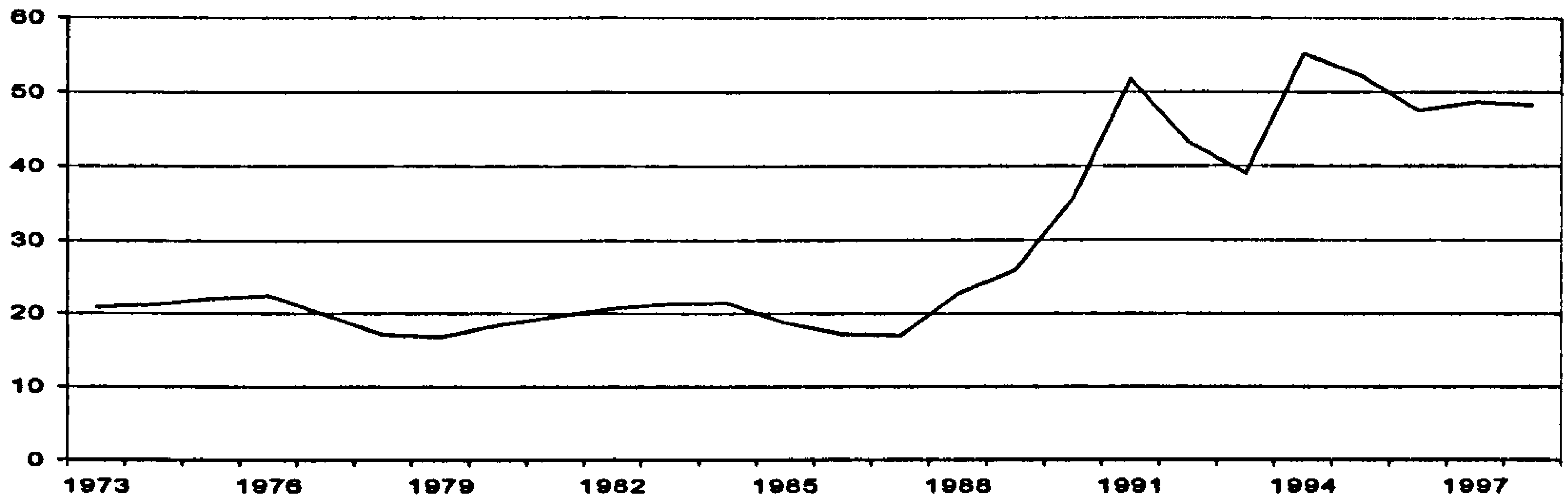




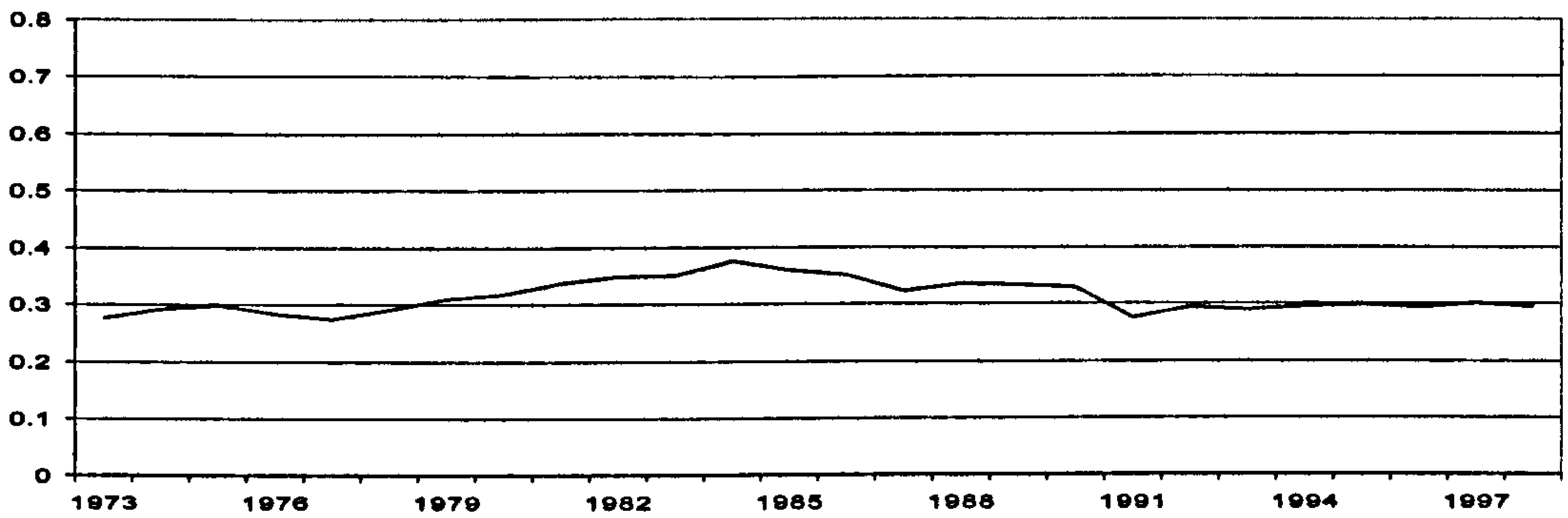




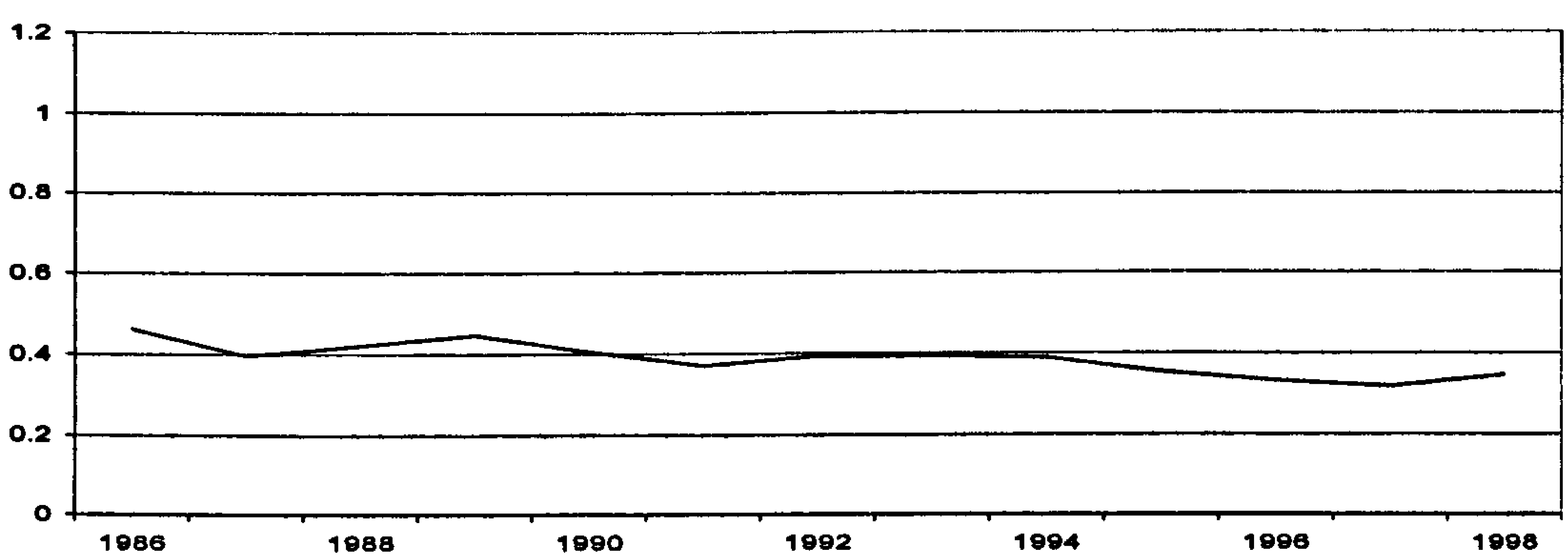
Algerian real exchange rate (1973-1998)



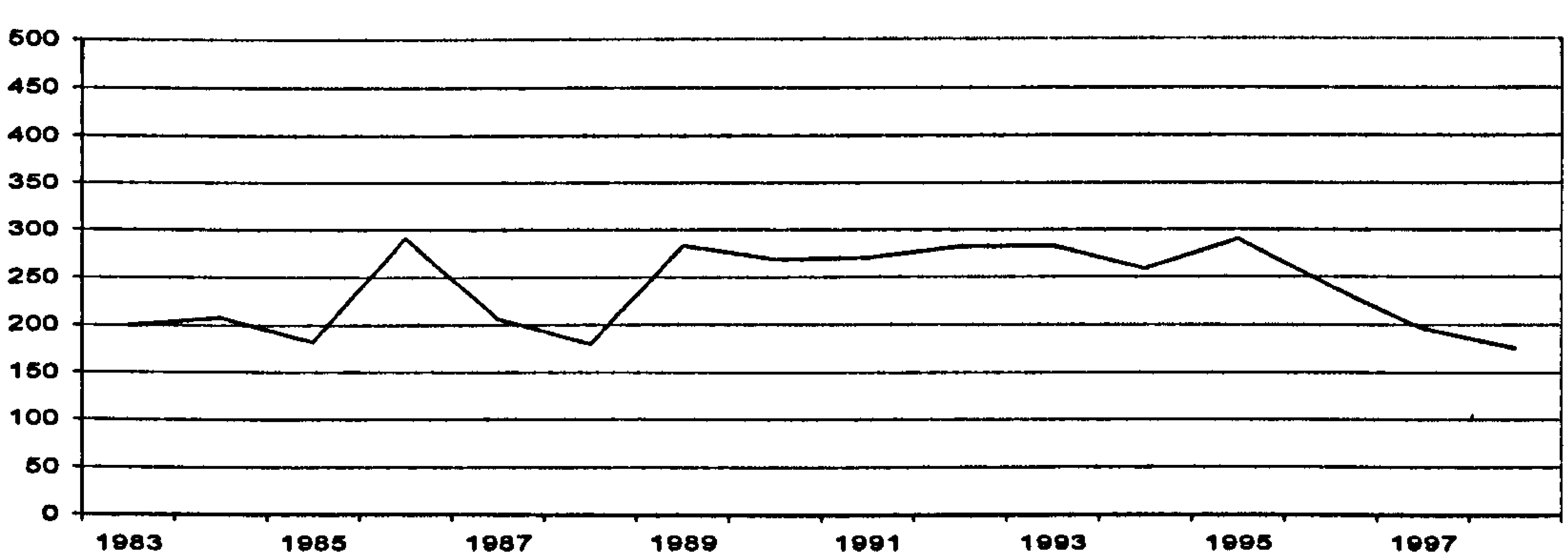
Kuwaiti real exchange rate (1973-1998)



Libyan real exchange rate (1986-1998)



Venezuelan real exchange rate (1983-1998)



Finally, the validity of PPP is further investigated using the ADF unit root test. Table 4-1 contains the results of the ADF tests. Dummy variables were added to the ADF regression to account for outliers in the dates shown in column two. Interestingly, after taking into account the outliers, the *t*-statistics shown in column three indicate that the real exchange rates in all countries are stationary at the conventional levels of significance. These findings globally confirm the stationarity of real exchange rates, implying the validity of the PPP assumption in our sample countries. Therefore, one may conclude that the key assumption in the traditional exchange rate models is generally acceptable in the sample under consideration.

Table 4-1: Results of the ADF unit root tests for the real exchange rate

Country	Outliers	ADF (level with dummies)	Intercept & trend	Decision
Canada	-	-3.58**	C	I(0)
Germany	1984	-3.49**	C	I(0)
Japan	1984	-3.67**	C&T	I(0)
UK	1984	-3.88**	C&T	I(0)
Algeria	1991, 1994	-3.58**	C&T	I(0)
Kuwait	1978	-4.10***	C&T	I(0)
Libya	-	-2.57**	C	I(0)
Venezuela	-	-3.81**	C	I(0)

Note: **, *** indicate level of significance at 5% and 1% respectively.

4.4 Devereux and Lane's model

The question of the principal determinants of exchange rate volatility has attracted a great deal of attention in the international finance literature over the last three decades. However, the results of Meese and Rogoff (1983) which suggested that exchange rate movements are highly unpredictable remain for the most part intact as described by Devereux and Lane (2003). Traditional exchange rate studies exclusively focus on the time series features of exchange rate compared to a single large currency such as the US dollar. Devereux and Lane (2003) took another perspective and concentrated on understanding the determinants that drive bilateral exchange rate volatility across countries. Devereux and Lane started empirically from the optimum currency areas (OCA) theory of Mundell (1961). The OCA hypothesis suggests that there are several criteria that make two countries or entities part of a common currency area. These factors include trade flow interdependence, similarities and dissimilarities in the vulnerability to economic shocks, and the sizes of the economies. The literature on the OCA theory has noticed that, based on their size, openness and the correlation of their business cycle with those of their neighbours, countries may be able to form a common currency according to the OCA theory, but in practice do not. Thus, most observers would dismiss the OCA theory when attempting to interpret the choice of exchange rate regime which suggest that the OCA hypothesis has relatively little predictive power (see Bayoumi and Eichengreen, 1997b). Bayoumi and Eichengreen (1997a and 1997b) therefore examined the role OCA theory can play in explaining the choice of exchange rate regime, and hence, the exchange rate behaviour, using exchange rate volatility rather than the exchange rate regime itself as the dependent variable. Their argument was that the variables indicated by the OCA theory help in interpreting the behaviour of bilateral exchange rates, assuming that these factors inform the decisions made about whether or not to form a currency union, and hence influence exchange rate behaviour across countries. Their results were largely supportive of the idea that the OCA hypothesis can explain a large part of the behaviour of exchange rates.

In addition to the variables suggested by OCA theory, a recent body of work points to the importance of financial elements in understanding exchange rates in emerging market economies. For example, Bernanke et al. (1999) highlighted the importance of the impact of the balance sheet in understanding the properties of business cycles.

Devereux and Lane (2001) and Eichengreen (2002) among others, extended these ideas to an open economy. The main finding of such studies is that balance sheet effects alongside the presence of external debt denominated in foreign currency may significantly affect the way in which exchange rate movements affect an economy. Given the presence of unhedged foreign-currency denominated debt, exchange rate changes can be important through their effects on the financial sector and balance sheets of firms. This is a different way of introducing the cost of exchange rate volatility from that of the conventional theory. Eichengreen and Hausmann (1999) depended on these ideas in stating that many emerging market countries may not be able to tolerate a large degree of exchange rate instability against their main creditors. The phenomenon of fear of floating provides a foundation for the hypothesis that OCA factors alone cannot give a full explanation of exchange rate variability, especially in emerging market economies.

Against this background, Devereux and Lane (2003) put their central hypothesis which is that in addition to the standard OCA factors, bilateral exchange rate volatility is related to the stock of bilateral financial claims across countries. Devereux and Lane developed a simple model of exchange rate choice for a small open economy subject to external disturbances, relying on the recent work of new open economy macroeconomics. The authors started with the simple Redux model which contains a single small economy producing traded and nontraded goods. They distinguished between two groups of countries by assuming that underdeveloped economies face credit constraints in the purchase of intermediate inputs, whereas industrial countries are not vulnerable to such restrictions. More precisely, while developed countries can freely borrow from international capital markets by issuing assets denominated in their currencies, developing countries are constrained in the international capital markets and need to issue debt in foreign currency when borrowing. Devereux and Lane concluded that, for an economy free of credit restrictions, exchange rate movements are desirable in adjusting to external shocks as proposed by OCA theory. In the case of an economy subject to various credit constraints in combination with external debt, exchange rate changes in response to external disturbances significantly decline. Their empirical findings support their expectations.

The literature on OCA theory concentrates on those criteria that make a common currency either more or less attractive across regions or countries. The main suggested factors are the symmetry of shocks to real output, the degree of trade linkages, the usefulness of money for domestic transactions, degree of labour mobility and the automatic stabilizers provided by federal governments. As pointed out by Bayoumi and Eichengreen (1997b) the latter two characteristics are only important in responding to shocks across regions within a country, but not between different countries. Therefore, the first three factors which matter for the responses to disturbances across countries will be considered here.

4.4.1 The theoretical rationale of Devereux and Lane's model

The theoretical justifications of the OCA characteristics and the others provided by Devereux and Lane regarding their relevance to exchange rate movements can be presented as follows.

1-Trade linkages

The transaction costs of international trade, such as the bid-ask spread and related commission fees; go down when a common currency is used in a region. Therefore, the more level trade with a major trading partner is, the more exchange rate stability is desirable. Thus, the size of trade increases with a stable exchange rate (Larrain and

Tavares, 2003). Moreover, the uncertainty concerning future realized profits would be significantly reduced, which may encourage more engagement in international transactions. In other words, a common currency or fixed exchange rates between two or more countries is associated with more certain future proceeds and hence more trade flows between these countries.

2-Asymmetric shocks

Kenen (1969) stated that in well-diversified economies the importance of asymmetric shocks would be of lesser significance than in less-diversified economies. This argument hinges on the hypothesis that positive shocks that hit some exports will be offset by negative shocks that affect other exports; that is, increasing demand for some exports will be compensated by falling demand for other exports. A country that produces a wide variety of goods and then exports diversified products will witness a slower decrease in overall production if it is exposed to reductions in external demand for its products. Thus, Kenen's argument is that good diversification reduces the possibility of asymmetric shocks and decreases their reversal impact. Hence, the larger the asymmetric country shocks, the more dissimilarity will be present between countries and the slower will be real adjustment, the less appropriate is the chance for a currency area and the more appropriate is the option for flexible exchange rate. Therefore, a highly diversified economy may prefer to form a currency area, whereas a less diversified economy needs a flexible exchange rate to adjust for outside disturbances (Horvath and Suomen Pankki. Siirtym talouksien tutkimuslaitos, 2003). Consequently, bilateral exchange rate variability is lower for pairs of countries whose output shocks are strongly correlated (Larrain B and Tavares, 2003).

3-The economy size

Bayoumi and Eichengreen (1997a) used country size to measure the reduction in the transaction costs of a national currency resulting from flexible exchange rates. The costs of a common currency due to losing independent monetary policy should be balanced with the gains, which will be largest for small economies in which there is little scope for using a separate national currency in international transactions. Thus, small economies should gain the most from the common currency's services, which are represented in the unit of account, means of payments and store of value.

Furthermore, Eichengreen and Masson (1998) stated that small economies which trade mainly with large neighbours and/or which have big tourism earnings benefit little from independent monetary policy. Small economies may gain more from decreases in transaction costs and uncertainty, as the common currency is a means of avoiding the weakness of domestic exchange markets, and avoiding substantial fluctuations (Larrain and Tavares, 2003).

McKinnon (1963) suggested that the degree of openness can be used as a measure of the benefits from using a common currency. Consider a small economy which has three categories of goods, namely exportables, x1, importables, x2, and nontradables, x3, where the ratio of x1 and x2 to x3 is high. The price of exportables, p1, and importables, p2, in terms of domestic currency will vary with the exchange rate movements under a regime of flexible exchange rate, whereas the price of nontradables, p3, can be assumed to be constant. Therefore, the goal of stable price level confronts the variability in exchange rate in a small open economy. The case is reversed for a large economy with sizable production of nontradables. Accordingly, small open economies may find it more beneficial to join large currency areas (Horvath, 2003). However, Bayoumi and Eichengreen (1997b) argued that the economic size may be a better measure of the benefits from a stable currency, because a comparison between the gains from national currencies of a large and relatively open economy and a smaller and more closed economy should make clear.

4-External financial links

Countries that are not subject to borrowing constraints are principally the developed economies, and these can easily issue debt in their own currencies in the international capital markets. For them exchange rate adjustments are desirable in response to external shocks. On the other hand, countries that are subject to various borrowing constraints are mainly developing countries which can only issue debt in foreign currencies. For these exchange rate adjustments in response to external disturbances are not desirable. So the external interdependence finance will be inversely linked to the volatility of exchange rate. The rationalization for this is as follows. Importers of intermediate products first need to borrow from foreign banks an amount of money to fund their purchases. Then these importers may buy their intermediates at a specific world price and as a consequence of default risk, the intermediate importers need to pay an extra risk premium per unit of imported inputs. Assuming that risk premium is an increasing function of the amount borrowed relative to net worth of intermediate input sector, which is fixed in terms of the nontradables, the balance sheet position of intermediate sector determines the response of risk premium to exchange rate. Devereux and Lane concluded that the higher the risk premium, the higher the elasticity of intermediate input prices with respect to exchange rates will be which the case for most developing countries is. Thus, given the existence of important credit restrictions for the intermediate inputs sector, exchange rate movements affect the real cost of intermediate inputs via the risk-premium response. Therefore, the direct benefits of exchange rate responses to terms of trade shocks are compensated for by the indirect costs, in terms of a rising risk-premium (cost of intermediate inputs), which itself is sensitive to exchange rate movements

5-Internal financial development

Devereux and Lane argued that when the domestic financial sector is more sophisticated, the financial frictions (risk premium) are likely to be less important. Thus, one expects a negative relationship between the development of the internal financial sector and exchange rate volatility in developing countries, which may suggest that domestic financial development helps to stabilize the exchange rate, for example through adding liquidity to financial markets including the foreign exchange market (Devereux and Lane, 2003).

4.4.2 The empirical specification of Devereux and Lane's model

Devereux and Lane form their exchange rate volatility model as below:

$$Vs_{ij} = b_0 + b_1 trade_{ij} + b_2 cycle_{ij} + b_3 size_{ij} + b_4 finance_j + b_5 extfin_{ij} + u \tag{4.9}$$

The signs of parameters are expected to be as follows: $b_1 < 0, b_2 > 0, b_3 > 0, b_4 < 0$ and $b_5 < 0$.

where Vs_{ij} is the volatility of the nominal bilateral exchange rate between the debtor country j and the creditor country i measured as the standard deviation of the log first difference of the bilateral exchange rate; $trade_{ij}$ is the log of the sum of exports and imports between i and j as a ratio of the country j's GDP, $cycle_{ij}$ is the asymmetric economic shocks approximated by the standard deviation of the growth rate differential between the two countries, $size_{ij}$ is the log of the product of the GDPs of i and j [log(GDP i *GDP j)]

These factors were derived from the OCA hypothesis.

The variables introduced by Devereux and Lane to measure the financial linkages are as follows:

$finance_j$ is the size of domestic financial sector and is measured as the ratio of liquid liabilities to GDP, $extfine_{ij}$ is the financial dependence of country j on country i which is approximated by the own-currency bank claims of country i on country j or by the former plus the long-term debt securities of country j held by country i in constant 1995 US dollars. This variable is entered in the form of $\log(1 + extfin_{ij} / GDP_j)$. The authors added the interaction term $extfin_{ij} * fin_j$ because they expected that external financial dependence would be less relevant for exchange rate policy when the domestic financial sector is more developed. Moreover, they added *GDP per capita* as an extra control variable.

Unlike in the conventional exchange rate models, Devereux and Lane (2003) directly addressed exchange rate volatility rather than its level. They also took a different view by concentrating on understanding the determinants that drive bilateral exchange rate volatility across a large group of countries and developing a model of exchange rate choice for a small economy. Their empirical findings revealed that in addition to OCA factors financial variables significantly help to explain exchange rate volatility.

4.5 Empirical literature on the OCA theory-based exchange rate variability models

As discussed in the previous section, OCA theory was regarded as having little predictive power. However, the OCA hypothesis has recently attracted some attention. For example, Bayoumi and Eichengreen (1997a) attempted to operationalize the theory of optimum currency areas, analysing the determinants of nominal exchange rate variability. They assumed that the factors pointed to by this theory inform the decision of whether to form a currency union, and hence influence exchange rate behaviour across countries. Thus, these factors help in explaining the behaviour of bilateral exchange rates. Bayoumi and Eichengreen used annual data for the period 1983-1992 on bilateral exchange rates for 21 developed countries. They included output disturbances, dissimilarity of export structures as proxies for asymmetric shocks, size of trade and size of the economy as regressors to exchange rate volatility. They found that these variables generally have significant impact on exchange rate variability with the correct expected signs. Bayoumi and Eichengreen (1997b) examined the importance of the factors pointed to by OCA theory in explaining the behaviour of nominal and real exchange rates, and reached similar results to those obtained in Bayoumi and Eichengreen (1997a). Moreover, they added four more variables suggested by the literature on the choice of exchange arrangements and one variable to measure the international regime. Overall, their results showed that the impact of the OCA theory variables on exchange rate variability did not change by expanding the original model. They also found that the results are similar when using either nominal or real exchange rate volatility. In addition, estimating their model using different periods, namely, the 1960s, 1970s and 1980s, showed that the OCA variables were relatively more important in the 1970s and 1980s. However, Bayoumi and Eichengreen indicated that the factors introduced by the OCA hypothesis do not provide a complete interpretation for the variability of bilateral exchange rates.

Bayoumi and Eichengreen (1998) used the same four factors from OCA theory and the same sample countries used in their past study 1997a to explain three dependent variables, including the exchange rate variability, and found similar results. They, again, showed that the variables from the OCA theory provided a better explanation for currency volatility in the 1980s when countries were able to choose their preferred exchange rate regime, compared to the 1960s and 1970s.

Larrain and Tavares (2003) aimed to assess the dollarization versus regional currency union as options for the economies of East Asia, South America and Central America. They used indicators of bilateral integration to examine the determinants of real exchange rate variability within each region and between each region and the US. They examined the impact of the intensity of trade, dissimilarity of exports, asymmetry of output shocks and the size of economy on bilateral real exchange rate volatility for 37 countries in the 1970s, 1980s and 1990s. Their results, in general, supported the theoretical expectations.

Devereux and Lane (2003) tested the determinants of bilateral exchange rate variability using a broad cross section of countries. They included in their model a set of standard OCA criteria, in particular trade interdependence, asymmetric economic shocks and country size and added two other financial series. The first of these captures the degree of internal financial depth and the second represents external financial dependence. Similar to the results of Bayoumi and Eichengreen (1997a 1997b and 1998) and Larrain and Tavares (2003) Devereux and Lane's results showed that the standard OCA variables play their expected role in exchange rate volatility for both developed and developing countries. In particular, trade linkages were negative and significant for industrial countries and some developing countries. The size of the economy was found to be significant and positive in all countries. Cycle was positive and significant for developed economies and negative and significant for developing countries. They also found that the internal finance and external financial linkages were negatively related to bilateral exchange rate volatility in developing economies. However, these variables were positively or insignificantly related to exchange rate variability in industrial nations.

4.6 Comparison between traditional exchange rate models and the Devereux-Lane's model

It is useful to make comparisons among the previous models in order to identify similarities and differences. We shall refer to the previous structural models which investigate the first moment of exchange rates; i.e. the flexible-price, sticky-price monetary models and the Redux model, as the conventional models. We first compare and contrast the conventional models with Devereux and Lane's model and then compare amongst the conventional models themselves.

With regard to Devereux and Lane model and the other set of conventional models, we can note the following points.

- Although the conventional models examine exchange rate levels, they can be reconstructed to examine exchange rate variability, as we did in the beginning of this chapter. Thus, both sets of models can be used to address the link between exchange rate instability and instability or movements in some economic fundamentals. The same measure of volatility can be used in both models.
- The Devereux and Lane model concentrates on explaining the exchange rate volatility of one country against another by considering some specific aspects of relationships, such as trade and financial bilateral factors. On the other hand, the conventional models can investigate the exchange rate volatility either on a bilateral or multilateral basis by considering specific fundamental variables.

- While the conventional models relate the level of exchange rate to a set of fundamentals in terms of relative (differential) forms, the Devereux-Lane model relates the exchange rate volatility to a set of bilateral linkages and a group of factors representing the similarities and dissimilarities between two economies. Thus, the conventional models look at the issue from a comparative viewpoint relative to Devereux-Lane model which looks at the matter from the perspective of the choice of exchange rate regime in the country under consideration.
- Both models can be used for time series data. However, as a result of data availability, the traditional models can be used for as short a horizon as a monthly basis, whereas the Devereux-Lane model, which was originally formed for cross section data, can be applied on time series but only on an annual basis. In this case it is expected, as previously mentioned, that volatility will be smoothed out, since variability gets lower as we move from higher frequency to lower frequency data.
- Since we have the exchange rate in terms of volatility in the Devereux-Lane model and we transferred the traditional models into a volatility form, the relative real income in the standard models and the cycle variable in the Devereux-Lane model appear to be the only similar (if not identical) factors in both models. Thus, such models can be integrated together to yield a combined model which excludes the repeated variable; i.e. the income differential.
- It seems that the Devereux-Lane model fits a small country which depends financially on or has strong trade linkages with a major partner. Thus, it tries to set an exchange rate policy whereby it can keep a stable exchange rate with

that partner. Therefore, it would seem that this model is more suitable in explaining the choice of the exchange rate policy. As such, this point applies more to developing countries, and we expect that the Devereux-Lane model would perform better for developing economies than for developed economies, as was found in Devereux and Lane (2003).

- It seems that the conventional models are suitable for a large country which does not depend financially on or trade mainly with one major partner. Therefore, it can leave its exchange rate to be primarily determined by market forces. Since this point applies more to developed countries, we would expect that the conventional models to perform better for such economies than for developing countries. Moreover, given that most underdeveloped countries link their currencies to a single currency or to a set of currencies of industrial countries, or choose to float them within limited bands[4], their currency prices are not entirely determined by free market forces. In contrast, since the conventional structural exchange rate models mainly rely on the hypothesis of free trade in exchange markets and other markets, such models are less likely to be able to capture and explain exchange rate volatility in developing economies.

As regards the conventional models, we first compare the flexible-price monetary model and sticky-price monetary model. In this regard we could say that the former is suitable for a long-run period when all prices are changeable, and thus PPP holds at all times. On the other hand, the fixed-price monetary model fits the short-run period when prices are sticky; therefore, there are deviations from PPP during this span of time.

[4] Even though some developing countries choose to float their currencies, this experience has been adopted only in recent years.

Regarding the monetary and Redux models, the only difference is that the Redux model links the exchange rate movements to relative consumption rather than relative real income as in the monetary model, in showing the role of the marginal utility of consumption in the decisions to hold money. Consequently, the conventional models, generally, are alike and all of them assume the validity of PPP in the long term at least.

4.7 An augmented equation

From section 4.2 we have exchange rate volatility models which were transformed from the conventional exchange rate models, and which relate exchange rate volatility to the volatility of a set of fundamental differentials. These variability equations will be used to test the volatility of bilateral nominal exchange rates. In section 4.4 we have presented the Devereux-Lane model that relates exchange rate variability to a set of factors which will be used to examine the volatility of bilateral nominal exchange rates. Therefore, we now have two different sets of models which are designed to examine bilateral nominal exchange rate variability in a sample of countries. It may be a good idea to combine both sets of models to form an augmented exchange rate model. We will then see how this hybrid model performs and whether the results obtained by the Devereux-Lane model would change by adding variables from our volatility models. We should bear in mind that such augmented functions are only applicable on an annual basis, since the variables included in the Devereux-Lane model cannot be found on a basis less than annual. Using annual observations implies a small sample consisting of 26 observations at most. As a result of the small number of observations and the great number of variables in the hybrid equations, we may experience a problem of over-parameterization. Thus, results from such an ad hoc augmented model should be interpreted with extra caution.

From the underlying calculations of the variables incorporated in the Devereux-Lane and our transformed models (shown in the following section) note that the variable *cycle* in the Devereux-Lane model is similar (or the same as) the variable $V\tilde{y}$ in our volatility models, hence, we will use only one of these to avoid the problem of multicollinearity. Therefore, the equations of the augmented model can be written as follows:

Flexible price-Devereux-Lane augmented model

$$\begin{aligned} Vs_{ij} &= b_0 + b_1 trade_{ij} + b_2 cycle_{ij} + b_3 size_{ij} + b_4 finance_j \\ &+ b_5 extfin_{ij} + b_6 V\tilde{m} + b_7 V\tilde{r} \end{aligned} \tag{4.10}$$

sticky price-Devereux-Lane augmented model

$$\begin{aligned} Vs_{ij} &= b_0 + b_1 trade_{ij} + b_2 cycle_{ij} + b_3 size_{ij} + b_4 finance_j \\ &+ b_5 extfin_{ij} + b_6 V\tilde{m} + b_7 V\tilde{r} + b_8 V\tilde{\pi}^e \end{aligned} \tag{4.11}$$

Redux-Devereux-Lane augmented model

$$\begin{aligned} Vs_{ij} &= b_0 + b_1 trade_{ij} + b_2 cycle_{ij} + b_3 size_{ij} + b_4 finance_j \\ &+ b_5 extfin_{ij} + b_6 V\tilde{m} \end{aligned} \tag{4.12}$$

The target equations to be empirically tested, therefore are now (4.5), (4.7), (4.8) for equations based on the traditional models using monthly and quarterly data. The volatility measures we intend to use are the standard deviation of percentage change in a series, the simple proxy and the ARCH model wherever appropriate, and equation (4.9) for the Devereux-Lane model in addition to our hybrid equations (4.10), (4.11) and (4.12) using annual data with the standard deviation as a proxy for volatility.

4.8 Data sources and definitions of variables

The main source of our data is the IMF's international financial statistics from the two versions, the publications and CD-ROM 2000. However, some data were collected from other sources, such as national sources and other international organizations. Our sample, as mentioned earlier, consists of the four oil-based developing countries; i.e. Algeria, Kuwait, Libya and Venezuela, and the four developed economies; i.e. Canada, Germany, Japan and the UK and the numeraire country is the US. Our sample lasts from the advent of floating exchange rate regimes in 1973 through 1998 using monthly and quarterly data for the equations (4.5), (4.7) and (4.8) and annual observations for equations (4.9)-(4.12). However, for the Venezuelan Bolivar and Libyan dinar exchange rates against the US dollar, the samples start in 1989 and 1986 respectively, because these exchange rates were fixed against the dollar until those dates. The Venezuelan Bolivar, however, was subject to several devaluations prior to 1989 and multiple exchange rate system was applied in the 1980s. Thus, since the Bolivar was not utterly fixed to the US dollar in that period of time, we may expand our sample for Venezuela especially when using annual observations.

We first define the variables used in the conventional models-based equations.

All variables are formed in terms of percentage changes before calculating the volatility measures. The way in which volatility is calculated was explained during the discussion of volatility measures in chapter three.

4.8.1 Proxies of the exchange rate and its determinants

- The exchange rate variable is the log nominal exchange rate defined as the number of domestic currency units per one US dollar. Our concentration on the volatility of nominal exchange rates instead of the real rates is due to the fact that nominal rates give an easier benchmark for comparison to a single currency (Bayoumi and Eichengreen, 1997). Furthermore, in related work Bayoumi and Eichengreen found both nominal and real exchange rates provided similar results. Moreover, Bini-Smaghi (1991) argued that the risk considered should concern nominal rather than real exchange rate, as the latter depends not only on the variance of the nominal rate but also on that of relative prices. Therefore, to distinguish between risk due to exchange rate changes and risk due to price movements, the nominal exchange rate should be considered. In addition, Medhora (1990) stated that given the relatively short time horizon of traders; it is more relevant for them to look only at nominal rate changes, as they move faster and more frequently than prices on a day to day basis. We compute the first difference of the log s, which is then used to calculate the volatility measure.

- Money supply is the log nominal money supply which is represented by seasonally adjusted M1, except for the UK where the narrow definition of money, M0, is used, because there is no data available about M1. Moreover, M2, money plus quasi money, does not exist for the whole period on a monthly basis for this country. Then we subtract the US log money supply from each country's log money supply, in order to derive the log money supply differential. Afterwards we take the first difference of the product variable which is then used to calculate the volatility measure.
- Monthly gross domestic product (GDP) series does not exist for both developed and developing economies, and quarterly series do exist for developed countries but not for developing countries. Therefore, a proxy is required. Most researchers use the index of industrial production (IP) as a proxy measure for the trend of GDP. Although the IP is a more restrictive

measure representing only the level of manufacturing, it should reflect the overall trend in GDP. However, this series is not available for our developing country sample, thus we approximate the GDP level by the crude petroleum production index (CPP) for the developing country sample because it represents the major source of GDP. Accordingly, we will use the IP and CPP as proxies for the GDP in the developed and developing economies respectively. We then calculate the differential log IP (i.e. subtract the US IP from the country under consideration's IP (or CPP)). After taking the first difference of the resulting term, we use it to produce the volatility measure.

- The interest rate is the Treasury bill rate for Canada, Germany, the UK and the US, the call money rate for Japan and the discount rate for the sample of the developing economies. Such different proxies are due to their availability. Although, these proxies differ from each other, they may, in general, represent the cost of holding money and move more or less together. In this case we calculate the percentage change of interest rate for each country, computing the differential term and then using it to calculate the volatility measure.
- The expected inflation rate is represented by the current inflation rate[5]; the inflation rate is calculated as the first difference of the log consumer price index, CPI. Then we take the differential term which is used to calculate the volatility measure.

[5] Proxies for the unobservable expected inflation rate are usually the long-term interest rate, the preceding twelve month period CPI or WPI inflation rates or with an inflation rate autoregression. Meese and Rogoff (1983) reported that using the long-term interest rate, a moving average of past inflation or future inflation proxies give similar results. Frankel (1981) stated that proxying the expected inflation by long-term interest rates or lagged actual inflation rates may introduce errors in variables and argued that the choice of the proxy does not seem to be crucial and the coefficient of the expected inflation rate is the one that has always appeared to be most robust across sample period and estimation technique. The current inflation rate has this merit in our sample.

- The consumption variable is the log per capita real consumption. We first calculate the differential term and take its first difference and using this to calculate the volatility measure.

4.8.2 The proxies used in Devereux-Lane's model

- *trade* as in Devereux and Lane is the log of the sum of bilateral exports and imports between the numeraire and the country concerned as a ratio to the GDP of the country under consideration. The sources are various issues of the UN's international trade statistics yearbook.
- *cycle* is the standard deviation of the differential growth rate; and growth rate is measured as the first difference of log real GDP.
- *size* is the log of the product of GDP's, where GDP is measured in constant US dollars.
- *Finance* is the internal financial depth, which is approximated by the ratio of private credit to GDP as in Levine et al. (2000) who argued that although liquid liabilities are commonly used as a measure of the financial sector development as in the Devereux and Lane (2003) this has some shortcomings. It may not precisely represent the effectiveness of the financial sector in improving informational asymmetries and facilitating transaction costs. Also, since liquid liabilities equal currency plus demand and the interest-bearing liabilities of banks and non-bank financial intermediaries, they include deposits of one financial intermediary in another. Thus it involves double counting. Levine et al. prefer to use the private credit as an indicator of the financial development. The private credit equals the value of credits given to the private sector by financial intermediaries divided by GDP. On the one

hand this isolates credit issued to private sector from that issued to governments and public firms and excludes credits issued by the central bank and development banks. On the other hand, although private credits do not evidently measure the improvement in information and transaction costs, higher levels of private credit can be explained as an indicator of higher levels of financial services and thus higher financial intermediary development (Levine et al, 2000).

- *extfin* represents the financial dependence of the country concerned j on the numeraire country i, which is approximated by either the sum of the own-currency bank claims of country i on country j and the long-term debt securities of country j held by country i, or by only the own-currency bank claims of country i on country j in constant 1995 US dollars as in Devereux and Lane's paper. Since Devereux and Lane used a cross section of countries, they needed no series of observations about bilateral external debt and they required only one observation for this proxy. For our method, in which we use a time series analysis, we require a series of data which lasts for the targeted period of 1973-1998. Since there is no published series of the above-mentioned variables, we have to use the total external debt which is denominated in US dollars, presuming that most of the external debt of developing countries is denominated in US dollars. Devereux and Lane stated that in most Latin American countries bond debt is denominated in US dollars, while the industrial nations bonds are often dominated in domestic currency. We generalize this assumption for our entire developing economies sample. The source of this variable is the World Bank (1998) World Development Indicators CD-ROM.

Some variables, such as the population (used to calculate per capita consumption) and inflation rate in some countries were transferred from annual or quarterly data into monthly data using interpolation by regression with time trends.

The standard deviation-based exchange rate volatility is computed as the standard deviation of the first difference log of the series. The standard deviation was calculated over three months to obtain quarterly data. For the differential terms the volatility is the standard deviation of the first difference of the differential terms. Some data was interpolated, as in the Kuwaiti case during the Gulf War.

Chapter five

Econometric Methodology

5.1 Introduction

The previous chapter presented the exchange rate volatility equations to be empirically estimated using time series data. This chapter outlines the econometric methodology to be used for the empirical research presented later in the thesis.

Most empirical studies that investigate the determinants of the exchange rate level, in particular those conducted prior to 1987, implicitly assume that the time series used for estimation are stationary (i.e. the data fluctuates around a constant mean with a finite constant variance and covariance).

It is widely known, however, that most economic time series are nonstationary, that is, they possess a unit root, (see for example, Hwang, 2001). This fact makes inferences from the regression of nonstationary variables misleading. Namely, if a nonstationary variable is regressed on another variable which is either stationary or nonstationary, a spurious regression is likely to emerge, in that the conventional t-values may tend to indicate a relationship between the variables when in fact no meaningful causal relationship exists. To shed some light on the concept of spurious regression, a simple exposition of this concept is given next.

Suppose we have two random walk models: $y_t = y_{t-1} + u_t$ and $x_t = x_{t-1} + v_t$ where u_t and v_t are white noise errors; i.e. normally distributed with zero mean and constant variance; and they are independent and serially uncorrelated. These two time series are nonstationary; more precisely they are integrated of order one. Since y_t and x_t are uncorrelated, any regression between them must produce zero R^2 and insignificant estimated coefficients. However, if their regression produces nonzero R^2 and statistically significant coefficients, this is well-known as a nonsense or spurious regression, which was first discovered by Yule (1926). Yule showed that spurious correlations may easily be found between independent nonstationary variables even if the sample size is very large. Half a century later, Granger and Newbold (1974) generated 100 pairs of independent random walk variables, such as those shown above, and ran linear OLS regressions. They computed the conventional t-values to test the significance of each regression. Their main result was that in 77 out of 100 simulations the statistics were highly significant, leading to a mistaken rejection of the null hypothesis of no relationship. Moreover, Granger and Newbold found very low Durbin-Watson statistic (DW) values which lead to underestimations of the standard errors, and hence, overstated t-values. Adding more random walk regressors increased the percentage of incorrect inferences (Johnston and DiNardo, 1997).

In a theoretical explanation of this issue Phillips (1986) has shown that in the regressions of independent random walk time series, the usual t-ratio and F-ratio significance tests do not possess limiting distributions but diverge as the sample size gets larger and larger. Therefore, the bias in this test for rejecting the hypothesis of no relationship will increase with sample size. He also showed that the DW statistic converges in probability to zero, while the determinant coefficient (R^2) has a non-degenerate limiting distribution as the sample size increases.

This phenomenon provides strong evidence that one should be extremely cautious in conducting regression analysis based on time series that exhibit nonstationarity.

Taking the first difference of a first order integrated variable produces a stationary variable. However, this approach excludes the opportunity of estimating any relationships between the variables in terms of levels, although some cointegration techniques imply the existence of such relationships (Davidson and MacKinnon, 1993). It also results in a loss of valuable long run movement in the data (Maddala, 2001). In addition, the difficulty of interpreting the results when the first differenced variables are used in the regression can make this a complicated task in most cases. Furthermore and may be less seriously, losing an observation and in turn reducing the degrees of freedom could be a problem, especially in small samples. As a result researchers have focused on finding long run relationships (cointegrating vectors) without the need to turn nonstationary variables into stationary ones by differencing them.

A great deal of research has therefore been devoted to finding a way to examine the long run relationships between nonstationary series. This literature has produced what is called cointegration analysis. To address this issue in more detail, we begin with a few definitions which were provided by Engle and Granger (1987).

Integration: a series without a deterministic component which has a stationary ARMA representation after differencing d times is said to be integrated of order d, denoted $x_t \sim I(d)$. Thus if d=0 the series is stationary or $I(0)$; if d=1 the series is integrated of order one (nonstationary) or $I(1)$, and if d=2, the variable is integrated of order two or I(2) and so forth.

Cointegration: the components of a vector x_t are considered to be cointegrated of order d, b ($x_t \sim I(d,b)$), if all components of x_t are I(d) and there exists an α such that $z_t = \alpha' x_t \sim (d-b)$, b>0. The vector α is cointegration vector. In the simplest case when d=b=1, if the components of x_t are all $I(1)$, the cointegration means the equilibrium error would be $I(0)$ and z_t will not drift away from the mean, thus there is a long run equilibrium relationship between the components of x_t. On the other hand, if all components of x_t are $I(1)$ and a linear combination of them is also $I(1)$, x_t is not cointegrated and its components drift away from each other more and more as time goes on. Therefore, there is no long run relationship between these components. In this case the relationship obtained, if any, from a regression between these components is spurious (Maddala, 2001).

Estimating and testing the long run equilibrium relationships has been conducted under the headings of *cointegration,* for example in Engle and Granger (1987) Stock (1987) and Johansen and Juselius (1990); *regression with integrated regressors,* for example in Phillips (1988) Park1(992) and Sims et al. (1990); and *common trends,* for example in Stock and Watson (1988) (Johansen, 1991).

Engle and Granger (1987) proposed a residual-based technique by which an equilibrium relationship can be estimated. If the relevant variables are integrated of the same order and the residuals from the cointegration regression are $I(0)$, then a statistically significant unique cointegrating vector can be found. After estimating the cointegrating equation, Engle and Granger suggested estimating the short run relationship through variants of the error correction model (ECM) by a two stage estimation method using the estimated coefficients from the cointegrating regression. The Engle and Granger two-step procedure has the advantage of the use of the OLS method for estimating the long run equation from which the short run dynamics can be modelled through the ECM methodology. However, this approach is dependent upon the assumption of the uniqueness of the estimated cointegrating vector. The test procedures do not have well defined limiting distributions, and thus, testing for cointegration is not a straightforward procedure. It is also applicable only when the relevant variables are integrated of the same order. Thereafter a contribution by Johansen (1988) provided a maximum likelihood estimation procedure by which some of the above problems are solved. This model, however, has its own disadvantages. In particular it fails to include I(0) variables in the regression. This would be a crucial drawback if economic theory indicates that these variables play an important role in defining equilibrium relationships. A recently developed method called the bounds testing approach has the capability to overcome this particular problem by including both I(0) and I(1) variables simultaneously in the regression. Details of such cointegrating vectors will be discussed in later sections.

The purpose of this chapter is to introduce the basic framework of the two main econometric methods which have been developed for estimating long run relationships between nonstationary time series. These methods include the Johansen multivariate cointegration approach, which is most often used in the empirical literature because of its superiority to some earlier approaches. The second method was developed by Pesaran et al. (2001). The main advantages and disadvantages of both approaches are also presented, and these two procedures are then used, when appropriate, to investigate the existence of possible level relationships between the variables by estimating our targeted equations introduced in chapter four.

5.2 The Johansen multivariate cointegration procedure

Johansen (1988) and Johansen and Juselius (1990) have developed a multivariate technique that provides maximum likelihood estimates of all the possible cointegrating vectors that may exist between a set of series. This procedure uses a general vector error correction model (VECM) and a reduced regression model through which the number of cointegrating vectors is determined by the rank of the long run matrix.

Despite the fact that there are several approaches to multivariate cointegration as mentioned above, the Johansen method is the most commonly applied technique and is widely programmed in econometric software packages (Patterson, 2000). Johansen begins with a general unrestricted vector autoregressive model:

$$X_t = A_1 X_{t-1} + A_2 X_{t-2} + \ldots\ldots + A_k X_{t-k} + \varepsilon_t \qquad (5.1)$$

where X_t is an $n \times 1$ vector of a set of variables which are commonly assumed to be I(1), A is a $n \times n$ matrix of unknown parameters and ε_t is a vector of Gaussian error terms.

Letting Δ represent the first difference operator, (5.1) can be reformulated in terms of a generalized vector autoregressive error correction form:

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \Gamma_2 \Delta X_{t-2} + \ldots\ldots + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-k} + \varepsilon_t \qquad (5.2)$$

where $\Gamma_i = -(I - A_1 - A_2 - \ldots\ldots - A_i)$, i =1, 2, ..., k. and

$\Pi = -(I - A_1 - A_2 - \ldots\ldots. - A_k)$.

Γ_i defines the dynamic adjustment in X_t and Π defines the long run solution to X_t. Π can be decomposed into two matrices such that $\Pi = \alpha\beta'$, where α is a matrix of error correction coefficients, and thus it reflects the speed of adjustment from disequilibrium to equilibrium, and β is a matrix of long run equilibrium coefficients. The lag length of the VAR model is set to ensure that ε_t is a white noise error term.

If X_t is I(1), then all ΔX_{t-i} will be stationary, therefore, for equation (5.2) to determine an equilibrium relationship, ΠX_{t-k} has to be I(0) if ε_t is stationary. The rank of Π determines the number of distinct cointegrating vectors (r) amongst the variables.

If the Π matrix has rank zero, implying no linear combination of X_t that are I(0), this means that there are no cointegrating vectors. Then the appropriate model is a VAR in first differences form.

If Π has full rank (r=n), it implies that all the variables in X_t are stationary. In this case there is no need for a VECM since there is no spurious regression and then the appropriate method is to estimate the standard Sims-type VAR in levels.

If Π has reduced rank 0<r<n, there are $r \leq n-1$ cointegrating vectors present in β. For example, if rank $\Pi=1$, there is a unique cointegrating vector, and if rank $\Pi=2$, there are two cointegrating relationships, and so forth.

A reduced rank regression approach can be used to identify the number of cointegrating vectors. This involves writing Eq (5.2) in the following form:

$$\Delta X_t + \alpha\beta' X_{t-k} = \Gamma_1 \Delta X_{t-1} + \Gamma_2 \Delta X_{t-2} + \ldots\ldots + \Gamma_{k-1}\Delta X_{t-k+1} + \varepsilon_t \tag{5.3}$$

Now by regressing the short run dynamics on ΔX_t and X_{t-k} respectively we obtain:

$$\Delta X_t = P_1 \Delta X_{t-1} + P_2 \Delta X_{t-2} + \ldots\ldots\ldots + P_{k-1}\Delta X_{t-k+1} + \eta_{0t} \tag{5.4}$$

$$X_{t-k} = T_1 \Delta X_{t-1} + T_2 \Delta X_{t-2} + \ldots\ldots\ldots + T_{k-1}\Delta X_{t-k+1} + \eta_{kt} \tag{5.5}$$

While equation (5.4) represents the stationary changes, ΔX_t, adjusted for short run dynamics, equation (5.5) shows the stationary level term, X_{t-k}, is corrected for short run dynamics. η_{0t} and η_{kt} are the unexplained components of ΔX_t and X_{t-k} by the short run dynamics. If Eq (5.4) and (5.5) are estimated separately, the OLS can be used to produce consistent estimates of the long run coefficients. However, since the combination of the cointegrating vectors links the variables together, maximum likelihood estimation must be used. To determine the cointegration rank, the maximum likelihood estimate of β (the cointegrating vector) can be obtained as the eigenvectors corresponding to the r largest eigenvalues from solving the following eigenvalue problem:

$$\left|\lambda S_{kk} - S_{k0}S_{00}^{-1}S_{0k}\right| = 0 \tag{5.6}$$

where S_{00} is the residual moment matrix from the least squares regression of (5.4), S_{kk} is the residual matrix from the regression of Eq (5.5) and S_{0k} is the cross-product moment matrix.

Using these eigenvalues, Johansen and Juselius (1990) provided two likelihood ratio tests, the trace and maximum eigenvalue tests, to test the number of cointegrating vectors.

The first is the trace test (also known as the likelihood test statistic) which is based on the stochastic matrix and is defined as:

$$\lambda_{trace} = -\mathrm{T}\sum_{i=R+1}^{n}\log(1-\hat{\lambda}_i) \tag{5.7}$$

where $R = 0,1,\dots\dots,n-1$, T is the number of usable observations and $\hat{\lambda}_i$ is the estimated value of the characteristic roots (the eigenvalues) from the estimated $\alpha\beta'$ matrix.

The critical values of the trace test are used to test the null hypothesis that the number of distinct cointegrating vectors is at most $r \leq R$ (there is no cointegrating vector) against the alternative that $r \geq R+1$ (there is one or more cointegrating relationships).

The second test is the maximum eigenvalue test which is based on the following function:

$$\lambda_{\max} = -\mathrm{T}\log(1-\hat{\lambda}_{R+1}) \tag{5.8}$$

This statistic tests the null hypothesis that the number of cointegrating vectors is r against the alternative of r= $R+1$.

In both of these tests the number of cointegrating vectors is identified by a sequential testing technique until the null hypothesis is accepted.

The multivariate Johansen approach is superior to the two-step Engle and Granger method in the following aspects: (i) it fully captures the underlying time series properties of the data; (ii) it does not a priori assume the presence of at most a single cointegrating relationship and explicitly tests for the number of cointegrating vectors; (iii) in contrast to the Engle and Granger method which is sensitive to the choice of the dependent variable, the Johansen technique assumes all variables to be endogenous; (iv) it offers a test statistics for the number of cointegrating vectors which has an exact limiting distribution; and (v) it was enhanced with a set of tests which can be used to directly test the linear restrictions imposed by economic theory.

In spite of its superiority to the Engle and Granger technique, the Johansen method raises some difficulties that empirical researchers should take into account. These problems can be listed as follows. Firstly, this approach is only applicable when the time series included are all integrated of the same order, i.e. I(1). Thus, if all or some of the variables under consideration are not I(1), the problem of spurious regression is likely to emerge. Secondly, this procedure is sensitive to the misspecification of the lag length of the VAR model. Therefore, one must be cautious in choosing the appropriate lag order. Thirdly, if more than one cointegrating vector has been found, the researcher has to face the difficulty of meaningful economic interpretation that results from the identification problem. Because of the seriousness of this problem, it will be discussed in more detail in the following section.

5.2.1 The identification problem

In the case of a single structural equation or a unique cointegrating vector, the economic interpretation of the long run coefficients can be achieved by means of normalization. Normalization is usually implemented by restricting one of the estimated coefficients (often the dependent lagged variable coefficient of the structural equation) to equal -1 and dividing all other coefficients by the negative value of the chosen normalizing variable coefficient. However, if one endogenous is excluded and all the exogenous variables are included in the system, the coefficients of the cointegrating vector will be a linear combination of the behavioural and reduced form coefficients. With multiple cointegrating vectors, economic interpretation is impossible without imposing additional restrictions to those usually used in cointegration analysis in order to achieve an exact identification. The presence of a cointegrating vector implies a stable long run equilibrium relationship among jointly endogenous variables arising from restrictions imposed by the economic theory of the long run relationship. If more than one cointegrating vector is found, there is no longer a unique equilibrium relationship to which the error correction model converges (Tawadros, 2001). Choosing a particular cointegrating equation among multiple cointegrating vectors implies that the individual chosen model is indeed valid. This issue is not discussed in Johansen method because it considers all variables in the system to be endogenous and does not categorize them into endogenous and exogenous variables, which is an essential procedure in estimating a single reduced form or structural equation. Therefore, it is impossible for a researcher to identify separate individual equilibrium relationships, because the Johansen technique can only impose and test the same restrictions across all the cointegrating vectors simultaneously. Consequently, Wickens (1996) has shown that unless prior information is available to impose constraints on a reduced form VECM, the structural system cannot be identified. Moreover, given the relationship between the reduced form and structural coefficients, an economic interpretation is not possible of the cointegrating vectors derived from the Johansen approach (Wickens, 1996). Therefore, it is unreliable to rely on the Johansen method in estimating a long run relationship among variables that have different orders of integration in addition to the identification and interpretation problems associated with this approach. Thus, this method should be applied only when its requirements are met. It is essential to look for another method that avoids or at least minimizes the difficulties related to Johansen. Fortunately, the relatively recent bounds testing approach achieves this task insofar as it has some feature which enable it to give more robust results than those of Johansen's. The following section provides more details about this recently developed approach.

5.3 *The bounds testing approach*

The main standard cointegration analysis methods, the two-step Engle and Granger cointegration approach and the Johansen multivariate cointegration approach, focus on cases in which the underlying time series are integrated of order one. This inevitably incorporates a certain degree of pre-testing of the order of integration, and hence introduces more uncertainty degree into the analysis of levels relationships (Cavanagh et al., 1995). Recently, Pesaran et al. (2001) proposed a new procedure known as bounds testing approach which avoids the pre-testing problem. The bounds testing procedure is used for testing the existence of a single long run relationship between a set of underlying variables (Pesaran et al., 2001). The bounds testing method to cointegration is employed within an autoregressive distributed lag (ARDL) framework. The statistic underlying the ARDL procedure is the Wald or *F*-statistic in a generalized Dickey-Fuller type regression, which is used to test the significance of lagged levels of the time series under consideration in a conditional unrestricted equilibrium correction model (UECM). This approach has two main advantages over the traditionally used cointegration methods. Firstly, it tests for the presence of a single level relationship between a dependant variable and regressor(s) when it is not known for certain whether the underlying variables are purely $I(0)$, purely $I(1)$ or mutually cointegrated. It is well known that standard cointegration analysis methods are applicable to nonstationary variables that are integrated of the same order. The pre-testing is particularly problematic in the analysis of unit root cointegration where the unit root tests typically have low power, and there is a switch in the distribution function of the test statistics as one or more roots of the regressor's process approach unity. Furthermore, the UECM is likely to have better statistical properties than the two-step Engel-Granger approach, since unlike the two-step procedure the UECM does not push the short run dynamics into the residual terms (Narayan and Smyth, 2005). Secondly, the bounds testing approach is robust in the case of small sample cointegration analysis (Tang, 2003). It is said that finite sample analysis can bias the likelihood ratio (LR) test of Johansen's (1988) method towards finding long-run relationship either too often or too infrequently (Cheung and Lai, 1993). In addition, the conventional cointegration methods become unreliable with small sample data (Mah, 2000).

The bounds testing approach involves two stages. In the first stage one should test for the existence of a long run relationship. Once an equilibrium relationship has been established, a further two step procedure is performed in estimating the long run coefficients.

To illustrate the bounds testing procedure, assume that we have X_t as a vector of I(d) independent variables, where $0 \le d \le 1$. The conditional autoregressive distributed lag error correction model (ARDL-ECM) formulation can be written as follows:

$$\Delta y_t = c_0 + c_1 t + \pi_y y_{t-1} + \pi_x X_{t-1} + \sum_{i=1}^{p-1} \varphi_i \Delta y_{t-i} + \sum_{j=1}^{q-1} \delta_j' \Delta X_{t-j} + \gamma' \Delta X_t + u_t \quad (5.9)$$

where c_0 is the intercept, t is the trend component, and π_y and π_x are the long run coefficient matrices for y_{t-1} and X_{t-1}. The short run dynamic structure of Δy_{t-i} and ΔX_{t-j} is set to ensure the residuals, u_t, are white noise error.

Equation (5.9) was written under the assumption that X_t is the long run forcing variable for y_t as there is no feedback from the level of y_t. This assumption restricts cointegration to cases where there is at most one conditional level relationship[6] between y_t and X_t regardless of the level of integration of the process X_t (Pesaran et al., 2001).

In order to test for the absence of a level relationship between y_t and X_t in equation (5.9), it is estimated by OLS, and the F-statistic for the joint hypothesis of $\pi_y = 0$ and $\pi_x = 0$ is calculated. The null hypothesis is tested through the exclusion of the lagged level variables y_{t-1} and X_{t-1} in equation (5.9). The null hypothesis is that $\pi_y = 0$ and $\pi_x = 0$ against the alternative that $\pi_y \ne 0$ and $\pi_x \ne 0$, $\pi_y \ne 0$ and $\pi_x = 0$ or $\pi_y = 0$ and $\pi_x \ne 0$.

[6] Since we are interested in the variances rather than the levels, this assumption may be more realistic.

The statistic underlying the bounds testing technique is the familiar Wald or F-statistic. Pesaran et al. (2001) have shown that the asymptotic distributions of the F-statistic are non-standard under the null hypothesis that there is no levels relationship between the variables included regardless of whether the regressors are $I(0)$, $I(1)$ or mutually cointegrated. Therefore, they provided two sets of asymptotic critical values that establish lower and upper bounds of significance which assume that all the regressors are, on the one hand $I(1)$, and, on the other, $I(0)$.

The asymptotic critical value bounds for the F-statistic are cited in Pesaran et al. (2001), pp. 300–301, Table CI(i)–(v)), for some significance level (10%, 5%, or 1%). If the computed F-statistic (test statistic) exceeds the upper critical value, $I(1)$, then it is possible to reject the null hypothesis. In case that the computed F-statistic falls below the lower critical value, $I(0)$, the null hypothesis of no cointegration cannot be rejected. When we find that the computed F-statistic falls within the critical value bounds, a conclusive inference cannot be made. Here, the time series properties of the data must be known before any conclusion can be drawn, as the bounds test is only applicable for $I(0)$ or $I(1)$ regressors.

The critical values are also available to accommodate a range of different deterministic components: (i) no intercept and no trend; (ii) restricted intercept and no trend; (iii) unrestricted intercept and no trend; (vi) unrestricted intercept and restricted trend and (v) unrestricted intercept and unrestricted trend. They also introduced the critical value bounds for the t-statistic associated with the coefficient of the lagged dependent variable in an unrestricted conditional ECM for the confirmation of the F-statistic results. The asymptotic distribution of this statistic is given for cases where all the regressors are $I(1)$, and where they are $I(0)$ or mutually cointegrated.

The asymptotic critical value bounds for the *t-test* are cited in Pesaran et al. (2001) pp. 303–304, Table CII(i), (iii) and (v)), for some significance level (10%, 5%, 2.5% and 1%). The critical values are available to accommodate a range of different deterministic components: (i) no intercept and no trend; (iii) unrestricted intercept and no trend and (v) unrestricted intercept and unrestricted trend.

Pesaran et al. (2001) stated that if the null hypothesis of no cointegration using the bounds procedure based on the Wald or F-statistic of Eq (5.9) is not rejected, one should proceed no further. On the other hand, if this null hypothesis is rejected, test $H_0 : \pi_y = 0$ using the bounds procedure based on the t-statistic. if $H_0 : \pi_y = 0$, is false, a large value of t_{π_y} should confirm the existence of a level relationship between y_t and X_t.

This approach regards X_t as long run forcing variables for y_t, so there is no feedback from the level of y_t in Eq (5.9). Given this assumption, it is presumed that the explanatory variables are not cointegrated among themselves and that, therefore, the size of the cointegrating space is restricted to unity (De Vita and Abbott, 2004). This assumption can be tested by first estimating an ARDL-ECM equivalent to (5.9) for each of the independent variables, and then testing for an absence of feedback via the statistical significance of the y_t estimated coefficient.

To this end, a variant of the bounds test suggested originally by Banerjee et al (1998) is used. This test is grounded on the t-statistic for $H_0 : \pi_y = 0$, from the estimation of the following equation using OLS:

$$\Delta X_t = c_0 + c_1 t + \pi_x X_{t-1} + \pi_y y_{t-1} + \sum_{i=1}^{p-1} \varphi_i \Delta X_{t-i} + \sum_{j=1}^{q-1} \delta_j' \Delta y_{t-j} + \gamma' \Delta y_t + u_t \quad (5.10)$$

If the null hypothesis cannot be rejected, then X_t is confirmed to be long run forcing regressors for y_t (De Vita and Abbott, 2004).

In the case of rejecting the null hypothesis, namely that there exists a conditional level relationship between y_t and X_t, in the second stage the long run relationship model derived from estimation (5.9), given that an optimal lag structure was used, is defined by:

$$y_t = \theta_0 + \theta_1 t + \theta X_t + v_t \tag{5.11}$$

where $\theta_0 = -c_0 / \pi_y$, $\theta_1 = -c_1 / \pi_y$ and $\theta = -\pi_x / \pi_y$.

Pesaran et al. (2001) referred to the importance of keeping the coefficients of lagged changes unrestricted when testing the hypothesis of no cointegration in Eq (5.9) in order to avoid the subject of these tests to a pre-testing problem. However, they advise the use of more parsimonious specifications when estimating level effects and short run dynamics. To this end, they used the ARDL approach for the estimation of the long run parameters. The equivalent ARDL specification of equation (5.9) can be written as follows:

$$y_t = c_0 + c_1 t + \sum_{i=1}^{p} \pi_{yi} y_{t-i} + \sum_{j=0}^{q} \pi_{xj} X_{t-j} + u_t \tag{5.12}$$

The long run coefficients in equation (5.11) can be obtained from estimating Eq (5.12) given that: $\pi_y = \sum_{i=1}^{p} \pi_{yi} - 1$; $\pi_x = \sum_{j=1}^{q} \pi_{xj}$

The associated short run parameters can then be obtained by estimating the conditional ECM using the same lag structure as used for the ARDL model above.

Bardsen (1989) reported that estimating Eqs (5.9) and (5.12) will give identical results. However, he reported that the ARDL-ECM equation (5.9) explicitly gives the short run dynamics in the differenced terms and the long run coefficients can easily be obtained by dividing the regressors' coefficients by the coefficient of the lagged regressand multiplied by negative sign. Bardsen indicated that Eq (5.9) has the merit of simplification in the computation of the variances of the coefficients regardless of the number of lags involved unlike the formula of (5.12) (Bardsen, 1989). Moreover, it is essential for equation (5.12) to involve continuous lags for the calculation of the long run parameters, whereas these parameters can be computed from Eq (5.9) whether the lags are continuous or not. Therefore, to avoid the problem of over parameterization and to save the degrees of freedom, it is more convenient to use discontinuous lags. Consequently, it seems that formula (5.9) is more appropriate in estimating level effects. For these reasons we intend to use the ARDL-ECM formula with our data sets.

Although the bounds testing procedure of Pesaran et al. (2001) has some merits over the traditionally used methods, it still suffers from some shortcomings. The bounds testing approach is not applicable when there are $I(2)$ variable(s) in a relationship. Moreover, it assumes the existence of only one cointegrating vector, if any, and excludes the possibility that more than one cointegrating vector exists. However, although the bounds testing approach requires the testing of whether or not the variables are $I(2)$ or less, which seems to be a pre-testing problem, we believe that this problem is not severe for two reasons. Firstly, since we are using variables in terms of variance they are unlikely to be higher than $I(1)$. Secondly, unit root tests may provide more reliable results concerning whether a variable is $I(2)$ or less. Therefore, we think these would give more credit for applying the bounds test. The second problem, i.e.

the assumption of unique cointegrating rank, can be tested by regressing equation (5.10) to each regressor as shown above. Thus, the bounds testing approach is the most appropriate method, in our view, for investigating the presence of equilibrium relationships.

5.4 Unit root tests

With regard to the application of the Johansen multivariate method, it is a prerequisite to make sure that all time series included are $I(1)$. Also, regarding the application of the bounds testing approach, which is applicable irrespective whether the variables are $I(0)$, $I(1)$ or mutually cointegrated, it is still necessary to ensure that no variable is integrated of an order higher than one. Therefore, it is vital procedure to investigate the properties of the underlying time series before using them in regressions and obtaining any meaningful inference. That is to examine the order of integration of each variable in our sample.

The literature has proposed several tests to investigate the stationarity of a series. These include, among others, the augmented Dicky Fuller (ADF), Phillips and Perron (PP), Elliott, Rothenberg and Stock (DFGLS) tests. Schwert (1987) has noted that the Phillips-Perron test statistic may reject the null hypothesis of unit root too often in the presence of the first order moving average process (Schwert, 1987). Moreover, simulation studies implemented by Dickey et al. (1986) and DeJong et al. (1992) have shown that the Dickey-Fuller class tests have low power in finite samples. In addition, Schwert (1989) and Perron and Ng (1996) have stated that the majority of unit root tests suffer from severe size distortions when the moving average polynomial of the first differenced series has a large negative root. While few economic variables have been shown to exhibit a negative serial correlation of the autoregressive type, many are found to have large negative moving average roots which lead to over-rejections of the unit root hypothesis.

One of the most important requirements when conducting unit root tests is the selection of the truncation lag. This is required for running the regression of the tests and also for constructing an autoregressive estimate of the spectral density at frequency zero. Simulation experiments, however, have shown a strong relationship between the selection of the lag length and the severity of size distortions and the extent of power loss (Lopez, 1997).

Perron and Ng (1996) analyzed a class of modified unit root tests which were initially proposed by Stock (1990) and showed that these modified tests are far more robust to size distortion than other unit root tests when the residuals have negative serial correlation. Exploiting the local GLS detrending of Elliott, Rothenberg and Stock (1996) which gives substantial power gains, Ng and Perron (2001) applied the idea of GLS detrending to the previously mentioned modified tests and showed that non-negligible size and power gains can be made when used along with an autoregressive spectral density estimator at frequency zero. Secondly, regarding optimal lag length, Ng and Perron (2001) suggested a class of modified information criteria that take better account of the cost of under fitting. They found the modified Akaike information criterion (MAIC) to lead to substantial size improvements over the standard criteria. Joining the two procedures, namely, the GLS detrending approach and the selection rule of lag length, produced methods which allow for setting unit root tests that have much improved size and power according to Ng and Perron (2001). Moreover, using the GLS detrending time series as estimator to the spectral density at frequency zero is shown to have preferred size and power implications. Thus, we intend to apply the Ng and Perron (NP) test in our analysis along with choosing the lag order as referred to by MAIC.

5.4.1 The Ng-Perron unit root test description

This section gives brief description of the unit root testing procedure developed by Ng and Perron (2001), who assumed that data series $\{y_t\}_{t=0}^{T}$ are generated by

$$y_t = d_t + u_t \qquad u_t = \alpha u_{t-1} + v_t \tag{5.13}$$

where $E(u_0^2) < \infty, v_t = \delta(L)e_t = \sum_{j=0}^{\infty} \delta_j e_{t-j}$ with $\sum_{j=0}^{\infty} j|\delta_j| < \infty$ and $\{e_t\} \sim iid(0, \sigma_e^2)$. The non-normal spectral density of v_t at frequency zero is given by $\sigma^2 = \sigma_e^2 \delta(1)^2$. Also, $d_t = \psi' z_t$ where z_t is a set of deterministic components. Ng and Perron considered $d_t = \sum_{i=0}^{P} \psi_i t^i$, with special focus on P=0, 1. They tested the null hypothesis of $\alpha = 1$ against $\alpha < 1$. The augmented Dickey Fuller (ADF) test is the t statistic for $\beta_0 = 0$ in the following autoregression:

$$\Delta y_t = d_t + \beta_0 y_{t-1} + \sum_{j=1}^{k} \beta_j \Delta y_{t-j} + e_{tk} \tag{5.14}$$

Perron and Ng (1996) analyzed the properties of three tests: MZ_α, MZ_t and *MSB*, collectively called as *M* tests. These tests are defined as follows:

$$MZ_\alpha = (T^{-1} y_T^2 - s_{AR}^2 \left(2T^{-2} \sum_{t=1}^{T} y_{t-1}^2 \right)^{-1}, \tag{5.15}$$

$$MSB = \left(T^{-2} \sum_{t=1}^{T} y_{t-1}^2 / s_{AR}^2 \right)^{-1/2}, \tag{5.16}$$

and $MZ_t = MZ_\alpha \times MSB$. An autoregressive estimate of the spectral density of v_t at zero frequency, s_{AR}^2, is given by

$$s_{AR}^2 = \hat{\sigma}_k^2 / (1 - \hat{\beta}(1))^2 \tag{5.17}$$

where $\hat{\beta}(1)=\sum_{i=1}^{k}\hat{\beta}_i$ and $\hat{\sigma}_k^2=(T-k)^{-1}\sum_{t=k+1}^{T}\hat{e}_{tk}^2$. ^ denotes OLS estimates from equation (5.14).

The MZ_α and MZ_t can be viewed as modified versions of the Phillips (1987) and Phillips-Perron (1988) Z_α and Z_t tests, referred to as the Z tests. The MSB test statistic is related to Bhargava's (1986) R_1 statistic. The Z tests suffer from severe size distortion when v_t has a negative moving average root. On the other hand, the M tests have been shown to have much smaller size distortion than most unit root tests, including the Z tests, if a suitable k is chosen.

Ng and Perron (2001) adapted the local to unity GLS detrending procedure proposed previously in Elliott et al. (1996). For any series $\{x_t\}_{t=0}^{T}$, define

$(x_0^{\overline{\alpha}},x_t^{\overline{\alpha}})\equiv(x_0,(1-\overline{\alpha}L)x_t)$, t=1,....,T for some chosen $\overline{\alpha}=1+\overline{c}/T$. The GLS detrended series is defined as

$$\widetilde{y}_t\equiv y_t-\hat{\psi}'z_t \tag{5.18}$$

where $\hat{\psi}$ minimizes $s(\overline{\alpha},\psi)=(y^{\overline{\alpha}}-\psi'z^{\overline{\alpha}})'(y^{\overline{\alpha}}-\psi'z^{\overline{\alpha}})$. If v_t is i.i.d. normal, the point optimal test of the null hypothesis $\alpha=1$ against the alternative $\alpha=\overline{\alpha}$ is the likelihood ratio statistic, $L=S(\overline{\alpha})-S(1)$, where $S(\overline{\alpha})=\min_\psi S(\overline{\alpha},\psi)$. Elliott et al. (1996) considered a feasible point optimal test that takes into consideration that v_t may be serially correlated. They proposed the statistic:

$$P_T=[S(\overline{\alpha})-\overline{\alpha}S(1)]/s_{AR}^2 \tag{5.19}$$

The value $\overline{c}$=-7.0 for P=0 and -13.5 for P=1. They also suggest the DF^{GLS} statistic as the t-statistic for $\beta_0=0$ from the following regression estimated by OLS:

$$\Delta\widetilde{y}_t=\beta_0\widetilde{y}_{t-1}+\sum_{j=1}^{k}\beta_j\Delta\widetilde{y}_{t-j}+e_{tk} \tag{5.20}$$

Ng and Perron (2001) used the GLS detrending approach to the M tests as well, which are referred to as the M^{GLS} tests. Ng and Perron examined the asymptotic properties of the M^{GLS} tests and calculated their critical values. They, also, considered two modified feasible point optimal tests which are given as follows:

$$P=0: \qquad MP_T^{GLS} = \left[\bar{c}^2 T^{-2} \sum_{t=1}^{T} \tilde{y}_{t-1}^2 - \bar{c} T^{-1} \tilde{y}_T^2 \right] / s_{AR}^2$$

$$P=1 \qquad MP_T^{GLS} = \left[\bar{c}^2 T^{-2} \sum_{t=1}^{T} \tilde{y}_{t-1}^2 + (1-\bar{c}) T^{-1} \tilde{y}_T^2 \right] \Big/ s_{AR}^2 \qquad (5.21)$$

When s_{AR}^2 is estimated based on equation (5.20), the test statistics will be called $\overline{M}^{GLS}$ and $\overline{Z}^{GLS}$ respectively. The M^{GLS} and Z^{GLS} test statistics use OLS detrended data to estimate s_{AR}^2 from equation (5.14), on the other hand.

The standard truncation lag selection criteria, such as the AIC and BIC, differ in the weight applied to overfitting, but all use k as the penalty to overfitting, thus, Ng and Perron argued that with integrated data this penalty may be a poor approximation to the cost of underfitting. The sequential t test, like the BIC, leads to less efficient estimates and, hence, to power losses. Ng and Perron, therefore, proposed modified criteria that set $k_{mic} = \arg\min_k MIC(k)$,

$$\text{Where } MIC(k) = \ln(\hat{\sigma}_k^2 + \frac{C_T(\tau_T(k)+k)}{T-k_{\max}} \qquad (5.22)$$

With $C_T > 0$ and $C_T / T \to 0$ as $T \to \infty$. $k_{\max}$ is an upper bound of k. This latter uses $k_{\max} = \text{int}\left(12(T/1000)^{1/4}\right)$, even though other choices are possible. Also, $\tau_T(k) = (\hat{\sigma}_k^2)^{-1} \hat{\beta}_0^2 \sum_{t=k_{\max}+1}^{T} \tilde{y}_{t-1}^2$, with $\hat{\sigma}_k^2 = (T-k_{\max})^{-1} \sum_{t=k_{\max}+1}^{T} \hat{e}_{tk}^2$. For the modified AIC (MAIC), C_T =2 and for modified BIC (MBIC), $C_T = \ln(T-k_{\max})$, respectively. Against this background we intend to use the Ng-Perron method with the modified Akaike criterion to search for the integration order of the data series in our study.

5.4.2 The ADF unit root test description

In addition to the Ng-Perron test, we will use the most commonly applied test, namely, the augmented Dickey-Fuller test in order to compare the results obtained using both methods to make sure that they are consistent. This test can be asymptotically reliable, which is the case for most our data sets. The augmented Dickey-Fuller (ADF) tests were computed by using the following regression equation:

$$\Delta Y_t = \beta_1 + \beta_2 t + \delta Y_{t-1} + \sum_{i=1}^{p} \alpha_i \Delta Y_{t-i} + \varepsilon_t \qquad (5.23)$$

Equation (5.23) tests the null hypothesis of a unit root ($H_0 : \delta = 0$) against the alternative of stationarity ($H_0 : \delta \langle 0$). β_1 and t are included to allow for the presence of significant drift and/or trend components. If Y_t follows an AR(p) process, then a number of lagged dependent variables need to be included to ensure ε_t is white noise error. The test is completed by deriving an OLS estimate of δ and comparing the calculated t statistic with the critical values (see, for instance, Dickey and Fuller, 1979; Abbott, 1999; and Gujarati, 2003).

Having introduced the unit root tests to be used in our study, the next chapter presents the empirical results of these tests.

Chapter six

Results of Unit Root Tests and the Estimated ARCH-based Volatility Measures

6.1 Introduction

The previous chapter introduced the econometric methodologies that will be used in the empirical research. The unit root tests which will be used to determine the order of integration of the time series involved in our study were also introduced. This chapter presents the results of unit root tests for all variables used in the study, using different data frequencies and volatility proxies. More precisely, the simple measure for monthly and quarterly data is used since we can obtain monthly volatility series from monthly series; the standard deviation for quarterly observations since we can only obtain quarterly volatility series from monthly series; and ARCH measures for monthly and quarterly data as well as the standard deviation for annual data. Before proceeding with the cointegration analysis, it is necessary to establish the order of integration of each variable in the equations. This step is necessary even for the bounds testing approach which allows us to test for the presence of long run relationship with a set of variables which are either $I(0)$ or $I(1)$, since it is still necessary to test for higher orders of integration than $I(1)$. To achieve this aim, several unit root tests have been proposed in the literature. For reasons given in the previous chapter, the N-P and ADF unit root tests are applied here to find out the order of integration of our data series. However, in using the N-P test, was encountered a difficult problem which precluded us from determining the integration order of the data sets. This problem is described below. Therefore, we decided to rely only on the results of the ADF test and some other tools. The results of the ADF tests are shown in section 6.2. The ADF tests were applied to all volatility series which were generated using the simple measure and the standard deviation for both monthly and quarterly observations. The ADF tests were also used to determine the integration order for the nominal exchange rates and the differenced variables (the differentials), which is a primary step in estimating volatility proxies based on the ARCH models. The estimation procedure for the ARCH-based variability measures, alongside the ADF results for monthly and quarterly data are presented in section 6.3. Section 6.4 shows the results of ADF tests for the annual data set which is incorporated in the Devereux-Lane model.

However, before introducing the results of unit root tests, and for convenience the symbols used to describe the relevant variables involved in the traditional models-based volatility equations are defined first. These series are described as follows:

$\tilde{c}$	consumption differential
$\tilde{\pi}$	inflation rate differential
$\tilde{m}$	money supply differential
$\tilde{r}$	interest rate differential
s	nominal exchange rate

$\tilde{y}$ income differential

SM volatility of a variable using the simple measure; so SM$\tilde{m}$, for example, refers to the volatility of the money supply differential using the simple measure.

SD volatility of a variable using the standard deviation; so SD$\tilde{m}$, for example indicates the volatility of the money supply differential using the standard deviation.

As emphasized in the past chapter, we intended to use both the N-P and ADF unit root tests. However, the application of the N-P tests with our data set gave results which made it difficult to determine the order of integration of the variables. This ambiguity occurred in many series, making it difficult to specify whether a series was $I(0)$, $I(1)$ or more. The problem encountered can be described as follows: if a variable was found to be nonstationary (or stationary at the 5% level of significance for example) in level form, its first difference is still nonstationary (or becomes nonstationary). Furthermore, the significance level of the coefficient of the lagged level variable in the test equation gets less, when we take the first difference. The problem still exists even after taking the second difference. To illustrate the problem more precisely, some examples of the results we obtained from the application of the N-P tests to the data under consideration are given in Table 6-1.

As can be seen from Table 6-1, the variable SM$\tilde{c}$ in Algeria is stationary in level, however, it becomes nonstationary in the first differenced form. The same situation can be noticed for the variables SMs in Algeria, and SM$\tilde{m}$ in Germany.

Table 6-1: Results of Ng-Perron modified unit root tests for monthly data using the simple measure

Country	Variable	MZ_{α}	MZ_{t}	*MSB*	*MPT*	Lags
Algeria	SM$\tilde{c}$	-9.11**	-2.13**	0.23*	2.69**	10
		(-0.29)	(-0.38)	(1.31)	(84.24)	15
Algeria	SM$\tilde{s}$	-23.73***	-3.44***	0.15***	1.03***	9
		(-0.16)	(-0.28)	(1.77)	(153.33)	15
Germany	SM$\tilde{m}$	-24.85***	-3.53***	0.14***	0.99***	5
		(0.06)	(0.12)	(2.17)	(241.91)	14
Canada	SM$\tilde{s}$	-5.52	-1.66*	0.30	4.44	10
		(-0.07)	(-0.15)	(2.34)	(270.76)	15

Notes: The sample period starts in 1973m3 and ends in 1998m12. * , ** and *** denote level of significance at 10%, 5% and 1% respectively using the asymptotic critical values of Ng and Perron (2001, table 1). The Ng-Perron tests for the first difference of each variable are shown in parentheses. The number of lags used for the Ng-Perron regressions (shown in last column) are based on the modified AIC. The MZ_{α} and MZ_{t} can be viewed as modified versions of the Phillips (1987) and Phillips-Perron (1988) Z_{α} and Z_{t} tests respectively. The *MSB* test statistic is related to Bhargava's (1986) R_1 statistic. MPT is the modified feasible point optimal test originally proposed by Elliott et al., 1996.

The SM$\tilde{s}$ in Canada is non-stationary either at level or first difference and the magnitude of the coefficient at level (-5.52) is larger than that in the first differenced form (-0.07). These peculiar findings occurred for many variables using different data frequencies and volatility measures.

Different lag length criteria and various spectral estimation methods were used in trying to cure this problem, yet it was still present. After correspondence with S. Ng and P. Perron[7] and oral discussion with M. Karananos, we decided to use the ADF test

[7] Consultations were made via the following email addresses: serena.ng@umich.edu and perron@bu.edu.

to determine the integration order associated with the modified AIC to determine the lag length. Also, for the sake of consistency, we applied the ADF test to all data series, including those where the previous difficulty did not occur. The problem using this test almost completely disappears. To avoid the drawbacks of the ADF test, as discussed in the past chapter, guidance was sought from other criteria in determining the order of integration of a series, in addition to the application of the ADF test. The other criteria applied were the line graphs, the correlogram, the AR(p) process where the sum of AR coefficients should be close to unity if the series is integrated of order one, and other unit root tests. These tests include Phillips and Peron (PP), Elliott, Rothenberg and Stock point-optimal (ERS point-optimal).

6.2 Results of the ADF unit root tests for volatility proxies generated from the SM and SD using monthly and quarterly observations

The results of the ADF unit root tests with our time series data are included in the following tables. Table 6-2 shows the order of integration for all 48 variables considered for all countries using the simple measure of volatility (SM) for monthly data covering the period from March 1973 to December 1998. Since for most variables reveal some outliers which may affect the stationarity tests results, dummy variables were included into the ADF equation at the dates indicated (appearing in column four) to account for the outliers[8] where we think that they may have affected the findings. The position of dummy variables is determined by visual examination of

[8] As after consultations with R. Harris, via email r.harris@socsci.gla.ac.uk, in such circumstances researchers often bootstrap because the introduction of dummies will affect the small sample critical values for the test. However, by referring to the graphs, correlogram and the sum of AR regression coefficients of the variables, it can easily be seen that the series looks stationary when ignoring the outliers. Therefore, it was thought that there was no need for bootstrapping and the series were considered as stationary.

the plots of a series to specify the extreme observations, and by reviewing any important changes which occurred in macroeconomic policies and economic or political events. Break points can also be determined as endogenous by an iterative regression process. However, this would involve an enormous number of regressions given the number of series, frequencies and countries being tested. Thus, it was thought reasonable to merely use a simple, conventional and well-known method. It is rare to include trends in ADF equations as the plots generally do not exhibit such time trend in volatility variables. However, trends were included when the plots showed their existence.

Table 6-2: Results of the ADF unit root tests for monthly data using the simple measure

Country	Variable	ADF Levels (lags)	ADF First differences	Outliers	ADF (levels with dummies)	Decision
Algeria	$SM\tilde{c}$	-2.96**(10)		1990m7	-4.60***	I(0)
	$SM\tilde{\pi}$	-2.73*(12)	-31.15***			I(1)
	$SM\tilde{m}$	-4.02***(11)				I(0)
	$SM\tilde{r}$	17.70***				I(0)
	SMs	-5.06***(7)				I(0)
	$SM\tilde{y}$	-4.46***(11)				I(0)
Canada	$SM\tilde{c}$	-6.61***(2)				I(0)
	$SM\tilde{\pi}$	-4.00***(12)		1991m1	-5.90***	I(0)
	$SM\tilde{m}$	-2.10(11)	-23.03***			I(1)
	$SM\tilde{r}$	-8.00***(2)				I(0)
	SMs	-4.39***(10)				I(0)
	$SM\tilde{y}$	-3.75***(6)				I(0)
Germany	$SM\tilde{c}$	-5.74***(2)				I(0)
	$SM\tilde{\pi}$	-3.37**(12)		1991m1	-4.95***	I(0)
	$SM\tilde{m}$	-5.67***(5)				I(0)
	$SM\tilde{r}$	-3.29**(15)				I(0)
	SMs	-2.81*	-27.99***			I(1)
	$SM\tilde{y}$	-11.55***				I(0)
Japan	$SM\tilde{c}$	-5.76***(2)				I(0)
	$SM\tilde{\pi}$	-6.93***(13)				I(0)
	$SM\tilde{m}$	-1.71(11)	-30.04***			I(1)

Country	Variable	ADF Levels (lags)	ADF First differences	Outliers	ADF (levels with dummies)	Decision
	SM$\tilde{r}$	-4.40***(5)				I(0)
	SMs	-4.74***(4)				I(0)
	SM$\tilde{y}$	-2.74*(15)	-25.43***			I(1)
Kuwait	SM$\tilde{c}$	-3.19**(13)				I(0)
	SM$\tilde{\pi}$	-3.09**(12)				I(0)
	SM$\tilde{m}$	-7.82***(2)				I(0)
	SM$\tilde{r}$	-3.60***(11)				I(0)
	SMs	-2.72*(11)	-44.97***			I(1)
	SM$\tilde{y}$	-3.76***(14)				I(0)
Libya	SM$\tilde{c}$	-5.64***(2)				I(0)
	SM$\tilde{\pi}$	-0.65(5)	-15.44***			I(1)
	SM$\tilde{m}$	-1.82(11)	-18.83***			I(1)
	SM$\tilde{r}$	-11.55***				I(0)
	SMs	-12.52***				I(0)
	SM$\tilde{y}$	-3.99***(11)				I(0)
UK	SM$\tilde{c}$	-2.83*(2)	-27.29***			I(1)
	SM$\tilde{\pi}$	-3.30**(12)				I(0)
	SM$\tilde{m}$	12.12***				I(0)
	SM$\tilde{r}$	-2.72*(11)	-29.80***			I(1)
	SMs	-4.64***(9)				I(0)
	SM$\tilde{y}$	-3.54***(15)				I(0)
Venezuela	SM$\tilde{c}$	-2.62*(12)	-21.51***			I(1)
	SM$\tilde{\pi}$	-7.74***				I(0)
	SM$\tilde{m}$	-2.10(11)	-18.12***			I(1)
	SM$\tilde{r}$	-69.36***				I(0)
	SMs	-10.98***				I(0)
	SM$\tilde{y}$	-2.86*	-14.66***			I(1)

Notes: The sample period starts in 1986m3 and 1989m3 in Libya and Venezuela respectively and in 1973m3 for the other countries, ending in 1998m12 for them all. * , ** and *** denote levels of significance at 10%, 5% and 1% respectively using the Mackinnon (1996) one-sided p-values. The number of lags used for the ADF regressions (shown in parentheses) are based on the modified AIC. When the number of lags is not reported, it is in fact zero. Note that the decision with respect to the order of integration is not exclusively based on the results of the ADF test.

From the results it can be seen that most variables are integrated of order zero except in thirteen cases. Such mixed results in terms of integration order imply that for monthly data, and when using the simple measure, the bounds testing method is the only suitable approach for examining cointegrating relationships. The Johansen method cannot be used in this case where there are mixed orders of integration.

Table 6-3 shows the results of the ADF test on all variables where the simple volatility proxy is used to approximate the volatility of variables for quarterly data. The results in tables 6-2 and 6-3 are quite similar, possibly because of the use of the same volatility measure (SM) for different frequencies.

Most variables in Table 6-3 can be considered as stationary; however, there are also about 15 non-stationary variables. These series are found to be $I(1)$. These findings imply that the Johansen method cannot be applied as there are $I(0)$ and $I(1)$ variables in each case.

Table 6-3: Results of the ADF unit root tests for quarterly data and using the simple measure

Country	Variable	ADF (levels)	ADF (first differences)	Outliers in	ADF (levels with dummies)	Decision
Algeria	SM$\tilde{c}$	-2.68*(3)	-14.48***	1988q4, 1989q4, 1990q3	-12.54***	I(0)
	SM$\tilde{\pi}$	-2.49*(1)	-17.10***			I(1)
	SM$\tilde{m}$	-1.74(11)	-17.35***			I(1)
	SM$\tilde{r}$	-10.62***				I(0)
	SMs	-1.47(12)	-16.33***	1991q1, 1994q2	-8.26***	I(0)
	SM$\tilde{y}$	-2.68*(3)	-20.72***			I(1)
Canada	SM$\tilde{c}$	-3.32**(7)				I(0)
	SM$\tilde{\pi}$	-2.31(8)	-16.76***	1980q1, 1982q2, 1991q1	-3.55***	I(0)
	SM$\tilde{m}$	-2.02	-13.41***			I(1)
	SM$\tilde{r}$	-1.93(8)	-16.58***			I(1)
	SMs	-4.43***(3)				I(0)
	SM$\tilde{y}$	-2.18(10)	-16.96***	1985q1	-6.76***	I(0)
Germany	SM$\tilde{c}$	-2.74*(7)	-17.24***	1991q1	-19.51***	I(0)
	SM$\tilde{\pi}$	-2.86*(4)	-16.15***			I(1)
	SM$\tilde{m}$	-9.55***				I(0)
	SM$\tilde{r}$	-2.54(6)	-18.98***			I(1)

Country	Variable	ADF (levels)	ADF (first differences)	Outliers in	ADF (levels with dummies)	Decision
	SMs	-3.71***(4)				I(0)
	SM$\tilde{y}$	-4.14***(3)				I(0)
Japan	SM$\tilde{c}$	-6.07***(1)				I(0)
	SM$\tilde{\pi}$	-3.53***(9)				I(0)
	SM$\tilde{m}$	-2.60*(1)	-19.02***			I(1)
	SM$\tilde{r}$	-4.62***(1)				I(0)
	SMs	-6.53***(1)				I(0)
	SM$\tilde{y}$	-3.27**(6)				I(0)
Kuwait	SM$\tilde{c}$	-3.53***(3)				I(0)
	SM$\tilde{\pi}$	-2.80*(4)	-17.32***			I(1)
	SM$\tilde{m}$	-4.30***(2)				I(0)
	SM$\tilde{r}$	-2.11(8)	-19.73***	1980q4	-6.24***	I(0)
	SMs	-2.31*(4)	-17.01***			I(1)
	SM$\tilde{y}$	-5.26***				I(0)
Libya	SM$\tilde{c}$	-1.40(7)	-11.29***	1987q1	-3.56**	I(1)
	SM$\tilde{\pi}$	-2.75*(2)	-7.70***			I(1)
	SM$\tilde{m}$	-1.54(8)	-9.82***	1986q4	-9.09***	I(0)
	SM$\tilde{r}$	-2.85*(1)	-8.87***			I(1)
	SMs	-6.72***				I(0)
	SM$\tilde{y}$	-6.79***(6)				I(0)
UK	SM$\tilde{c}$	-2.29(8)	-14.52***	1979q2	-2.57*	I(1)
	SM$\tilde{\pi}$	-3.96***(3)				I(0)
	SM$\tilde{m}$	-9.84***				I(0)
	SM$\tilde{r}$	-1.60(12)	-19.72***	1980q4	-2.88*	I(1)
	SMs	-2.95**(6)		1981q2, 1987q4, 1992q4	-5.22***	I(0)
	SM$\tilde{y}$	-2.80*(11)	-15.92***		-4.33***	I(0)
Venezuela	SM$\tilde{c}$	-6.65***				I(0)
	SM$\tilde{\pi}$	-5.05***				I(0)
	SM$\tilde{m}$	-3.16(9)	-10.22***			I(1)
	SM$\tilde{r}$	-7.94***				I(0)
	SMs	-8.27***				I(0)
	SM$\tilde{y}$	-1.54	-8.73***			I(1)

Notes: The sample period starts in 1986q2 and 1989q2 in Libya and Venezuela respectively and in 1973q2 for the rest of the countries and it ends in 1998q4 for them all. * , ** and *** denote levels of significance at 10%, 5% and 1% respectively using the Mackinnon (1996) one-sided p-values. The number of lags (shown in parentheses) used for the ADF regressions are based on the modified AIC. When the number of lags is not reported, it is in fact zero. Note that our decision with respect to the order of integration is not exclusively based on the results of the ADF test.

Thus, the bounds testing approach must be used in order to investigate the presence of cointegrating relationships between the regressors and the dependent variable.

Table 6-4 contains the results of the ADF tests of the 48 variables in all countries using the standard deviation as a measure of volatility for quarterly data. Note that some variables again turned out to be stationary after taking into account the effects of the outliers. Note also that there are different numbers of variables were found to be $I(1)$ when using the SD: 23 $I(1)$ variables using the SD as opposed to 15 using the SM.

The results in Table 6-4 imply that the Johansen method of regression cannot be applied into investigating the assumed relationships, in that the target equation contains variables that have different orders of integration, whereas this method requires all variables in an equation to be integrated of the same order.

Table 6-4: Results of the ADF unit root tests for quarterly data and using the standard deviation

Country	Variable	ADF (levels)	ADF (first differences)	Outliers	ADF (levels with dummies)	Decision
Algeria	SD$\tilde{c}$	-1.91(8)	-15.43***			I(1)
	SD$\tilde{\pi}$	-1.90(9)	-18.33***			I(1)
	SD$\tilde{m}$	-3.92***(5)				I(0)
	SD$\tilde{r}$	-10.44***				I(0)
	SDs	-1.20(12)	-18.26***	1991q1,q3, 1994q2	-6.93***	I(0)
	SD$\tilde{y}$	-4.65***(3)				I(0)
Canada	SD$\tilde{c}$	-5.75***(1)				I(0)
	SD$\tilde{\pi}$	-1.85(11)	-14.85***			I(1)
	SD$\tilde{m}$	-1.90(3)	-12.77***			I(1)
	SD$\tilde{r}$	-4.40***(2)				I(0)
	SDs	-9.62***				I(0)
	SD$\tilde{y}$	-1.72(7)	-20.70***			I(1)
Germany	SD$\tilde{c}$	-1.45(8)	-14.18***	1993q3, 1998q3	-2.90**	I(0)
	SD$\tilde{\pi}$	-2.04(7)	-21.97***	1991q1	-2.73*	I(1)
	SD$\tilde{m}$	-5.48***(1)				I(0)

Country	Variable	ADF (levels)	ADF (first differences)	Outliers	ADF (levels with dummies)	Decision
	SD$\tilde{r}$	-3.12**(4)				I(0)
	SDs	-4.51***(4)				I(0)
	SD$\tilde{y}$	-2.45(9)	-15.31***			I(1)
Japan	SD$\tilde{c}$	-4.16***(3)				I(0)
	SD$\tilde{\pi}$	-2.56(7)	-16.47***			I(1)
	SD$\tilde{m}$	-1.77(3)	-13.68***			I(1)
	SD$\tilde{r}$	-1.27(12)	-15.57***	1984q1, 1998q4	-2.22	I(1)
	SDs	-3.16**(3)		1978q4, 1998q4	-4.49***	I(0)
	SD$\tilde{y}$	-2.27(7)	-17.86***		-4.29***	I(0)
Kuwait	SD$\tilde{c}$	-1.57(7)	-18.27***			I(1)
	SD$\tilde{\pi}$	-0.95(11)	-20.64***			I(1)
	SD$\tilde{m}$	-7.81***				I(0)
	SD$\tilde{r}$	-2.25(8)	-15.62***			I(1)
	SDs	-2.60(3)	-18.16***			I(1)
	SD$\tilde{y}$	-2.95**(5)		1990q3	-4.85***	I(0)
Libya	SD$\tilde{c}$	-1.61(8)	-9.74***			I(1)
	SD$\tilde{\pi}$	-2.19(5)	-11.25***			I(0)
	SD$\tilde{m}$	-2.56(1)	-9.13***			I(1)
	SD$\tilde{r}$	-1.82	-8.98***			I(1)
	SDs	-7.27***				I(0)
	SD$\tilde{y}$	-1.94(5)	-11.71***	1987q3	-4.50***	I(0)
UK	SD$\tilde{c}$	-5.04***(1)				I(0)
	SD$\tilde{\pi}$	-2.71*(9)	-14.64***			I(1)
	SD$\tilde{m}$	-10.20***				I(0)
	SD$\tilde{r}$	-1.25(10)	-19.18***			I(1)
	SDs	-1.65(12)	-17.65***			I(1)
	SD$\tilde{y}$	-1.30(12)	-16.24***	1979q1	-6.80***	I(0)
Venezuela	SD$\tilde{c}$	-2.87*(1)	-8.38***			I(1)
	SD$\tilde{\pi}$	-10.69***				I(0)
	SD$\tilde{m}$	-1.28(8)	-5.19***(1)			I(1)
	SD$\tilde{r}$	-8.47***(1)				I(0)
	SDs	-6.00***(1)				I(0)
	SD$\tilde{y}$	-1.97(3)	-10.41***			I(1)

Notes: The sample period starts in 1986q2 and 1989q2 in Libya and Venezuela respectively and in 1973q2 for the rest of countries and ends in 1998q4 for them all. * , ** and *** denote levels of significance at 10%, 5% and 1% respectively using the Mackinnon (1996) one-sided p-values. The number of lags (shown in parentheses) used for the ADF regressions are based on the modified AIC. When the number of lags is not reported, it is in fact zero. Note that our decision with respect to the order of integration is not exclusively based on the results of the ADF test.

On the other hand the bounds testing approach appears to be the most appropriate method to examine the presence of cointegrating relationships because none of the variables appeared to be integrated of order higher than one.

From tables 6.2-6.4 we can see that there are 24 out of 48 variables which have identical results as to the order of integration of a series. For example, the Algerian inflation differential is I(1) using all volatility proxies and data frequencies. From these three tables it can be concluded that since we have $I(0)$ and $I(1)$ variables in all countries, the bounds testing method is the only valid approach to be used for testing the existence of long run relationships. The Johansen method, thus, would be invalid for the three conventional models.

6.3 The ARCH-based measures of volatility

As mentioned in chapter three, an ARCH-based variability proxy was used alongside the simple measure and the standard deviation. This section reports empirically estimation of this measure. In fact the discussion of this measure has been postponed to this point because its estimation involves using unit root tests as well as estimating some time series models.

ARCH models revolve around the second moment of the data; in particular, they specify the variance of a series as conditional on past realizations. Thus, they can be used to approximate measures centred on prediction errors. The first step in estimating volatility proxies based on ARCH models is to investigate the order of integration of the nominal exchange rates and the other differential variables using the

ADF test,[9] as the ARCH model is only applied to stationary series. Therefore, the following section presents the results of the unit root tests for the differential variables.

6.3.1 Results of the ADF unit root tests for monthly differential data

Table 6-5 depicts the results of the ADF test for all data series in all our sample countries using monthly data. The trend is included alongside the intercept in the ADF equation when the plot of a series shows the existence of a time trend. From the table below the following can be noted. Interestingly, all inflation rate differentials are *I*(0), except for the Libyan case. The Venezuelan consumption differential is the only stationary variable among its counterparts in the other countries. The rest of the variables appeared to be integrated of order one. According to the ADF test some stationary series were found; however, when consulting the other criteria, they appeared to be *I*(1). These variables include the Japanese consumption differential, the Venezuelan percentage interest rate difference and the exchange rate of the Venezuelan Bolivar against the US dollar. Thus, we assume that these variables are *I*(1). By contrast, some variables which appear *I*(1) based on the ADF test appear to be *I*(0) according to the other criteria. This applies to the Algerian and German inflation rate differentials. Therefore, we assume that they are *I*(0).

Table 6-5: Results of the ADF unit root tests for monthly differential data

Country	Variable	ADF (level)	ADF (first differenced)	Intercept & trend	Decision
Algeria	$\tilde{c}$	-3.76***(12)	-22.89***	C	I(1)
	$\tilde{\pi}$	-2.25(14)	-30.79***	C	I(0)
	$\tilde{m}$	-2.20(1)	-3.06**(15)	C&T	I(1)
	$\tilde{r}$	-1.35(1)	-16.00***	C	I(1)
	s	-1.44(8)	-4.56***(7)	C&T	I(1)

[9] The ADF test was performed because we faced the same problem with the N-P test as mentioned previously.

Country	Variable	ADF (level)	ADF (first differenced)	Intercept & trend	Decision
	$\tilde{y}$	-1.23(13)	-26.89***	C&T	I(1)
Canada	$\tilde{c}$	-1.62(11)	-25.94***	C	I(1)
	$\tilde{\pi}$	-3.05**(11)		C	I(0)
	$\tilde{m}$	-2.54(12)	-2.62***(11)	C&T	I(1)
	$\tilde{r}$	-2.65*(6)	-12.19***(1)	C	I(1)
	s	-2.10(13)	-3.29**(10)	C&T	I(1)
	$\tilde{y}$	-2.83(1)	-4.13***(14)	C&T	I(1)
Germany	$\tilde{c}$	-2.72*(2)	-26.29***	C	I(1)
	$\tilde{\pi}$	-2.19(11)	-27.98***	C	I(0)
	$\tilde{m}$	-0.57(6)	-3.15**	C&T	I(1)
	$\tilde{r}$	-2.16(15)	-20.48***	C	I(1)
	s	-1.98	-3.76***(12)	C&T	I(1)
	$\tilde{y}$	-1.94(1)	-3.59***(15)	C&T	I(1)
Japan	$\tilde{c}$	-3.48**(8)	-28.37***	C&T	I(1)
	$\tilde{\pi}$	-4.06***(11)		C	I(0)
	$\tilde{m}$	-2.33(15)	-2.25**(13)	C	I(1)
	$\tilde{r}$	-2.36(8)	-12.54***	C	I(1)
	s	-1.98	-4.04***(11)	C&T	I(1)
	$\tilde{y}$	-0.72(3)	-3.64***(13)	C	I(1)
Kuwait	$\tilde{c}$	-1.99(10)	-23.22***	C&T	I(1)
	$\tilde{\pi}$	-3.75***		C	I(0)
	$\tilde{m}$	-2.05(4)	-4.78***(7)	C	I(1)
	$\tilde{r}$	-1.90(8)	-14.36***	C&T	I(1)
	s	-2.87(4)	-10.14***(2)	C&T	I(1)
	$\tilde{y}$	-2.64(1)	-4.54***(9)	C&T	I(1)
Libya	$\tilde{c}$	-2.70(2)	-21.37***	C&T	I(1)
	$\tilde{\pi}$	-1.21(6)	-14.25***	C	I(1)
	$\tilde{m}$	-2.75(2)	-3.94***	C&T	I(1)
	$\tilde{r}$	-1.33	-2.59***(8)	C	I(1)
	s	-2.25	-8.73***(1)	C&T	I(1)
	$\tilde{y}$	-1.39(12)	-20.08***	C	I(1)
UK	$\tilde{c}$	-2.60(11)	-27.35***	C&T	I(1)
	$\tilde{\pi}$	-2.10(11)	-26.14***	C	I(0)
	$\tilde{m}$	-2.58*(11)	-29.86***	C	I(1)
	$\tilde{r}$	-2.25(15)	13.60***	C	I(1)
	s	-2.12	-4.42***(10)	C&T	I(1)
	$\tilde{y}$	-2.42(1)	-5.86***(7)	C&T	I(1)
Venezuela	$\tilde{c}$	-3.86***(1)		C	I(0)
	$\tilde{\pi}$	-2.96**(4)		C	I(0)

Country	Variable	ADF (level)	ADF (first differenced)	Intercept & trend	Decision
	$\tilde{m}$	-1.71(5)	-5.58***(11)	C&T	I(1)
	$\tilde{r}$	-3.04**	-6.18***(2)	C	I(1)
	s	-3.86**	-11.06***	C&T	I(1)
	$\tilde{y}$	-2.43(12)	-22.02***	C&T	I(1)

Notes: The sample period starts in 1986m3 and 1989m3 in Libya and Venezuela respectively and in 1973m3 for the rest of countries and ends in 1998m12 for them all. * , ** and *** denote levels of significance at 10%, 5% and 1% respectively using the Mackinnon (1996) one-sided p-values. The number of lags (shown in parentheses) used for the ADF regressions are based on the modified AIC. When the number of lags is not reported, it is in fact zero. C and T denote the inclusion of intercept and trend respectively in the ADF level equation. Note that our decision with respect to the order of integration is not exclusively based on the results of the ADF test.

Since ARCH models are only applied to stationary series, one should first take the first difference of the $I(1)$ series. The second step in estimating an ARCH-based volatility proxy is to examine the presence of an ARCH effect in the series by modelling it for example by a time series model. Therefore, we specify an ARMA(p,q)[10] models for the data using the Box-Jenkins approach for the purpose of investigating the existence of an ARCH effect in the data series. The time series models chosen for this purpose are shown in column (3) in Table 6-6[11]. The fitted models were chosen depending on the following criteria: (1) the lowest value of the Akaike information criterion (AIC) and the Schwarz Bayesian information criterion (SBC); (2) the significance of the coefficients as indicated by t-values; (3) the sum of AR coefficients not being close to unity and the MA roots lying outside the unit circle, indicating that is the model is stationary and invertible respectively; and (4) the absence of autocorrelations as indicated by the Ljung-Box Q-statistic.

[10] In essence, an estimated equation of ARMA (2,3), for instance, reveals the inclusion of the first two autoregressive terms and the first three moving average terms. However, the notation ARMA[(2),(1,4)] is used here to indicate that only the autoregressive term at lag 2 and the moving average terms at lags 1 and 4 are included in the model. This can also be found in Enders, W. (1995) *Applied econometric time series*. Series in Probability and Mathematical Statistics. New York; Chichester, U.K. and Toronto: Wiley.

[11] The diagnostic tests indicate the absence of autocorrelation; however, the Jarque-Bera test rejects the normal distribution assumption of the residuals in most cases. Harris and Sollis (2003) indicated that non-normality is an inherent feature of the errors from regression models for financial data. Moreover, Karananos suggested that this can also be true for other macroeconomic series.

The Lagrange multiplier (LM) test proposed by Engle (1982), which tests for both ARCH and GARCH effects was used. The LM test is performed by estimating the following equation by OLS:

$$\hat{u}_t^2 = \alpha_0 + \alpha_1 \hat{u}_{t-1}^2 + \alpha_2 \hat{u}_{t-2}^2 + \dots\dots\dots + \alpha_p \hat{u}_{t-p}^2 + v_t$$

where $\hat{u}$ is the fitted residuals from the original regression. Engle (1982) recommended testing the null hypothesis $H_0 : \alpha_1 = \alpha_2 = \dots\dots = \alpha_P = 0$ using the LM principle. This test statistic is calculated as the sample size multiplied by the R^2 for the regression involving the fitted OLS residuals. Under the null hypothesis, this statistic is distributed as chi-squared (χ^2) with p degrees of freedom. The F version of the LM test for the ARCH (not provided here to save space) gives almost the same findings as the LM statistic shown in column (4) of Table 6-6. The ARMA models are first fitted to the exchange rates data, from which the presence of an ARCH effect is tested. We start with the exchange rates data series, because there is no point in examining the ARCH effect in other series, if it is not present in the exchange rate series (the dependent variable). The ARCH effect tests of the exchange rate revealed that it is present only in the exchange rates of Algeria, Japan, Kuwait and the UK at lags shown in column (5) of table 6-6. The rest of the exchange rates data series seem to be free from ARCH effects as indicated by the LM test statistic. Therefore, one does not need to study the presence of ARCH effect in the other differential time series data for countries other than Algeria, Japan, Kuwait and the UK. For these four economies the ARCH effect test results for the other data series are shown in column (5) as well. The findings indicate the existence of ARCH effect in: all series in Algeria except the inflation rate differential; only two series in Japan, the consumption differential and the interest rate differential; all data series in Kuwait except for the Kuwaiti income differential; and all of the data series for the UK. Before proceeding to fit an ARCH type model to our data series, it is interesting to compare the results of the ARCH effect tests of the exchange rate series with those of previous empirical studies. Most of these studies differ from ours in terms of the period of study, the frequency of the data and the sample considered. For example, Lastrapes (1989) studied the existence of ARCH effect in the exchange rates of the British pound, Canadian dollar, Deutschmark, Japanese yen and Swiss franc against the US dollar using weekly data for the period 1976-1986. Pozo (1992) used real monthly exchange rate series between the British pound and the US dollar for the period 1900-1940 (Pozo, 1992b). Frommel and Menkhoff (2003) used daily data on the exchange rate of the US dollar, Japanese yen, pound sterling, Swiss franc, French franc and the Canadian dollar versus the Deutschmark for the period 1973-1998 (Frommel and Menkhoff, 2003). The closest study to ours may be that performed by McKenzie and Brooks (1997). They used monthly data of the US-German exchange rate for the period 1973-1992 and found that the ARCH (1,1) model is adequate for the data. However, in our study we did not find an ARCH effect in the exchange rate series between Germany and the US. This may be due to the usage of different span, where we studied a longer period (1973-1998) than theirs.

Table 6-6: Results of the LM test for the ARCH effect using monthly data

Country	Variable	ARIMA(p,d,q)	LM statistic	ARCH effect at lag
Algeria	$\tilde{c}$	ARIMA[(10,12),1,(1,2,7)]	70.96(0.00)[7]	2,7
	$\tilde{\pi}$	ARIMA[(3,4,13),0,0] SAR(12)	7.92(0.24)[6]	NO
	$\tilde{m}$	ARIMA[0,1,(1)]	10.32(0.00)[1]	1
	$\tilde{r}$	ARIMA[0,1,(12)]	22.58(0.03)[12]	12
	s	ARIMA[(1,2,8),1,(5,31)]	39.11(0.00)[8]	8
	$\tilde{y}$	ARIMA[(10,13,14,24,29),1,(1,5,26,34,36)] SAR(12)	26.54(0.00)[8]	3,8
Canada	$\tilde{c}$			

Country	Variable	ARIMA(p,d,q)	LM statistic	ARCH effect at lag
	$\tilde{\pi}$			
	$\tilde{m}$			
	$\tilde{r}$			
	s	ARIMA[0,1,(11,13)]	3.59(0.06)[1]	No
	$\tilde{y}$			
Germany	$\tilde{c}$			
	$\tilde{\pi}$			
	$\tilde{m}$			
	$\tilde{r}$			
	s	ARIMA[(17),1,0]	7.80(0.65)[10]	No
	$\tilde{y}$			
Japan	$\tilde{c}$	ARIMA[(1,2,3,4,5),1,(10,33)] SMA(12)	21.94(0.00)[3]	3
	$\tilde{\pi}$	ARIMA[(5,7,24),0,0] SAR(12)	10.14(0.75)[14]	No
	$\tilde{m}$	ARIMA[(1,2,6,9,14),1,(12,30)] SAR(12)	5.10(0.16)[3]	No
	$\tilde{r}$	ARIMA[(1,7,15),1,(9,12,20)]	74.34(0.00)[6]	1,6
	s	ARIMA[0,1,(17,21)]	15.43(0.01)[5]	5
	$\tilde{y}$	ARIMA[(1,17,22),1,0]	13.39(0.00)[3]	3
Kuwait	$\tilde{c}$	ARIMA[0,1,(1,2,12)]	6.73(0.01)[1]	1
	$\tilde{\pi}$	ARIMA[(1,12),0,0]	29.19(0.00)[12]	12
	$\tilde{m}$	ARIMA[(2),1,0]	39.52(0.00)[1]	1
	$\tilde{r}$	ARIMA[(1,2,6,8,20),1,(1,14)]	70.93(0.00)[12]	1,2,12
	s	ARIMA[(1,4,17),1,0]	19.55(0.01)[8]	8
	$\tilde{y}$	ARIMA[0,1,(2,11,14)]	0.00(0.94)[1]	No
Libya	$\tilde{c}$			
	$\tilde{\pi}$			
	$\tilde{m}$			
	$\tilde{r}$			
	s	ARIMA[0,1,(4)]	0.01(0.91)[1]	No
	$\tilde{y}$			
UK	$\tilde{c}$	ARIMA[(1,2),1,(3,6,12,21)]	20.59(0.00)[6]	6
	$\tilde{\pi}$	ARIMA[(1,20),0,0] SAR(12)	22.94(0.00)[9]	1, 9
	$\tilde{m}$	ARIMA[(3,6,7,9,11,16),1,(12)] SAR(12)	15.69(0.00)[1]	1
	$\tilde{r}$	ARIMA[(7,14,15),1,(1,19)]	71.06(0.00)[6]	1,2,4,6
	s	ARIMA[0,1,(14)]	15.38(0.00)[4]	1,4
	$\tilde{y}$	ARIMA[(1,24),1,(20)]	15.12(0.00)[1]	1
Venezuela	$\tilde{c}$			
	$\tilde{\pi}$			
	$\tilde{m}$			

Country	Variable	ARIMA(p,d,q)	LM statistic	ARCH effect at lag
	$\tilde{r}$			
	s	ARIMA[(17),1,0]	2.32(0.68)[4]	N0
	$\tilde{y}$			

Notes: numbers in squared brackets in the fourth column refer to the lag at which the LM ARCH test was performed. SAR and SMA refer to seasonal autoregressive and moving average terms.

The third step in estimating variability from the ARCH models is to specify the best fitted ARCH model for the data found to contain an ARCH effect.

Given that these tests indicate the possibility of the presence of ARCH effects in the indicated cases, it is appropriate to fit ARCH models to the data. Several types of ARCH models were estimated, including ARCH, GARCH and EGARCH, with different lags. It is likely that several ARCH models will have significant coefficients and thus will be found to be reliable for the data. Since economic theory gives little guidance for the selection of an optimal model amongst those fitted to the data, we follow McKenzie (1997) in adopting the best fitted model. McKenzie has proposed a three-step procedure for the selection of the optimal model from potential candidates. Firstly, non-converting models are excluded in addition to the models with singularity problems. Secondly, models that have statistically insignificant parameters are excluded as well as models whose ARCH and GARCH coefficients sum up to one or more, in which case the model is explosive whereby a disturbance to a market does not disappear over time and therefore must be excluded. Thirdly, as the AIC and SBC are constructed around the goodness of fit within the first moment, they may not be suitable for the ARCH models which revolve around the second moment of the data (McKenzie, 1997). McKenzie adopted the method introduced by Pagan and Schwert (1990). Pagan and Schwert proposed an alternative model selection criterion. They suggested regressing $\hat{u}_t^2 = a_0 + a_1\sigma_t^2 + v_t$ for each reliable ARCH model, where $\hat{u}_t$

are the residuals and σ_t^2 the fitted values of the conditional variance. Then, the model with the highest explanatory power (R^2) should be chosen. Applying these steps to our data, the best fitted ARCH models for each data series are shown in Table 6-7.

Table 6-7: The fitted ARCH models for monthly data

Country	Variable	ARCH model	
		Mean equation	Variance equation
Algeria	$\tilde{c}$	ARIMA[0,1,(2,7)]	ARCH(2)
	$\tilde{\pi}$		
	$\tilde{m}$	ARIMA[(1),1,0]	GARCH(1,1)
	$\tilde{r}$	ARIMA[0,1,(12)]	explosive
	s	ARIMA[(5,8),1,0]	explosive
	$\tilde{y}$	ARIMA[(10,13,24,29),1,(1,5,26,34,36)] SAR(12)	GARCH(1,1)
Japan	$\tilde{c}$	ARIMA[(1,2),1,(3)]	GARCH(1,1)
	$\tilde{\pi}$		
	$\tilde{m}$		
	$\tilde{r}$	ARIMA[(1,7,12),1,0]	explosive
	s	ARIMA[(1),1,(17,21)]	GARCH(1,1)
	$\tilde{y}$		
Kuwait	$\tilde{c}$	ARIMA[(25),1,(1,2)]	GARCH(1,1)
	$\tilde{\pi}$	ARIMA[(1,12),0,0]	GARCH(1,1)
	$\tilde{m}$	ARIMA[(2,3),1,0]	ARCH(1)
	$\tilde{r}$		explosive
	s	ARIMA[(4,17),1,0]	GARCH(1,1)
	$\tilde{y}$		
UK	$\tilde{c}$	ARIMA[(2,3,12,21,24),1,(1)]	ARCH(6) contains negative coeffcients
	$\tilde{\pi}$	ARIMA[(1),0,(6,12,20,24,36)]	GARCH(1,1)
	$\tilde{m}$	ARIMA[(1,7,8,22,23,26),1,(12,24,36)]	ARCH(1)
	$\tilde{r}$	ARIMA[0,1,(1,3,7)]	GARCH(1,1)
	s	ARIMA[0,1,(14)]	GARCH(1,1)
	$\tilde{y}$	ARIMA[(1),1,0]	GARCH(1,1)

The ARCH effect LM test was then applied again to make sure that the chosen models have successfully accounted for all ARCH effects in the data series. The results reveal that there is no ARCH effect left in the data. The correlogram of the standardized residuals indicate the absence of autocorrelation problem and the correlogram of the squared standardized residuals support the absence of any more ARCH effects in the data. The fitted values of each ARCH model (the prediction values of the conditional variance) are then used to approximate the volatility in each data series.

6.3.2 Results of the ADF unit root tests for quarterly differential data

Following the same procedure for quarterly observations the findings of the ADF tests are given first. Table 6-8 shows the results of the stationarity tests for the quarterly exchange rates and differential data series. The results are the same as those for monthly data except in two cases; namely, the Algerian consumption differential and the Libyan inflation rate differential.

Table 6-8: Results of the ADF unit root tests for quarterly differential data

Country	Variable	ADF (levels)	ADF (first differences)	Intercept & trend	Decision
Algeria	$\tilde{c}$	-4.15***(3)		C	I(0)
	$\tilde{\pi}$	-2.48(3)		C	I(0)
	$\tilde{m}$	-1.98(8)	-3.43**(7)	C&T	I(1)
	$\tilde{r}$	-1.12	-5.18***(2)	C	I(1)
	s	-1.49(2)	-5.18***	C&T	I(1)
	$\tilde{y}$	-1.15(4)	-2.94**(9)	C	I(1)
Canada	$\tilde{c}$	-3.81**(1)	-3.97***(4)	C&T	I(1)
	$\tilde{\pi}$	-3.54***(2)		C	I(0)
	$\tilde{m}$	-1.51(2)	-2.93**(6)	C&T	I(1)
	$\tilde{r}$	-2.15(7)	-10.25***	C	I(1)
	s	-1.02(3)	-2.91**(6)	C	I(1)
	$\tilde{y}$	-1.31	-5.01***(2)	C	I(1)
Germany	$\tilde{c}$	-2.21	-5.61***(1)	C	I(1)
	$\tilde{\pi}$	-2.17(3)	-13.33***	C	I(0)
	$\tilde{m}$	-0.27(2)	-2.63*(6)	C	I(1)
	$\tilde{r}$	-2.19(7)	-2.94**(4)	C	I(1)
	s	-1.84	-3.83***(3)	C	I(1)

Country	Variable	ADF (levels)	ADF (first differences)	Intercept & trend	Decision
	$\tilde{y}$	-0.08	-3.89***(3)	C	I(1)
Japan	$\tilde{c}$	-1.56(2)	-3.23**(5)	C	I(1)
	$\tilde{\pi}$	-5.17***(3)		C	I(0)
	$\tilde{m}$	-1.32(2)	-3.87***(2)	C	I(1)
	$\tilde{r}$	-2.15(7)	-7.70***	C	I(1)
	s	-0.83	-3.94***(3)	C	I(1)
	$\tilde{y}$	-0.73(1)	-3.41**(10)	C	I(1)
Kuwait	$\tilde{c}$	-0.90(4)	-15.32***	C	I(1)
	$\tilde{\pi}$	-3.71***(3)		C	I(0)
	$\tilde{m}$	-2.16(4)	-3.77***(8)	C	I(1)
	$\tilde{r}$	-1.17(7)	-3.22**(4)	C	I(1)
	s	-1.59(2)	-4.73***(3)	C	I(1)
	$\tilde{y}$	-2.65*	-9.70***	C	I(1)
Libya	$\tilde{c}$	-2.21(3)	-11.80***	C&T	I(1)
	$\tilde{\pi}$	-13.61***		C	I(0)
	$\tilde{m}$	-3.08(1)	-1.63(7)	C&T	I(1)
	$\tilde{r}$	-1.39(8)	-1.61(4)	C	I(1)
	s	-1.59(2)	-8.39***	C&T	I(1)
	$\tilde{y}$	-1.83(6)	-1.46(4)	C	I(1)
UK	$\tilde{c}$	-2.59(1)	-3.53***(6)	C&T	I(1)
	$\tilde{\pi}$	-3.14**(7)		C	I(0)
	$\tilde{m}$	-1.49(1)	-4.09***(3)	C&T	I(1)
	$\tilde{r}$	-3.10**	-8.69***	C	I(1)
	s	-2.24(2)	-2.97**(6)	C	I(1)
	$\tilde{y}$	-0.48(1)	-9.37***	C	I(1)
Venezuela	$\tilde{c}$	-4.57***		C	I(0)
	$\tilde{\pi}$	-3.34**		C	I(0)
	$\tilde{m}$	-1.85	-2.75*(2)	C&T	I(1)
	$\tilde{r}$	-2.86*	-6.59***	C	I(1)
	s	-2.11(1)	-10.06***	C&T	I(1)
	$\tilde{y}$	-3.04**	-5.13***	C	I(1)

Notes: The sample period starts in 1986q2 and 1989q2 in Libya and Venezuela respectively and in 1973q2 for the rest of countries and ends in 1998q4 for them all. * , ** and *** denote levels of significance at 10%, 5% and 1% respectively using the Mackinnon (1996) one-sided p-values. The number of lags (shown in parentheses) used for the ADF regressions are based on the modified AIC. C and T denote the inclusion of intercept and trend respectively. When the number of lags is not reported, it is in fact zero. Note that our decision with respect to the order of integration is not exclusively based on the results of the ADF test.

Applying the Box-Jenkins procedure to the quarterly data in selecting the appropriate ARMA models gave different results, which are displayed in Table 6-9. The results of the ARCH effect test results are shown in column (5). Clearly, the Kuwaiti exchange rate is the only series appears to have an ARCH effect. This support the idea that high frequency time series data are more likely to contain an ARCH effect compared to low frequency data. Thus, the rest of Kuwaiti data series were tested as well. The ARCH effect was also found in the Kuwaiti interest rate and income differentials as well as in the exchange rate.

Table 6-9: Results of the LM test for the ARCH effect using quarterly data

Country	Variable	ARIMA(p,d,q)	LM statistic	ARCH effect at lag
Algeria	$\tilde{c}$			
	$\tilde{\pi}$			
	$\tilde{m}$			
	$\tilde{r}$			
	s	ARIMA[(1),1,(2)]	3.03(0.08)[1]	NO
	$\tilde{y}$			
Canada	$\tilde{c}$			
	$\tilde{\pi}$			
	$\tilde{m}$			
	$\tilde{r}$			
	s	ARIMA[(3),1,0]	3.75(0.44)[4]	NO
	$\tilde{y}$			
Germany	$\tilde{c}$			
	$\tilde{\pi}$			
	$\tilde{m}$			
	$\tilde{r}$			
	s	ARIMA[(3,4,22),1,0]	3.44(0.97)[10]	NO
	$\tilde{y}$			
Japan	$\tilde{c}$			
	$\tilde{\pi}$			
	$\tilde{m}$			
	$\tilde{r}$			
	s	ARIMA[(16),1,0]	1.94(0.16)[1]	NO
	$\tilde{y}$			
Kuwait	$\tilde{c}$	ARIMA[0,1,(1,4,5)]	3.47(0.32)[3]	NO

Country	Variable	ARIMA(p,d,q)	LM statistic	ARCH effect at lag
	$\tilde{\pi}$	ARIMA[(10),0,(1,4,6)]	4.87(0.68)[7]	NO
	$\tilde{m}$	ARIMA[(16,24),1,0]	6.57(0.16)	NO
	$\tilde{r}$	ARIMA[(1,2,3,7),1,(16)]	26.72(0.00)[4]	2,4
	s	ARIMA[(2),1,(20)]	12.52(0.00)[3]	3
	$\tilde{y}$	ARIMA[0,1,(6,9)]	28.08(0.00)[1]	1
Libya	$\tilde{c}$			
	$\tilde{\pi}$			
	$\tilde{m}$			
	$\tilde{r}$			
	s	ARIMA[(2,11),1,0]	0.69(0.95)[4]	NO
	$\tilde{y}$			
UK	$\tilde{c}$			
	$\tilde{\pi}$			
	$\tilde{m}$			
	$\tilde{r}$			
	s	ARIMA[(1,3,7),1,0]	3.49(0.06)[1]	NO
	$\tilde{y}$			
Venezuela	$\tilde{c}$			
	$\tilde{\pi}$			
	$\tilde{m}$			
	$\tilde{r}$			
	s	ARIMA[(6),1,0]	5.61(0.47)[6]	NO
	$\tilde{y}$			

Having checked the presence of ARCH effect; the same steps are followed as were applied to monthly data in finding the optimal ARCH models for the Kuwaiti data. The best ARCH models are displayed in Table 6-10.

Given the results in Table 6-10, the ARCH model was explosive in the differential interest rate, so it had to be excluded. Accordingly, there are only two ARCH models for which it is believed pointless to go further in estimating ARCH-based volatility measures for these quarterly data. Thus, this volatility measure will be used merely for the monthly data.

Table 6-10: The fitted ARCH models for the Kuwaiti quarterly data

Country	Variable	ARCH model	
		Mean equation	Variance equation
Kuwait	$\tilde{c}$		
	$\tilde{\pi}$		
	$\tilde{m}$		
	$\tilde{r}$	ARIMA[(1,2,3,7),1,(16)]	explosive
	s	ARIMA[(2,10),1,(20)]	ARCH(3)
	$\tilde{y}$	ARIMA[0,1,(6,9)]	GARCH(1,1)

6.4 Results of the ADF unit root tests for the variables of Devereux-Lane's model using annual data

Next the variables used in the Devereux and Lane model are examined. Testing for the stationarity of these time series using the ADF test produced the results shown in Table 6-11. The findings indicate that the variables in all countries are either stationary or integrated of order one. These mixed results imply that the bounds testing approach should be applied to investigate the presence of cointegrated relationships between the series when using Devereux-Lane model. Therefore, we conclude that the Johansen approach cannot be applied to any of our equations as the unit root tests indicate the existence of variables that have mixed orders of integration; i.e. $I(0)$ and $I(1)$ for both the conventional models and Devereux-Lane model. On the other hand, since none of the variables in any of the models appears to be integrated of order higher than one, the use of the bounds testing method is strongly favoured as a valid approach to testing and estimating the long run relationships for our equations.

The results of the ADF tests for our data series indicate the suitability of the bounds testing approach in examining the presence of equilibrium long run relationships in our volatility equations.

Table 6-11: Results of the ADF test for annual data using the standard deviation

Country	Variable	ADF (levels)	ADF (first differences)	Outliers in	Intercept & trend	Decision
Algeria	cycle	-4.44***			C	I(0)
	finance	-0.21	-4.19***		C	I(1)
	s	-6.55***(5)		1991,1994	C	I(0)
	size	-2.91	-4.44***		C & T	I(1)
	trade	-1.53	-5.19***		C	I(1)
	extfin	-3.89**(1)	-3.70**		C & T	I(1)
Canada	cycle	-2.74*	-6.48***		C	I(1)
	finance	-4.12**(3)	-3.69**		C & T	I(1)
	s	-5.06***			C	I(0)
	size	-0.62	-3.64**(2)		C & T	I(1)
	trade	-1.57(1)	-3.24**		C & T	I(1)
Germany	cycle	-4.91***			C	I(0)
	finance	-0.79	-4.32***		C	I(1)
	s	-5.73***			C	I(0)
	size	-1.84	-4.81***		C & T	I(1)
	trade	-1.81	-5.19***		C	I(1)
Japan	cycle	-3.80***			C	I(0)
	finance	-1.63(1)	-2.92*		C & T	I(1)
	s	-4.19***			C	I(0)
	size	-2.42	-3.51**		C	I(1)
	trade	-13.24***		1985	C	I(0)
Kuwait	cycle	-6.34***		1990	C	I(0)
	finance	-2.01(1)	-3.15**		C	I(1)
	s	-3.95***		1987	C	I(0)
	size	-3.93**	-6.22***		C & T	I(1)
	trade	-3.82**			C & T	I(0)
Libya	cycle	-1.98(2)			C	I(0)
	finance	-3.89**(1)			C	I(0)
	s	-4.18***			C	I(0)
	size	-5.96***(2)			C & T	I(1)
	trade	-1.51	-3.33**		C	I(1)
UK	cycle	-3.24**			C	I(0)
	finance	-2.36*(1)	-3.13**		C & T	I(1)

Country	Variable	ADF (levels)	ADF (first differences)	Outliers in	Intercept & trend	Decision
	size	-1.78	-3.87***		C & T	I(1)
	s	-3.71**			C	I(0)
	trade	-2.71*			C	I(0)
Venezuela	cycle	-1.94			C	I(0)
	finance	-3.35(1)	-2.35		C & T	I(1)
	s	-3.35**(1)			C	I(0)
	size	-3.13	-3.64**(1)		C & T	I(1)
	trade	-4.00*	-3.59**	1996	C	I(1)
	extfin	-2.49	-2.61		C & T	I(1)

Notes: the sample period starts in 1986, 1983 and 1973 for Libya, Venezuela and the rest of countries respectively. Lags (between parentheses) are according to the SBC. When the number of lags is not reported, it is in fact zero. Note that our decision with respect to the order of integration is not exclusively based on the results of the ADF test.

Moreover, the results of estimating an ARCH model-based volatility proxies show that a few ARCH measures of volatility can be obtained using monthly data and almost none using quarterly data. The following chapter presents the empirical findings from the tests of the long run relationship.

Chapter seven

Cointegration Analysis

7.1 Introduction

Chapter five introduced the framework for the econometric methodologies applied for the purposes of this study. Chapter six then performed the unit root tests with the data series in order to determine their order of integration. This was a necessary step in deciding upon a potentially valid methodology to examine the presence of any equilibrium relationships in the equations developed earlier in the thesis. In the previous chapter it was found that the data sets involved are either $I(0)$ or $I(1)$, but none has been proven to be $I(2)$ or higher, for all models with deferent data frequencies and volatility measures. Accordingly, it was concluded that the bounds testing approach is the preferred method to test for the existence of long run relationships since it is applicable irrespective of whether the involved variables are purely $I(0)$, purely $I(1)$ or mutually cointegrated. This chapter provides the empirical findings from the application of this method to the conventional-based models and the Devereux-Lane and the augmented Devereux-Lane models. However, first it is necessary to decide whether the covariance terms in the equations (4.5)-(4.8) in chapter four are relevant, and hence should be included, or irrelevant, and hence should be omitted. In fact only in a few cases was it found that these covariances proxies matter for our main target equations, as reported in section 7.2. Section 7.3 applies the bounds testing approach to the target models using different volatility measures and data frequencies. The process of cointegration analysis for each model, frequency and measure involves three steps. In the first step the presence of a long run relationship between the dependent variable (exchange rate volatility) and its regressors was tested. In the second step, if evidence of a cointegrating relationship was found, the hypothesis of a unique cointegration relationship exists is tested, namely, to test for the absence of any feedback from the exchange rate volatility towards the regressors. If this hypothesis is accepted, the third step is to estimate the long run coefficients of the relationship, using the so-called "delta" method. This process, in particular, is applied to the following: (1) the conventional-based volatility models using monthly data and the simple measure; (2) the conventional-based volatility models using quarterly data and the simple measure; (3) the conventional-based volatility models using quarterly data and the standard deviation measure; (4) the conventional-based volatility models using monthly data and an ARCH-based measure; (5) the Devereux-Lane model using annual data and the standard deviation measure; and (6) the augmented Devereux-Lane model using annual data and the standard deviation measure. Comments, explanations and comparisons of the empirical results are then provided in the subsequent chapter of the thesis.

7.2 Approximating the covariances

As outlined above, our main interest is to discover whether a level relationship exists between the dependent variable (the volatility of exchange rate) and the regressors (the volatilities of macroeconomic fundamentals). For example, using equation (4.5) in chapter four (rewritten below for convenience), the existence of an equilibrium relationship between the volatility of exchange rate (Vs) and the volatility of the differentials of money supply ($V\tilde{m}$), industrial production ($V\tilde{y}$) and interest rate($V\tilde{r}$) is tested for:

$$Vs = V\tilde{m} + \alpha_1^2 V\tilde{y} + \alpha_2^2 V\tilde{r} + 2\alpha_1 Cov(\tilde{m},\tilde{y}) + 2\alpha_2 Cov(\tilde{m},\tilde{r}).$$

However, it is necessary to approximate the covariance terms in the above equation. Therefore, before starting the cointegration analysis for our main target equations; i.e. equations (4.5), (4.7) and (4.8), for the traditional-based volatility models, proxies should be found for the covariances which appear in those models.

A covariance between two variables refers to how these variables co-vary with one another. If the two variables are independent, their covariance is zero. However, if these two variables are not independent, this implies that they are correlated or related to each other with some sort of relationship. Therefore, one can investigate if a long run relationship exists between these two variables using a regression method. If the results support the presence of such long run relationships, a series of observations is needed to approximate the covariances that appear in our equations. The fitted values obtained from a significant relationship can be used as proxy for the covariance. Employing the fitted values as proxies for the covariance gives us a number of observations which match the number of observations included in the main equations. On the other hand, calculating the covariance for a given period will only give a single value for that period. Hence, we would then have an insufficient number of observations that would not match the number of data in the main equations.

Accordingly, before the parameters $2\alpha_1, 2\alpha_2$ (the coefficients of $Cov(\tilde{m},\tilde{y})$ and $Cov(\tilde{m},\tilde{r})$ respectively) are included in equation (4.5), for example, we have to check if long run relationships exist between the differential terms $\tilde{m}$, $\tilde{y}$ on the one hand and between $\tilde{m}$, $\tilde{r}$ on the other. If the results indicate the presence of such relationships, the long run coefficients whereby one can obtain the fitted values of such relationships are calculated. These fitted values will then be used as proxy for the covariances appearing in the equation above. These values are considered to be the best series for approximating to what extent the two variables co-vary with one another. Afterwards, the exchange rate variability is regressed on these proxies individually and jointly with the other differentials' volatilities in the main equations. If these proxies are found to be significant, they will be retained in the model. However, if they are found to be insignificant, they will be dropped from the model. Therefore, these steps will be followed for the covariance of each pair of variables in the main equations in order to decide whether or not such proxies are relevant to exchange rate volatility.

The ADF test results for the differentials were shown in table (6.5) in chapter six. Since these indicate that most of the variables are mixed in terms of their orders of integration, the bounds testing method is used to test if a long run relationship is present between such differential terms. The bounds testing approach[12] is used to all covariances appearing in equations (4.5), (4.7) and (4.8) for both monthly and quarterly observations. To save space and to avoid repeating the discussion of cointegration tests using the bounds testing method, the details of the tests and estimations are not reported here, but the final results of the cointegration tests for these relationships are shown.

For the monthly data, only three cases were found in which the covariance proxies seem to be relevant to exchange rate volatility. These three cases are the covariance between $\tilde{\pi}$ and $\tilde{m}$ in equation (4.7) for both Germany and the UK, and that between $\tilde{c}$ and $\tilde{m}$ in equation (4.8) for Germany. Therefore, for these three cases the fitted values will be included in the estimation of the main equations using monthly data.

[12] Full details of the bounds testing approach will be introduced when we address the main equations.

The same steps were performed to the quarterly data; however, no single pair was found to be relevant to exchange rate volatility. Therefore, they were excluded from the main equations when using quarterly data sets.

7.3 Results of the bounds testing approach to exchange rate volatility

We now come to the empirical investigation of the existence of equilibrium relationships between the volatility of exchange rates and the volatility of the fundamentals proposed by some economic theories. We start with equations (4.5), (4.7) and (4.8) introduced in chapter four, which were derived from the two versions of the monetary models of exchange rates and the Redux model. These equations are first rewritten without the covariances terms which were found to be insignificant in most cases as reported above. These proxies of the covariance terms will be included, however, where appropriate. Moreover, those equations are first will be written in a simple formula which, in turn, will be rewritten in the form of the ARDL-ECM model. Equation (4.7) therefore, can be written as follows.

$$vs_t = \beta_0 + \beta_1 vm_t + \beta_2 vy_t + \beta_3 vr_t + \beta_4 v\pi_t + \varepsilon_{0t} \tag{7.1}$$

where, vs is the exchange rate volatility, the symbols vm, vy, vr and $v\pi$ indicate the volatility of differentials in the money supply, industrial production index, short-run interest rate and inflation rate respectively; and ε is the stochastic error term. The signs of the coefficients (β's) are expected to be positive although they might be negative since there is no clear theoretical prediction of their signs. Equation (7.1) is based upon the sticky price real interest monetary model of exchange rate; hence, hereafter we will name it as SP.

The flexible price monetary model of exchange rates assumes that the coefficient of the inflation rate is zero, thus equation (4.5) can be rewritten as:

$$vs_t = \alpha_0 + \alpha_1 vm_t + \alpha_2 vy_t + \alpha_3 vr_t + \varepsilon_{1t} \tag{7.2}$$

This equation will be called FP hereafter.

The Redux model of exchange rate determination differs slightly from the above models insofar as it relates exchange rate changes to the changes in the money supply differential and consumption differential. Thus, equation (4.8) can be formed as follows:

$$vs_t = \gamma_0 + \gamma_1 vm_t + \gamma_2 vc_t + \varepsilon_{2t} \tag{7.3}$$

where vc is the volatility of per capita consumption difference. This equation will be referred to as RER.

To implement the bounds testing approach, one first modelled the above equations as a conditional ARDL-ECM model. Therefore, the above equations are further rewritten as follows starting with equation (7.1):

$$\begin{aligned}\Delta vs_t &= c_{s0} + c_{s1}t + \eta_{s1} vs_{t-1} + \eta_{s2} vm_{t-1} + \eta_{s3} vy_{t-1} + \eta_{s4} vr_{t-1} + \eta_{s5} v\pi_{t-1} \\ &+ \sum_{sj=1}^{sm1} \alpha_{sj} \Delta vs_{t-sj} + \sum_{sj=0}^{sm2} \beta_{sj} \Delta vm_{t-sj} + \sum_{sj=0}^{sm3} \gamma_{sj} \Delta vy_{t-sj} + \sum_{sj=0}^{sm4} \lambda_{sj} \Delta vr_{t-sj} \\ &+ \sum_{sj=0}^{sm5} \delta_{sj} \Delta v\pi_{t-sj} + \varphi_s D_t + \xi_{st}\end{aligned} \tag{7.4}$$

where c_{s0} and t are the drift and trend components, D is a vector of dummy variables included to allow for outlier observations in the series, ξ_s is assumed to be a vector of white noise error processes, and η_s's are the long run coefficient matrices for the lagged level dependent and forcing variables. The changes in the variables, referred to as Δ, represent the short run parameters. The short run dynamic structure was set to ensure that the residuals, ξ_s's, are white noise errors. The small sub-letter $_s$ refers to the sticky-price model.

Applying the same procedure to equations (7.2) and (7.3) gives us the following forms of ARDL-ECM:

$$\Delta vs_t = c_{f0} + c_{f1}t + \eta_{f1}vs_{t-1} + \eta_{f2}vm_{t-1} + \eta_{f3}vy_{t-1} + \eta_{f4}vr_{t-1} + \sum_{fj=1}^{fm1} \alpha_{fj}\Delta vs_{t-fj}$$
$$+ \sum_{fj=0}^{fm2} \beta_{fj}\Delta vm_{t-fj} + \sum_{fj=0}^{fm3} \gamma_{fj}\Delta vy_{t-fj} + \sum_{fj=0}^{fm4} \lambda_{fj}\Delta vr_{t-fj} + \varphi_f D_t + \xi_{ft}$$

(7.5)

$$\Delta vs_t = c_{r0} + c_{r1}t + \eta_{r1}vs_{t-1} + \eta_{r2}vm_{t-1} + \eta_{r3}vc_{t-1} + \sum_{rj=1}^{rm1} \alpha_{rj}\Delta vs_{t-rj}$$
$$+ \sum_{rj=0}^{rm2} \beta_{rj}\Delta vm_{t-rj} + \sum_{rj=0}^{rm3} \gamma_{rj}\Delta vc_{t-rj} + \varphi_r D_t + \xi_{rt}$$

(7.6)

The small sub-letters $_f$ and $_r$ refer to the flexible-price and Redux models respectively.

In order to test for the absence of level relationships between exchange rate volatility (vs) and the volatility of the macroeconomic fundamental differentials series in equation (7.4) above, it is first estimated by OLS and the F-statistics for the joint hypothesis of $\eta_{s1} = \eta_{s2} = \eta_{s3} = \eta_{s4} = \eta_{s5} = 0$ are calculated. The null hypothesis is tested through the exclusion of the lagged level variables vs_{t-1}, vm_{t-1}, vy_{t-1}, vr_{t-1} and $v\pi_{t-1}$ in equation (7.4). The null hypothesis is that $\eta_s's = 0$ against the alternative that at least one of the $\eta_s's \neq 0$. The joint null hypothesis for equation (7.5) is $\eta_{f1} = \eta_{f2} = \eta_{f3} = \eta_{f4} = \eta_{f5} = 0$ against the alternative that at least one of the $\eta_f{}'s \neq 0$. The joint null hypothesis for equation (7.6) is $\eta_{r1} = \eta_{r2} = \eta_{r3} = 0$ against the alternative that at least one of the $\eta_r{}'s \neq 0$.

Pesaran et al. (2001) have shown that the asymptotic distributions of the *F*-statistic are non-standard under the null hypothesis that there is no levels relationship between the variables included regardless of whether the regressors are *I*(0), *I*(1) or mutually cointegrated. Therefore, Pesaran et al. (2001) have provided two asymptotic critical value bounds test for cointegration when the system's variables are *I*(*d*) (where $0 \leq d \leq 1$): a lower value assuming only *I*(0) regressors, and an upper value assuming purely *I*(1) regressors. If the test statistic exceeds the upper critical value, we can conclude that a level relationship exists. If the test statistic falls below the lower critical value, we cannot reject the null hypothesis of no cointegration. If the statistic falls within the respective bounds, inference would be inconclusive. Critical values are also made available to encompass a range of different deterministic components. High *t*-values for the coefficients η_{s1} in (7.4), η_{f1} in (7.5) and η_{r1} in (7.6) confirm the presence of a level relationship (Pesaran et al., 2001).

The selection of an optimal lag length for the ARDL-ECM equations is decided, firstly, to ensure an absence of serial correlation in the estimated residuals and the problem of endogenous regressors (as in Corbin 2004), and secondly, on the basis of the Akaike information criterion (AIC) and the general to specific criterion (GSC)[13]. In Practice, it was found that the chosen lag length by the AIC is usually supported by the GSC. Following Pesaran et al. (2001) and to ensure comparability of results for different choices of lag lengths, all estimations used the same sample period with a

[13] The general-to-specific criterion means including a fairly high lag structure and then omitting the last lag until the last included lag is significant.

specific number of first observations reserved for the construction of lagged variables. The inclusion of a linear trend is decided depending upon the appearance of the plots of the series in the particular sample period under consideration and upon the significance of the trend coefficient in the regression. This empirical work is started by applying the bounds testing approach to monthly observations.

7.3.1 Results of the bounds testing tests for cointegration using monthly data and the simple volatility measure

The results of estimation of the ARDL-ECM model of equation (7.4) for monthly data and using the simple measure of volatility are presented in Table 7-1 showing each country, its computed bounds F-statistic, upper F-critical value, computed bounds t-statistic, upper t-critical value, AIC, and the chosen lag structure respectively.

Clearly, we can see from the table above that the computed F-statistic exceeds the upper bound F-critical value in all countries. This implies that equilibrium long run relationships exist between the exchange rate volatility and the volatility of some macroeconomic fundamentals in all economies considered. Moreover, the computed t-values for the lagged level dependent variable exceed the upper t-critical values at the 5% level of significance in all cases except in Kuwait where it is significant at the10% level. Such high t-values confirm the presence of long run relationships in the sample under consideration. Therefore, it is concluded that cointegration exists between the volatility of exchange rate and the volatilities of the fundamentals suggested by the sticky-price monetary model using monthly data and a simple measure of volatility.

Table 7-1: Results of the bounds F-tests for cointegration analysis of the SP using monthly data and the SM proxy

Country	F-statistic	Upper F-cv	t-statistic	Upper t-cv	AIC	Lags
Algeria	5.66**	4.01	-4.75**	-3.99	-10.05	7
Canada	5.22**	4.01	-5.02**	-3.99	-16.08	9
Germany	5.88**	3.79	-5.50**	-4.19	-13.06	8
Japan	10.00**	4.57	-4.73**	-4.36	-12.72	4
Kuwait	6.62**	4.01	-3.73*	-3.99	-17.21	7
Libya	5.40**	4.01	-6.81**	-3.99	-12.03	2
UK	4.40**	3.79	-5.86**	-4.19	-12.89	7
Venezuela	33.49**	4.01	-4.12**	-3.99	-7.43	9

The upper F and t critical values shown in the table are for the 5% level of significance. The period of estimation starts in 1974m04 for all countries except for Libya and Venezuela when it starts in 1986m03 and 1989m03 respectively and ends in 1998m12 for all countries. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. Note that the critical values of F and t-statistics for Germany and UK differ from that of other countries as the proxy for the covariance between $\widetilde{m}$ and $\widetilde{\pi}$ is included in the regression of both countries. Linear trend is included for the Japanese case. **, * indicate significance at 5% and 10% levels respectively.

Next the presence of level relationship is investigated between exchange rate volatility and volatilities of some macroeconomic fundamentals suggested by the flexible-price exchange rate monetary model (FP). The results are introduced in Table 7-2 using monthly data and a simple measure of volatility. Again as can be seen that the F-statistic exceeds the 5% upper critical value in all eight cases considered, thus providing strong evidence in favour of the existence of a cointegrating level relationship between the exchange rate variability and the suggested regressors.

Table 7-2: **Results of the bounds F-tests for cointegration analysis of the FP using monthly data and the SM proxy**

country	F-statistic	Upper F-cv	t-statistic	Upper t-cv	AIC	Lags
Algeria	7.00**	4.35	-4.83**	-3.78	-10.08	7
Canada	6.96**	4.35	-3.86**	-3.78	-16.15	10
Germany	8.02**	4.35	-5.90**	-3.78	-13.10	8
Japan	7.46**	5.07	-3.94*	-4.16	-12.71	9
Kuwait	8.55**	4.35	-3.80**	-3.78	-17.27	7
Libya	4.85**	4.35	-4.53**	-3.78	-12.12	9
UK	10.77**	4.35	-6.85**	-3.78	-12.98	3
Venezuela	33.28**	4.35	-4.77**	-3.78	-7.14	8

The upper F and t critical values shown in the table are for the 5% level of significance. The period of estimation starts in 1974m04 for all countries except for Libya and Venezuela when it starts in 1986m03 and 1989m03 respectively, and ends in 1998m12 for all countries. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. Linear trend is included for the Japanese case. **, * indicate significance at the 5% and 10% levels respectively.

The t-values also reinforce the conclusion reached using the F-statistic at the traditional level of significance, in that the coefficient of the lagged level dependent variable (η_{f1}) is statistically significant at the 5% level of significance in all cases except at 10% level of significance in Japan. This emphasizes the existence of long run relationships.

Results of the estimation of the bounds equation for the RER to test for cointegration are displayed in Table 7-3. From this table it can be noticed that the computed F-statistic is significant at 5% level of significance in all eight countries, indicating that long run links exist between the exchange rate fluctuation and the variability, as measured by the simple measure, in the money supply and consumption differentials using monthly data. These results are emphasized by the significant coefficient of the lagged level dependent variable (η_{r1}) in (7.6) as judged by their high t-values which exceed the upper critical values at the usual level of significance. The results of equation (7.6), which is based on the Redux model, resemble the results of the equations based on two versions of exchange rate monetary model, namely, the SP and FP models. As such, they give the same conclusion with respect to the presence of long run relationships between exchange rate variability and its suggested regressors.

Table 7-3: Results of the bounds F-tests for cointegration analysis of the RER using monthly data and the SM proxy

Country	F-statistic	Upper F-cv	t-statistic	Upper t-cv	AIC	Lags
Algeria	17.15**	4.85	-9.24**	-3.53	-10.21	2
Canada	8.09**	4.85	-4.03**	-3.53	-16.14	10
Germany	8.48**	4.35	-6.48**	-3.78	-13.13	8
Japan	10.33**	4.85	-3.79**	-3.53	-12.77	9
Kuwait	11.79**	4.85	-4.78**	-3.53	-17.32	7
Libya	5.06**	4.85	-4.68**	-3.53	-11.95	6
UK	14.25**	4.85	-7.06**	-3.53	-13.01	3
Venezuela	7.35**	4.85	-6.24**	-3.53	-5.35	2

The upper F and t critical values shown in the table are for the 5% level of significance. The period of estimation starts in 1974m03 for all countries except for Libya and Venezuela when it starts in 1986m03 and 1989m03 respectively, and ends in 1998m12 for all countries. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. Note that the critical values of F and t-statistics for Germany differ from that of other countries as the proxy for the covariance between $\widetilde{m}$ and $\widetilde{c}$ is included in its regression. **, * indicate significance at the 5% and 10% levels respectively.

Accordingly, it can be concluded that using monthly data and a simple measure of volatility, cointegrating relations exist between exchange rate volatility and the volatilities of some fundamentals as originally suggested by traditional exchange rate models. However, as emphasized in chapter five, the bounds testing approach assumes that the size of cointegrating space is unity, and thus it is important to ascertain whether the independent variables are not cointegrated among themselves and that they really are long run forcing.

This approach regards the variables on the right hand side of the model considered as long run forcing variables for the volatility of exchange rate, so there is no feedback from the level of vs_t. Given this assumption, it is presumed that the explanatory variables are not cointegrated among themselves and that, therefore, the size of the cointegrating space is restricted to unity (De Vita and Abbott, 2004).

To establish whether the regressors were in fact long run forcing, and hence to confirm the uniqueness of the cointegrating relations found, bounds test equations were first re-estimated for each regressor as a dependent variable against the short run dynamics of exchange rate volatility, the remaining regressors and the lagged levels of the dependent variable and the regressors[14].

The null hypothesis to be tested in the equation of footnote 14 below when the volatility of money stock differential is the dependent variable, is that $H_0 : \eta_{sm1} = 0$ against the alternative that $\eta_{sm1} \neq 0$.

If this null cannot be rejected, then the variables on the right hand side of the model considered are confirmed to be long run forcing regressors for vs_t (De Vita and Abbott, 2004).

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[14] To make this clear, the volatility of money supply differential in equation (7.4) is taken as an example. Thus equation (7.4) will be as follows:

$$\Delta vm_t = c_{sm0} + c_{sm1}t + \eta_{sm1} vs_{t-1} + \eta_{sm2} vm_{t-1} + \eta_{sm3} vy_{t-1} + \eta_{sm4} vr_{t-1} + \eta_{sm5} v\pi_{t-1} + \sum_{sj=0}^{sm1} \alpha_{smj} \Delta vs_{t-sj}$$
$$+ \sum_{sj=1}^{sm2} \beta_{smj} \Delta vm_{t-sj} + \sum_{sj=0}^{sm3} \gamma_{smj} \Delta vy_{t-sj} + \sum_{sj=0}^{sm4} \lambda_{smj} \Delta vr_{t-sj} + \sum_{sj=0}^{sm5} \delta_{smj} \Delta v\pi_{t-sj} + \varphi_{sm} D_t + \xi_{smt}$$

This procedure is performed on each regressor for equations (7.4)-(7.6) and the results of the bounds *t*-statistics are presented in the subsequent section.

7.3.1.1 Results of bounds testing *t*-tests for long run forcing using monthly data and the simple measure

The results of the bounds *t*-tests for traditional-based volatility models using monthly data and the simple volatility measure are shown in the following tables. From Table 7-4 below it can be clearly seen that the computed *t*-statistics fall below the lower critical *t*-values (introduced by Pesaran et al., 2001, tables CII (iii) and CIII (v)) in all countries except Venezuela. Therefore, we can conclude that feedback from exchange rate volatility towards regressors is indeed absent except towards the volatility of inflation differential in the case of Venezuela.

Table 7-4: Results of long run forcing *t*-tests of the SP using monthly data and the SM proxy

Country	*vm*	*vy*	*vr*	$v\pi$	$Cov(\pi, m)$
Algeria	-0.72	-1.14	-0.19	0.22	-
Canada	0.08	-1.38	2.29	-0.12	-
Germany	-0.27	0.42	1.41	1.37	-1.61
Japan	0.08	-0.80 t	1.58	0.65 t	-
Kuwait	-0.14	-0.54	0.10	0.31	-
Libya	-1.41	0.57	-2.45	-2.67 t	-
UK	0.04	0.95	0.81 t	-0.44	2.39
Venezuela	-0.62	0.14	-0.08	5.72**	-

The upper bound of the critical *t*-values at %5 is -3.99 for k=4 and -4.19 for k=5 for unrestricted intercept and no trend, where k is the number of regressors. The upper bound of the critical *t*-values at 5% is -4.36 for k=4 and -4.52 for k=5 for unrestricted intercept and unrestricted trend. The lower *t* critical value at 5% is -2.86 and -3.41 with and without trend respectively. t refers to the inclusion of trend in the regression. ** refers to a significant statistic at 5% level.

Thus we then proceed to derive the long run estimates by means of the bounds testing method for all countries. Although, this method should not be applied in the case of Venezuela because of the presence of more than one cointegrating relationship, it is conducted anyway for the sake of comparison, so that the results for different measures of volatility and/or frequency can be compared.

Table 7-5 contains the results of *t*-statistic bounds tests to check the absence of feedback from the exchange rate volatility to the regressors in equation (7.5).

Table 7-5: Results of long run forcing *t*-tests of the FP using monthly data and the SM proxy

Country	*vm*	*Vy*	*vr*
Algeria	-0.79	-0.90	-0.61
Canada	-0.08	-1.16	2.31
Germany	-1.03	-0.17	1.43 t
Japan	-0.29 t	-0.74 t	1.05
Kuwait	-0.18	-0.51	0.08
Libya	-0.53 t	0.60	-0.31
UK	-0.09	0.42 t	0.60 t
Venezuela	-0.22	-1.96	4.94**

The upper bound of the critical *t*-values at 5% is -3.78 for k=3 for unrestricted intercept and no trend, where k is the number of regressors. The upper bound of the critical *t*-values at 5% is -4.16 for k=3 for unrestricted intercept and unrestricted trend. The lower *t* critical value at 5% level is -2.86 and -3.41 with and without trend respectively. t refers to the inclusion of trend in the regression. ** refers to a significant statistic at 5% level.

The computed *t*-statistics in Table 7-5 reveal the existence of feedback from exchange rate variability to the volatility of interest rate differential in Venezuela only. Thus, the assumption of presence of unique level relationship as presumed above cannot be rejected for any of all eight countries for equation (7.5) except for Venezuela.

Therefore, the long run coefficients of the FP will be estimated using the bounds testing approach for these countries in addition to Venezuela for the purpose of comparison.

The results of the bounds *t*-statistics to check that the regressors in equation (7.6) are indeed long run forcing are shown in Table 7-6 below. The findings indicate that the null hypothesis is rejected at 5% level of significance in all countries in the sample meaning that no feedback exists from the exchange rate variability to the independent variables in these countries using equation (7.6). Thus, the bounds testing method will be applied for all countries in estimating the equilibrium parameters of the RER model.

Table 7-6: Results of long run forcing *t*-tests of the RER using monthly data and the SM proxy

Country	*vm*	*vc*	Cov(c,m)
Algeria	-0.69	-2.22 t	-
Canada	-0.63	1.43	-
Germany	0.03	0.20	-1.66
Japan	0.65	-1.09	-
Kuwait	0.80	-2.29 t	-
Libya	0.19 t	0.51	-
UK	0.28	2.66	-
Venezuela	-0.68 t	1.08	-

The upper bound of the *t* critical values at 5% level of significance is -3.53 for k=2 and -3.78 for k=3 for unrestricted intercept and no trend, where k is the number of regressors. The upper bound of the *t* critical values at 5% is -3.95 for k=2 and -4.16 for k=3 for unrestricted intercept and unrestricted trend. The lower *t* critical value at 5% is -2.86 and -3.41 with and without trend respectively. t refers to the inclusion of trend in the regression. ** refers to a significant statistic at 5% level.

7.3.1.2 The estimation of long run coefficients using monthly data and the simple volatility measure

Pesaran et al. (2001) referred to the importance of keeping the coefficients of lagged changes unrestricted when testing the hypothesis of no cointegration in order to avoid subjecting these tests to a pre-testing problem. However, they advised the use of more parsimonious specification when estimating level effects and short run dynamics. Therefore, this advice is taken here to save degrees of freedom and to avoid any over-parameterized specification using the same lag structure criteria; i.e. non-serial correlation, non-endogenous regressors, AIC and GSC, as used before, to choose the appropriate lag length in estimating the long run parameters of equations (7.1)-(7.3) for the SP, FP and RER, respectively.

For the estimation of the long run parameters there are two competing models which can give identical results as reported Bardsen (1989). These are the ARDL and ARDL-ECM, as described in chapter five. However, Bardsen reported that the ARDL-ECM (equations (7.4)-(7.6)) explicitly give the short run dynamics in the differenced terms, and the long run coefficients can easily be obtained by dividing the regressors' coefficients by the coefficient of the lagged level dependent variable multiplied by negative sign. Bardsen indicated that the ARDL-ECM also has the merit of the simplification in the computation of the variances of the coefficients regardless of the number of lags involved unlike with the ARDL form (Bardsen, 1989). Moreover, it is essential for the ARDL to involve continuous lags for the calculation of the long run parameters, whereas these parameters can be computed from the ARDL-ECM whether the lags are continuous or not. Therefore, to avoid the problem of over parameterization and to save degrees of freedom, it is more convenient to use discontinuous lags by omitting differenced variables that have low *t*-values; or more precisely, coefficients with t-value less than 1.6. Consequently, the current form of our equations (7.4)-(7.6); that is the ARDL-ECM model, is used to obtain the long run parameters of equations (7.1)-(7.3).

The 'delta' method (Δ -method) is used to provide the estimates of long run coefficients $\beta's$, $\alpha's$ and $\gamma's$ in equations (7.1)-(7.3) and their standard errors. The 'delta' method has a number of advantages, in that the estimated level coefficients are consistent irrespective of whether the underlying series are $I(0)$ or $I(1)$, and this method performs particularly well in small samples (Corbin, 2004).

The inclusion of a deterministic linear trend in the regressions is decided according to the plots of the series and the significance of trend in the regression.

If the time series plots depict the existence of outliers, dummy variables are included to take into account their effects[15].

The results of the estimation of the long run coefficients of the SP are shown in Table 7-7 below.

It can be clearly seen from table above that none of the regressors, except the intercept, appear to be significant in determining the volatility of exchange rate according to t-values of the coefficients in the Algerian and Kuwaiti cases. The case in which there is no one single significant coefficient suggests extreme multicollinearity[16] although the F-bounds test indicate the presence of cointegrating relationship. The volatility of money supply differential is significant in both Japan and the UK but with different signs. It is also positive and significant at 10% level of significance in Germany.

[15] The inclusion of a 'one-off' dummy variables dose not affect the asymptotic theory developed by Pesaran et al. (2001) and the critical values, if the fraction of periods in which the dummy variables are non-zero tend to zero with the sample size. In the present study the fraction of non-zero dummy observations is at maximum 1%.

[16] As suggested by Pesaran via email address: mhp1@cam.ac.uk.

Table 7-7: Estimates of the long run coefficients of the SP model using monthly data and the SM proxy;

$$vs_t = \beta_0 + \beta_1 vm_t + \beta_2 vy_t + \beta_3 vr_t + \beta_4 v\pi_t + \varepsilon_{0t}$$

Country	β_0	β_1	β_2	β_3	β_4	*trend*	Cov(π, m)	$\bar{R}^2$	$\chi^2_{sc}(4)$	χ^2_H	$\chi^2_{FF}(1)$
Algeria	0.0003** (2.18)	-0.15 (-0.40)	0.02 (0.31)	-0.0003 (-0.23)	-0.50 (-0.76)	-	-	0.47	3.32[0.51]	17.54[0.35]	2.73[0.13]
Canada	0.00002** (4.03)	0.02 (1.11)	0.13** (2.08)	0.001** (4.82)	-0.62 (-0.75)	-	-	0.78	4.41[0.35]	36.60[0.81]	3.50[0.07]
Germany	0.0002** (4.96)	0.28* (1.83)	-0.04 (-0.37)	-0.00001 (-0.06)	15.57** (2.79)	-	0.03 (0.36)	0.57	4.05[0.40]	33.33[0.50]	3.70[0.055]
Japan	0.0001 (1.51)	-0.12** (-2.28)	-0.46 (-1.54)	0.004* (1.70)	2.86 (0.68)	0.0000006 (1.55)	-	0.65	1.43[0.84]	72.77[0.000]	0.16[0.69]
Kuwait	0.00002** (2.36)	-0.004 (-0.62)	-0.00009 (-0.55)	0.0004 (0.64)	-0.06 (-0.39)	-	-	0.50	1.25[0.87]	47.42[0.14]	1.31[0.25]
Libya	-0.000001 (-0.06)	-0.0006 (-0.03)	0.06** (2.56)	0.02** (2.31)	-0.19 (-0.21)	-	-	0.96	2.53[0.64]	114.21[0.000]	0.13[0.72]
UK	0.00006 (1.23)	0.66** (3.39)	-0.18 (-1.47)	0.002 (1.39)	0.17 (0.37)	-	0.003 (0.16)	0.68	10.27[0.04]	93.94[0.000]	0.34[0.56]
Venezuela	-0.00001 (-0.06)	-0.05 (-1.04)	0.003 (0.02)	0.02** (5.92)	10.76** (5.60)	-	-	0.99	8.86[0.07]	40.35[0.54]	31.27[0.000]

Notes: Dummy variables were excluded from regressions for Algeria and Kuwait because they caused serial correlation, heteroscedasticity and/or misspecification problems. White's method was used to correct for heteroscedasticity in Japan and Libya. The Newey-West method was used to correct for serial correlation and heteroscedasticity in the UK.**, * indicate significance at the 5% and 10% levels respectively. $\chi^2_{sc}(4)$, χ^2_H and $\chi^2_{FF}(1)$ denote chi-squared statistics to test for no residual serial correlation up to four order, homoscedasticity and no functional form misspecification respectively with p-values given in square brackets.

The volatilities of income and interest rate differentials appear to be significant in cases of Canada and Libya. They also have positive signs indicating that an increase in these two variables leads to a rise in the volatility of exchange rate in both countries. The variability of inflation rate differential seems to be significant only in the case of Germany. It is positive and very large. In the Venezuelan case, the only significant variable is the volatility of interest rate differential. This variable is also positive and significant at 10% level of significance in Japan. Note that for the purpose of consistency although the trend in Japan and covariance proxy in Germany and UK were not significant, they were kept as they were significant in the less parsimonious regressions. This will be the case for the rest of results. The regression fits reasonably well and passes the diagnostic tests against autocorrelation, heteroscedasticity and misspecification in almost all countries. However, it failed the functional form misspecification test at the 0.05 level in Venezuela. This failure may be linked to the presence of some non-linear effects or other reasons that our model is incapable of taking into account. The adjusted R squared ranges from 0.47 in Algeria to 0.99 in Venezuela.

The results of estimating the FP, as shown in Table 7-8, are very similar to those of the SP. The only difference is that the volatility of money supply differential, which was significant at 10% in Germany for SP, is now positive and statistically significant at 5% level of significance. Therefore, we can say that the equations based on the sticky-price and flexible-price exchange rate monetary models have given similar findings.

Table 7-8: Estimates of the long run coefficients of the FP model using monthly data and the SM proxy;

$$vs_t = \alpha_0 + \alpha_1 vm_t + \alpha_2 vy_t + \alpha_3 vr_t + \varepsilon_{1t}$$

Country	α_0	α_1	α_2	α_3	*trend*	$\overline{R}^2$	$\chi^2_{sc}(4)$	χ^2_H	$\chi^2_{FF}(1)$
Algeria	0.0002** (2.24)	-0.20 (-0.55)	0.02 (0.32)	-0.0006 (-0.30)	-	0.47	2.07[0.72]	16.23[0.18]	3.48[0.08]
Canada	0.00002** (4.06)	0.01 (0.96)	0.14** (2.13)	0.001** (4.87)	-	0.78	4.48[0.34]	37.33[0.63]	3.69[0.06]
Germany	0.0002** (6.42)	0.76** (2.20)	-0.05 (-0.45)	-0.000005 (-0.02)	-	0.56	5.25[0.26]	23.16[0.62]	3.69[0.06]
Japan	0.0002** (2.56)	-0.12** (-2.33)	-0.45 (-1.56)	0.004* (1.67)	0.0000004 (0.98)	0.65	1.51[0.82]	70.54[0.000]	0.13[0.72]
Kuwait	0.00002** (3.11)	-0.005 (-0.82)	-0.00008 (-0.63)	0.0002 (0.38)	-	0.49	7.44[0.11]	32.06[0.27]	1.27[0.26]
Libya	-0.000003 (-0.15)	-0.0004 (-0.02)	0.06** (2.57)	0.03** (2.41)	-	0.96	2.55[0.64]	114.17[0.000]	0.12[0.73]
UK	-0.0001 (-1.09)	1.83** (2.49)	-0.10 (-0.90)	0.002 (1.27)	-	0.68	20.72[0.000]	92.03[0.000]	1.07[0.30]
Venezuela	-0.0004 (-1.36)	-0.04 (-0.76)	0.70 (1.18)	0.05** (4.73)	-	0.99	10.92[0.028]	65.23[0.000]	36.78[0.000]

Notes: Dummy variables were excluded from regression for Algeria and Kuwait because they caused serial correlation, heteroscedasticity and/or misspecification problems. White's method was used to correct for heteroscedasticity in Japan and Libya. The Newey-West method was used to correct for serial correlation and heteroscedasticity in the UK and Venezuela. **, * indicate significance at %5 and 10% level respectively. $\chi^2_{sc}(4)$, χ^2_H and $\chi^2_{FF}(1)$ denote chi-squared statistics to test for no residual serial correlation up to four order, homoscedasticity and no functional form misspecification respectively with p-values given in square brackets.

Table 7-9 contains the results of estimating the level effects of the RER, which is grounded in the Redux model. From the table, the volatility of money supply differential is statistically significant at 5% in Germany, Japan and the UK, which is consistent with the two models above. The volatility of consumption differential is only significant in Algeria and Canada with a positive sign.

All significant variables in the three models have the expected positive sign except the volatility of money supply differential in Japan. Thus it can be concluded that an increase in volatilities of these fundamentals do indeed cause a rise in the exchange rate volatility. Overall, except in a few cases, the models explain 50 per cent or more of the behaviour of exchange rates and pass a set of diagnostic tests used in most cases.

Table 7-9: Estimates of the long run coefficients of the RER model using monthly data and the SM proxy;

$$vs_t = \gamma_0 + \gamma_1 vm_t + \gamma_2 vc_t + \varepsilon_{2t}$$

Country	γ_0	γ_1	γ_2	*trend*	$\mathrm{Cov}(c,m)$	$\overline{R}^2$	$\chi^2_{sc}(4)$	χ^2_H	$\chi^2_{FF}(1)$
Algeria	0.000009 (0.33)	-0.18 (-1.65)	0.27** (4.25)	-	-	0.68	9.30[0.054]	181.84[0.000]	9.43[0.002]
Canada	0.00003** (6.30)	0.01 (0.71)	0.007** (2.94)	-	-	0.75	8.12[0.09]	41.18[0.29]	0.87[0.35]
Germany	-0.002 (-1.05)	1.07** (2.68)	0.003 (0.16)	-	-0.005 (-1.13)	0.56	1.59[0.81]	28.58[0.73]	6.26[0.012]
Japan	0.0002** (2.96)	-0.12** (-2.06)	-0.02 (-1.60)	0.0000004 (1.01)	-	0.64	3.09[0.54]	64.57[0.000]	0.05[0.82]
Kuwait	0.00002** (3.10)	-0.003 (-0.51)	0.0002 (0.26)	-	-	0.50	7.35[0.12]	39.64[0.06]	0.94[0.33]
Libya	0.00007** (4.05)	0.03 (1.55)	0.002 (0.41)	-	-	0.94	6.60[0.16]	7.10[0.42]	0.80[0.37]
UK	-0.00008 (-0.85)	1.66** (2.54)	0.0008 (0.07)	-	-	0.68	12.50[0.014]	97.61[0.000]	1.35[0.25]
Venezuela	0.003 (1.38)	-1.02 (-1.12)	0.11 (0.52)	-	-	0.55	0.05[0.99]	6.20[0.91]	0.02[0.88]

Notes: Dummy variables were excluded from regression for Algeria and Kuwait because they caused serial correlation, heteroscedasticity and/or misspecification problems. White's method was used to correct for heteroscedasticity in Algeria and Japan. The Newey-west method was used to correct for serial correlation and heteroscedasticity in the UK. **, * indicate significance at 5% and 10% level respectively. $\chi^2_{sc}(4)$, χ^2_H and $\chi^2_{FF}(1)$ denote chi-squared statistics to test for no residual serial correlation up to four order, homoscedasticity and no functional form misspecification respectively, with p-values given in square brackets.

7.3.2 Results of the bounds testing tests for cointegration using quarterly data and the simple volatility measure

The equations (7.4)-(7.6) are now re-estimated by OLS using quarterly data and the same simple volatility measure. The maximum lag length will be set at eight quarters for all countries except for Libya and Venezuela where it is set at four lags as a result of smaller number of observations. We start with equation (7.4) to compute the *F*-statistic of the bounds test to discover if cointegrating relationships exist between the exchange rate volatility and volatility of the fundamentals based on variables introduced by the sticky-price monetary model for exchange rate (SP). The results of the bounds *F*-statistics for cointegration in equation (7.4) are demonstrated in Table 7-10. All F and *t*-critical values are for an unrestricted intercept and no trend as there is no trend included in any of the regressions (7.4)-(7.6) using quarterly data.

The results of the *F*-statistics reveal that long run relationships exist between exchange rate variability and variability of the fundamental differentials suggested by the sticky-price monetary-based volatility model of exchange rate (SP) for all countries except Algeria, where the computed *F*-value falls within the inconclusive band. The computed *t*-values confirm the presence of such level relationships at 5% in all cases found by the *F*-statistic, except for the Kuwaiti case where the *t*-value falls in the inconclusive area, as well as the Algerian and the Canadian cases where the *t*-values are significant at the 10% level. Overall, we can conclude that level relationships are present in the SP in all countries except Algeria when using quarterly data set and a simple volatility measure. This result is globally consistent with those of the same model using monthly data as shown in Table 7-1.

Table 7-10: Results of the bounds *F*-tests for cointegration analysis of the SP using quarterly data and the SM proxy

Country	*F*-statistic	*t*-statistic	AIC	Lags
Algeria	2.87inc	-3.94*	-8.28	4
Canada	5.80**	-3.67*	-14.87	3
Germany	4.42**	-4.32**	-10.90	3
Japan	5.48**	-4.31**	-10.81	3
Kuwait	5.15**	-3.58inc	-15.10	2
Libya	4.04**	-4.30**	-11.51	2
UK	6.70**	-4.67**	-11.38	2
Venezuela	22.51**	-6.75**	-6.23	0

The lower and upper *F*-critical values are 2.86 and 4.01 respectively at 5%, 2.45 and 3.52 respectively at 10% level of significance. The lower and upper *t*-critical values are -2.86 and -3.99 respectively at 5%, -2.57 and -3.66 respectively at 10% level of significance. The period of estimation starts in 1975q3 for all countries except for Libya and Venezuela when it starts in 1986q2 and 1989q2 respectively and ends in 1998q4 for all countries. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. inc means that the computed statistic falls within the inconclusive band. **, * refer to significance at 5% and 10% levels respectively.

Table 7-11 displays the results of the bounds *F* and *t* cointegration tests for the FP based on the flexible-price monetary model of exchange rate (FP). The similarity is apparent of the results for equations based on both the SP and FP. As shown in Table 7-11, the bounds *F*-statistics indicate the existence of cointegrating relationships in all countries except for Algeria, where the computed F falls below the lower critical value at the 5% level, but in the inconclusive area considering 10% level.

Moreover, the *t*-values confirm the findings of the *F*-statistics of level relationships at 5% in all seven countries except Kuwait where the *t*-value is significant at 10%. Therefore, by and large we can conclude that cointegrating relationships exist for all countries except Algeria using the FP. These findings generally resemble those for monthly data in Table 7-2.

Table 7-11: Results of the bounds F-tests for cointegration analysis of the FP using quarterly data and the SM proxy

Country	F-statistic	t-statistic	AIC	Lags
Algeria	3.06inc	-4.47**	-8.28	4
Canada	6.43**	-3.88**	-14.87	3
Germany	5.66**	-4.46**	-10.96	3
Japan	7.16**	-4.23**	-10.88	3
Kuwait	4.75**	-3.72*	-15.12	3
Libya	5.25**	-4.01**	-11.74	3
UK	8.37**	-4.86**	-11.43	2
Venezuela	30.83**	-5.91**	-6.13	0

The lower and upper F-critical values are 3.23 and 4.35 respectively at 5%, 2.72 and 3.77 respectively at 10% level of significance. The lower and upper t-critical values are -2.86 and -3.78 respectively at 5%, -2.57 and -3.46 respectively at 10% level of significance. The period of estimation starts in 1975q3 for all countries except for Libya and Venezuela when it starts in 1986q2 and 1989q2 respectively and ends in 1998q4 for all countries. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. inc refers to that the computed statistic falls within the inconclusive band. **, * refer to significance at 5% and 10% levels respectively.

The results of the bounds testing F-statistics and the t-statistics for cointegration for equation (7.6) using quarterly data are displayed in Table 7-12. The computed F-statistics indicate that level relationships exist for all countries. The t-values of the bounds testing support the results of the F-statistic test at 5% level, except for Canada and Kuwait at 10%. These findings favour the conclusion that long run relationships indeed exist between volatility of exchange rate and volatility of money supply differential and consumption differential using quarterly data for all economies considered. These results are largely similar to those of the same model using monthly data set shown in Table 7-3.

Table 7-12: Results of the bounds *F*-tests for cointegration analysis of the RER using quarter data and the SM proxy

Country	*F*-statistic	*t*-statistic	AIC	Lags
Algeria	5.54**	-4.03**	-8.47	5
Canada	5.28**	-3.22*	-14.80	6
Germany	6.12**	-4.36**	-11.15	6
Japan	7.08**	-4.10**	-10.85	4
Kuwait	8.62**	-3.46*	-15.21	2
Libya	7.08**	-18.30**	-13.64	3
UK	9.68**	-5.00**	-11.40	2
Venezuela	6.45**	-9.52**	-6.15	1

The lower and upper *F*-critical values are 3.79 and 4.85 respectively at 5%, 3.17 and 4.14 respectively at 10% level of significance. The lower and upper *t*-critical values are -2.86 and -3.53 respectively at 5%, -2.57 and -3.21 respectively at 10% level of significance. The period of estimation starts in 1975q3 for all countries except for Libya and Venezuela when it starts in 1986q2 and 1989q2 respectively and ends in 1998q4 for all countries. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. **, * refer to significance at 5% and 10% levels respectively.

The results for the quarterly data are slightly different from those using monthly data, in that for the latter level relationships were found to be presents in all equations for all countries, whereas for less frequent data (quarterly) no level relationships were found to be exist in Algeria using the SP and FP.

7.3.2.1 Results of bounds testing *t*-tests for long run forcing using quarterly data and the simple measure

In order to establish whether the regressors based on variables originally proposed by the traditional models were in fact long run forcing, and hence confirming the uniqueness of the cointegrating vectors found, equations (7.4)-(7.6) were re-estimated for each regressor as a dependent variable. Then, a variety of the bounds test suggested originally by Banerjee et al. (1998) which is based on the *t*-test for the null hypothesis that the coefficient of the lagged exchange rate volatility is zero, is performed. The results of the *t*-test for equation (7.4) are shown in Table 7-13.

The bounds *t*-statistic tests reveal that more than one level relationship exists in the SP for Venezuela. Thus the assumption of unique cointegrating relationship is accepted in six countries, that is Canada, Germany, Japan, Kuwait, Libya and the UK. Algeria had already been excluded since no level relationship was found to exist according to the *F*-tests when exchange rate volatility was considered as a dependent variable.

Table 7-13: Results of long run forcing *t*-tests of the SP using quarterly data and the SM proxy

Country	*vm*	*vy*	*vr*	$v\pi$
Canada	-0.33	-0.38	-1.23	-1.14
Germany	-0.17	-0.62	-0.97	0.50
Japan	1.06	0.33	1.14	1.02
Kuwait	-0.49	-0.28	0.55	-0.68 t
Libya	-0.62	0.50	0.38	2.04
UK	0.27	-1.06	0.35	1.59
Venezuela	-1.37	-0.50	1.23	3.88**

The upper bound of the critical *t*-values at 5% is -3.99 for k=4 for unrestricted intercept and no trend. The upper bound of the critical *t*-values at 5% is -4.36 for k=4 for unrestricted intercept and unrestricted trend. The lower *t* critical value at 5% is -2.86 and -3.41 with and without trend respectively. t refers to the inclusion of trend in the regression. ** refers to a significant statistic at 5% level.

The results for the monthly and quarterly data show that there is more than one cointegrating relationship for Venezuela, since feedback from exchange rate variability to the volatility of inflation rate differential was found in both data sets.

Table 7-14 contains the results of the bounds *t*-tests for the absence of feedback in the FP using quarterly data. Note that the assumption of one long run relationship is rejected only in Venezuela. Therefore, the absence of feedback is confirmed for the rest of the sample in which case we may proceed to estimate level effects using the bounds testing approach for all countries including Venezuela.

Table 7-14: Results of long run forcing *t*-tests of the FP using quarterly data and the SM proxy

Country	*vm*	*Vy*	*vr*
Canada	-0.79	-0.43	-1.10
Germany	-0.62	0.38	0.21 t
Japan	0.94	-0.44	0.79
Kuwait	-0.67	-0.36	-0.12
Libya	-1.62	0.40	0.71
UK	0.61	-0.43	-0.25
Venezuela	0.78	-1.69	3.83**

The upper bound of the *t*-critical values at 5% is -3.78 for k=3 for unrestricted intercept and no trend. The upper bound of the *t*-critical values at 5% is -4.16 for k=3 for unrestricted intercept and unrestricted trend. The lower t critical value at 5% is -2.86 and -3.41 with and without trend respectively. t refers to the inclusion of trend in the regression. ** refers to a significant statistic at 5% level.

Again the findings of the *t*-tests using monthly data (as reported in Table 7-5) and the quarterly data with the simple measure as proxy for volatility are consistent.

Table 7-15 shows the results of testing for no feedback being present in the RER. The results indicate the validity of the assumption of unique level relationship in all cases, confirming the validity of applying the bounds testing approach in estimating the coefficients of long run relationships.

Table 7-15: Results of long run forcing *t*-tests of the RER using quarterly data and the SM proxy

Country	*vm*	*vc*
Algeria	-0.42 t	-0.74
Canada	-1.59	-0.54
Germany	-0.27	0.40
Japan	1.68	0.55
Kuwait	-0.09	-1.55
Libya	-0.77	0.15
UK	-0.15	-0.50
Venezuela	0.60	0.28

The upper bound of the *t*-critical values at 5% is -3.53 for k=2 for unrestricted intercept and no trend. The upper bound of the *t*-critical values at 5% is -3.95 for k=2 for unrestricted intercept and unrestricted trend. The lower t critical value at 5% is -2.86 and -3.41 with and without trend respectively. t refers to the inclusion of trend in the regression. * refers to a significant statistic at 5% level.

Generally speaking the results of the tests for long run forcing for monthly and quarterly data for the three models are identical in terms of the countries that have unique or more than one cointegrating relationship and for the variables that appear to be cointegrated among themselves.

7.3.2.2 The estimation of long run coefficients using quarterly data and the simple volatility measure

Using the same method and steps as used for monthly data, the results of the estimation of long run coefficients of the SP for quarterly data are shown in Table 7-16 below.

Table 7-16: Estimates of the long run coefficients of the SP model using quarterly data and the SM proxy;

$$vs_t = \beta_0 + \beta_1 vm_t + \beta_2 vy_t + \beta_3 vr_t + \beta_4 v\pi_t + \varepsilon_{0t}$$

Country	β_0	β_1	β_2	β_3	β_4	$\overline{R}^2$	$\chi^2_{sc}(4)$	χ^2_H	$\chi^2_{FF}(1)$
Canada	0.00004 (1.59)	0.04 (1.21)	0.09 (0.76)	0.003** (2.00)	2.11** (2.14)	0.68	3.23[0.52]	20.06[0.86]	0.15[0.70]
Germany	0.00007 (0.56)	5.43** (5.71)	0.96 (1.52)	0.004* (1.69)	-11.15** (-2.20)	0.69	2.59[0.63]	47.22[0.10]	6.81[0.01]
Japan	0.0008** (5.34)	-0.38 (-0.50)	-0.14 (-0.15)	0.006** (2.13)	-9.66* (-1.74)	0.67	6.18[0.19]	29.99[0.88]	8.45[0.004]
Kuwait	0.00006** (2.12)	0.0009 (0.15)	0.001 (1.55)	-0.0003 (-0.39)	-0.56 (-0.94)	0.53	2.44[0.66]	11.13[0.97]	3.80[0.051]
Libya	0.00009 (1.16)	0.11** (2.02)	0.07** (1.99)	0.005 (1.19)	-0.008* (-1.83)	0.94	4.94[0.29]	15.41[0.93]	8.23[0.004]
UK	0.0003** (4.65)	0.12** (4.02)	1.35* (1.96)	-0.0006 (-0.42)	-0.25 (-0.40)	0.85	6.80[0.15]	27.40[0.90]	3.81[0.053]
Venezuela	0.002 (1.22)	0.12 (0.45)	-0.61 (-0.82)	-0.002 (-0.62)	4.34** (3.39)	0.87	1.43[0.84]	10.67[0.83]	0.02[0.90]

Notes: Dummy variables were excluded from regressions for Kuwait because they cause serial correlation, heteroscedasticity and/or misspecification problems. **, * indicate significance at 5% and 10% level respectively. $\chi^2_{sc}(4)$, χ^2_H and $\chi^2_{FF}(1)$ denote chi-squared statistics to test for no residual serial correlation up to four order, homoscedasticity and no functional form misspecification respectively with p-values given in square brackets.

From the table above we can say that the volatility of the money supply differential is positive and significant in Germany, Libya and the UK. The income differential variability is positive and significant in Libya and the UK. The variability of interest rate differential is statistically significantly different from zero and has a positive sign in Canada, Germany and Japan. The inflation rate differential volatility is positive and significant in Canada and Venezuela and negative and significant in Germany, Japan and Libya. There is no single significant variable in the case of Kuwait except for the intercept. This again may be due to the presence of severe multicollinearity.

The results of FP in the Table 7-17 clearly show that the volatility of income differential is statistically insignificant in all countries considered. This result is identical to that for monthly data as shown in Table 7-8. The volatility of money supply differential is positive and significant in Canada, Germany, Libya and the UK. The variability of interest differential has positive sign and is significantly different from zero in Canada, Japan and Venezuela. Again, none of the regressors appears to be significant in the Kuwaiti data except for the constant term.

Table 7-17: Estimates of the long run coefficients of the FP model using quarterly data and the SM proxy;

$$vs_t = \alpha_0 + \alpha_1 vm_t + \alpha_2 vy_t + \alpha_3 vr_t + \varepsilon_{1t}$$

Country	α_0	α_1	α_2	α_3	$\overline{R}^2$	$\chi^2_{sc}(4)$	χ^2_H	$\chi^2_{FF}(1)$
Canada	0.00004** (2.21)	0.07** (2.16)	0.06 (0.56)	0.003** (2.36)	0.66	3.02[0.55]	17.74[0.72]	0.18[0.67]
Germany	0.00004 (0.31)	4.84** (5.09)	0.96 (1.19)	0.002 (1.07)	0.68	1.28[0.86]	46.31[0.12]	6.44[0.01]
Japan	0.0006** (4.93)	0.07 (0.10)	-0.20 (-0.21)	0.006** (1.98)	0.66	4.00[0.41]	28.51[0.81]	7.68[0.006]
Kuwait	0.00004* (1.89)	-0.0007 (-0.11)	0.001 (1.47)	-0.0002 (-0.28)	0.53	2.42[0.66]	6.26[0.99]	3.72[0.06]
Libya	0.00009 (1.09)	0.12** (2.11)	0.03 (1.25)	0.002 (0.41)	0.93	4.28[0.37]	18.31[0.50]	6.49[0.01]
UK	0.0004** (5.83)	0.08** (4.14)	0.92 (1.41)	-0.0005 (-0.36)	0.85	6.44[0.17]	9.98[0.99]	2.12[0.15]
Venezuela	0.0004* (1.82)	-0.10 (-0.42)	-0.72 (-0.91)	0.04** (2.45)	0.84	1.55[0.82]	4.21[0.99]	0.15[0.70]

Notes: Dummy variables were excluded from regression for Kuwait because they cause serial correlation, heteroscedasticity and/or misspecification problems. **, * indicate significance at 5% and 10% level respectively. $\chi^2_{sc}(4)$, χ^2_H and $\chi^2_{FF}(1)$ denote chi-squared statistics to test for no residual serial correlation up to four order, homoscedasticity and no functional form misspecification respectively with p-values given in square brackets.

Table 7-18 gives estimates of equilibrium parameters of the RER model using quarterly data and the simple measure of volatility. The results show that the volatility of money supply difference has a significant positive impact on exchange rate volatility in Canada, Germany, Japan, Libya and the UK. The findings also show that the variability of consumption difference is statistically significant only in Algeria, with a positive sign.

The variability of money supply differential is positively significant in Canada, Germany, Japan Libya and the UK. The volatility of the consumption differential is positive significant in Algeria only. Again, except for the constant term, there is no one single significant variable in the Kuwaiti case or in Venezuela for the RER model. In general, all three models so far have a good fit and pass the tests of autocorrelation, heteroscedasticity and misspecification at the 95% significance level, except in small number of cases.

Table 7-18: Estimates of the long run coefficients of the RER model using quarterly data and the SM proxy;

$$vs_t = \gamma_0 + \gamma_1 vm_t + \gamma_2 vc_t + \varepsilon_{2t}$$

Country	γ_0	γ_1	γ_2	$\overline{R}^2$	$\chi^2_{sc}(4)$	χ^2_H	$\chi^2_{FF}(1)$
Algeria	0.0003 (0.78)	-0.56 (-1.23)	0.75** (4.13)	0.59	1.49[0.83]	3.62[0.99]	1.38[0.24]
Canada	0.00007** (4.83)	0.09** (2.08)	0.17 (0.51)	0.61	0.84[0.93]	29.20[0.51]	2.44[0.12]
Germany	0.0001 (0.86)	5.69** (4.41)	-0.33 (-1.02)	0.70	4.95[0.29]	43.61[0.05]	6.74[0.01]
Japan	0.0006** (4.37)	1.05* (1.95)	2.10 (1.41)	0.64	2.43[0.66]	18.89[0.53]	7.86[0.005]
Kuwait	0.00005** (2.05)	-0.002 (-0.32)	0.0001 (0.03)	0.49	4.91[0.30]	3.36[0.99]	3.53[0.06]
Libya	0.00009 (1.50)	0.13** (2.28)	0.002 (0.36)	0.95	3.58[0.47]	8.63[0.98]	2.45[0.12]
UK	0.0004** (6.72)	0.08** (3.87)	-0.35 (-0.56)	0.84	6.12[0.19]	14.75[0.79]	2.18[0.14]
Venezuela	0.002 (0.65)	0.09 (0.25)	0.13 (1.12)	0.86	6.41[0.17]	3.71[0.96]	0.17[0.68]

Notes: Dummy variables were excluded from regression for Kuwait because they cause serial correlation, heteroscedasticity and/or misspecification problems.**, * indicate significance at 5% and 10% level respectively. $\chi^2_{sc}(4)$, χ^2_H and $\chi^2_{FF}(1)$ denote chi-squared statistics to test for no residual serial correlation up to four order, homoscedasticity and no functional form misspecification respectively with p-values given in square brackets.

7.3.3 Results of the bounds testing tests for cointegration using quarterly data and the standard deviation as a volatility measure

Equations (7.4)-(7.6) are now estimated by OLS using quarterly data and the standard deviation as a volatility measure. The maximum lag length will be set at eight quarters for all countries except for Libya and Venezuela where it is set at four lags as a result of the small number of observations. Foe equation (7.4) we compute the F-statistic of the bounds test to find out the existence of cointegrating vector between the exchange rate volatility and volatility of the fundamentals based on variables introduced by the sticky-price monetary model of exchange rate. The results of the bounds F-statistics for cointegration in equation (7.4) (the SP model) are shown in Table 7-19. All F and t-critical values are for an unrestricted intercept and no trend, as there is no trend included in any of the regressions (7.4)-(7.11). The results of the F-statistics reveal that long run relationships exist for all countries between exchange rate variability and variability of the fundamental differentials based on variables suggested by the sticky-price monetary model of exchange rate (SP). The computed t-values confirm the results of F-tests for the presence of such level relationships at 10% or less in all cases, except for the Venezuelan case where the t-value falls in the inconclusive area when considering level of significance of 5% level. Overall, we can conclude that a level relationship is present for the SP in all countries when using the quarterly data set and the standard deviation as a volatility measure.

Table 7-19: Results of the bounds *F*-tests for cointegration analysis of the SP using quarterly data and the SD proxy

Country	*F*-statistic	*t*-statistic	AIC	Lags
Algeria	5.18**	-3.94*	-6.07	1
Canada	4.60**	-4.33**	-8.29	3
Germany	4.89**	-3.99**	-6.94	6
Japan	5.55**	-5.22**	-6.69	1
Kuwait	5.27**	-3.88*	-9.08	1
Libya	7.65**	-4.36**	-7.59	3
UK	7.34**	-5.10**	-6.77	1
Venezuela	11.60**	-3.52inc	-4.36	1

The lower and upper *F*-critical values are 2.86 and 4.01 respectively at 5%, 2.45 and 3.52 respectively at 10% level of significance. The lower and upper *t*-critical values are -2.86 and -3.99 respectively at 5%, -2.57 and -3.66 respectively at 10% level of significance. The period of estimation starts in 1975q3 for all countries except for Libya and Venezuela when it starts in 1986q2 and 1989q2 respectively and ends in 1998q4 for all countries. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. inc refers to that the computed statistic falls within the inconclusive band. **, * refer to significance at 5% and 10% levels respectively.

Table 7-20 displays the results of bounds *F* and *t* cointegration tests for the FP based on the flexible-price monetary model of exchange rate (FP). It is clear that the results for the SP and FP are similar. From Table 7-20 the bounds *F*-statistics indicate the existence of cointegrating relationships in all countries. Moreover, the *t*-values confirm the findings of the *F*-statistics of level relationships at 5% in all eight countries. Therefore, by large we can conclude that cointegrating relationships exist for all countries using the FP model.

Table 7-20: Results of the bounds *F*-tests for cointegration analysis of the FP using quarterly data and the SD proxy

Country	*F*-statistic	*t*-statistic	AIC	Lags
Algeria	5.46**	-4.22**	-6.07	1
Canada	5.35**	-4.50**	-8.29	3
Germany	5.02**	-4.14**	-6.86	6
Japan	7.25**	-5.03**	-6.72	1
Kuwait	6.68**	-3.96**	-9.13	1
Libya	6.29**	-4.35**	-7.30	3
UK	6.58**	-4.82**	-6.78	2
Venezuela	12.07**	-6.62**	-4.16	0

The lower and upper *F*-critical values are 3.23 and 4.35 respectively at 5%, 2.72 and 3.77 respectively at 10% level of significance. The lower and upper *t*-critical values are -2.86 and -3.78 respectively at 5%, -2.57 and -3.46 respectively at 10% level of significance. The period of estimation starts in 1975q3 for all countries except for Libya and Venezuela when it starts in 1986q2 and 1989q2 respectively and ends in 1998q4 for all countries. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. * *refers to significance at 5% level.

The results of the bounds testing *F*-statistics and *t*-statistics for cointegration for the RER model using quarterly data appear in Table 7-21. The computed *F*-statistics indicate that level relationships exist for all countries. The *t*-values support the results of the *F*-statistic test at 5% level in all cases. These findings lead us to suggest that long run relationships do indeed exist between the volatility of the exchange rate and the volatility of the different fundamentals suggested by the RER model using quarterly data for all countries of the sample. Overall, the results from the quarterly data using the standard deviation are slightly different from those using the simple measure as a proxy for variability, in that for the simple measure that level relationships were found to be present for all countries except for Algeria using equations (7.4) and (7.5).

Table 7-21: Results of the bounds F-tests for cointegration analysis of the RER using quarterly data and the SD proxy

Country	F-statistic	t-statistic	AIC	Lags
Algeria	7.08**	-7.66**	-6.75	3
Canada	4.96**	-4.55**	-8.26	4
Germany	10.70**	-5.34**	-6.95	2
Japan	7.02**	-3.54**	-6.66	2
Kuwait	7.66**	-4.75**	-9.17	1
Libya	7.44**	-5.92**	-7.51	3
UK	7.07**	-5.00**	-6.77	2
Venezuela	5.93**	-4.36**	-3.80	2

The lower and upper F-critical values are 3.79 and 4.85 respectively at 5%, 3.17 and 4.14 respectively at 10% level of significance. The lower and upper t-critical values are -2.86 and -3.53 respectively at 5%, -2.57 and -3.21 respectively at 10% level of significance. The period of estimation starts in 1975q3 for all countries except for Libya and Venezuela when it starts in 1986q2 and 1989q2 respectively and ends in 1998q4 for all countries. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. * *refers to significance at 5% level.

7.3.3.1 Results of bounds testing t-tests for long run forcing using quarterly data and the standard deviation

In order to establish whether the regressors based on variables originally proposed by traditional models, were in fact long run forcing, which would hence confirm the uniqueness of the cointegrating relations found, equations (7.4)-(7.6) were re-estimated for each regressor as a dependent variable. Then, the bounds the t-tests were applied to test for the absence of feedback from the exchange rate variability. The results of t-tests for equation (7.4) using the standard deviation for quarterly data as a volatility measure are shown in Table 7-22.

The bounds t-statistic tests reveal that the assumption of unique cointegrating vector in equation (7.4) is accepted in all cases including Venezuela. Thus the bounds testing method will be applied to estimate the long run coefficients of the SP model for all countries.

Table 7-22: Results of long run forcing *t*-tests of the SP using quarterly data and the SD proxy

Country	*vm*	*vy*	*vr*	*vπ*
Algeria	-1.83	-1.78	-1.16	1.06
Canada	0.93	-0.95	0.49	1.34
Germany	0.20	-0.10	0.83 t	0.11
Japan	-0.54	-0.71	0.74	0.23 t
Kuwait	0.68	1.51	-0.67 t	-1.63 t
Libya	-0.09	0.81	-0.80	-1.44
UK	2.34	-0.02	-0.31	0.91
Venezuela	-0.83	-1.62	2.09	2.04

The upper bound of the critical *t*-values at 5% is -3.99 for k=4 and -4.19 for k=5 for unrestricted intercept and no trend, where k is the number of regressors. Upper bound of the critical *t*-values at 5% is -4.36 for k=4 and -4.52 for k=5 for unrestricted intercept and unrestricted trend. The lower *t* critical value at 5% is -2.86 and -3.41 with and without trend respectively. t refers to the inclusion of trend in the regression. ** refers to a significant statistic at 5% level.

The results in Table 7-22 for the SP model using the standard deviation as a proxy for variability for the quarterly data set are generally in line with the results in Table 7-13 for the same model using the simple measure as a proxy for volatility for the same frequency of data, with one exception being Venezuela.

Table 7-23 contains the results of the bounds *t*-tests for the absence of feedback in the FP. It can be noticed that the assumption of one long run relationship cannot be rejected in all countries of the sample. Therefore, we proceed to estimate the long run parameters of the FP model.

Table 7-23: Results of long run forcing *t*-tests of the FP using quarterly data and the SD proxy

Country	*vm*	*Vy*	*vr*
Algeria	-1.44	-1.72	-0.14
Canada	0.89	-1.16	0.58
Germany	-0.24	-0.89	2.26 t
Japan	0.43	-0.47	1.48
Kuwait	-0.49	1.52	-0.59
Libya	-0.30	-0.29	-0.93
UK	1.75	-0.41	0.02 t
Venezuela	-0.21	-1.40	1.95

The upper bound of the *t*-critical values at 5% is -3.78 for k=3 for unrestricted intercept and no trend. The upper bound of the *t*-critical values at 5% is -4.16 for k=3 for unrestricted intercept and unrestricted trend. The lower *t* critical value at 5% is -2.86 and -3.41 with and without trend respectively. t refers to the inclusion of trend in the regression. ** refers to a significant statistic at 5% level.

Table 7-24 shows the results of the *t*-tests for testing the absence of any more cointegrating vectors in the RER for quarterly data using the standard deviation as a measure of fluctuation. The results indicate the validity of the assumption of unique level relationship in all economies considered, confirming the validity of applying the bounds testing approach in estimating the coefficients of long run relationships for all cases of the sample.

Table 7-24: Results of long run forcing *t*-tests of the RER using quarterly data and the SD proxy

Country	*vm*	*vc*
Algeria	-3.25	-0.74
Canada	0.63	2.09
Germany	0.97	1.49 t
Japan	0.75	-0.22
Kuwait	-0.31	-1.10
Libya	0.33	0.08
UK	0.91	1.54
Venezuela	-0.83	-1.50

The upper bound of the *t*-critical values at 5% is -3.53 for k=2 for unrestricted intercept and no trend. The upper bound of the *t*-critical values at 5% is -3.95 for k=2 for unrestricted intercept and unrestricted trend. The lower *t* critical value at 5% is -2.86 and -3.41 with and without trend respectively. t refers to the inclusion of trend in the regression. ** refers to a significant statistic at 5% level.

The *t*-tests of long run forcing indicate the absence of causality from exchange rate volatility towards the regressors in all countries. These results strongly resemble that shown in Table 7-15 for quarterly data using the simple measure of volatility.

Generally speaking the tests of long run forcing for monthly and quarterly data using different measures of volatility did not significantly differ from one another in terms of the countries that have more than one cointegrating relationship and the variables that appear to be cointegrated among themselves.

7.3.3.2 The estimation of long run coefficients using quarterly data and the standard deviation

Using the same method and steps used for monthly and quarterly data for the simple measure, the results of the estimation of long run coefficients of the SP for quarterly data using the standard deviation as proxy for variability are shown in Table 7-25 below.

The results in Table 7-25 show that the coefficient of volatility money supply differential has an inverse relationship with exchange rate volatility in all countries under consideration except for the UK. However, it is only statistically significant in Algeria, Canada, Kuwait, the UK and Venezuela at 5% level. The variability of differenced income is statistically significant at 5% level in all countries except Japan where it is significant at 10% level. It has a positive link with the volatility of exchange rate in all countries except for Algeria and Kuwait. The interest rate difference variability parameter is significant in Algeria, Germany and Japan with a positive sign in Algeria and Japan and a negative sign in Germany. It is not significantly different from zero in the other countries. The variability of inflation differential coefficient is positive and significant only in Germany and is negative and significant in Japan and the UK. It is insignificant at the conventional levels in the other countries.

Table 7-25: Estimates of the long run coefficients of the SP model using quarterly data and the SD;

$$vs_t = \beta_0 + \beta_1 vm_t + \beta_2 vy_t + \beta_3 vr_t + \beta_4 v\pi_t + \varepsilon_{0t}$$

Country	β_0	β_1	β_2	β_3	β_4	$\overline{R}^2$	$\chi^2_{sc}(4)$	χ^2_H	$\chi^2_{FF}(1)$
Algeria	0.01**	-0.28**	-0.08**	0.02**	0.005	0.92	2.19[0.70]	47.64[0.11]	1.09[0.30]
	(6.85)	(-2.42)	(-2.23)	(2.11)	(0.05)				
Canada	0.005**	-0.17**	0.33**	0.008	-0.32	0.72	7.42[0.10]	35.06[0.65]	1.40[0.24]
	(5.66)	(-2.16)	(2.24)	(1.59)	(-0.86)				
Germany	0.004	-0.19	0.97**	-0.01**	2.66**	0.82	2.93[0.57]	39.22[0.84]	0.09[0.76]
	(1.40)	(-1.16)	(3.76)	(-2.11)	(4.42)				
Japan	0.008	-0.06	0.73*	0.09**	-2.12**	0.66	3.38[0.50]	49.45[0.57]	2.23[0.14]
	(1.28)	(-0.49)	(1.73)	(2.39)	(-2.18)				
Kuwait	0.009**	-0.31**	-0.03**	0.03	-0.41	0.66	7.47[0.11]	39.21[0.78]	0.41[0.52]
	(3.13)	(-2.08)	(-1.98)	(0.96)	(-0.89)				
Libya	-0.0005	-0.10	0.23**	0.04	3.48	0.86	6.25[0.18]	45.24[0.12]	2.14[0.14]
	(-0.18)	(-1.65)	(3.24)	(0.72)	(1.11)				
UK	-0.005	1.87**	1.55**	0.003	-6.02**	0.68	4.32[0.36]	58.96[0.75]	3.94[0.05]
	(-0.62)	(2.41)	(3.05)	(0.16)	(-3.22)				
Venezuela	-0.03	-0.76*	1.38**	0.25	1.67	0.97	4.88[0.30]	36.31[0.83]	0.77[0.38]
	(-1.29)	(-1.99)	(2.21)	(1.29)	(0.56)				

Notes: Dummy variables were excluded from regressions for Japan because they caused serial correlation, heteroscedasticity and/or misspecification problems. **, * indicate significance at 5% and 10% level respectively. $\chi^2_{sc}(4)$, χ^2_H and $\chi^2_{FF}(1)$ denote chi-squared statistics to test for no residual serial correlation up to four order, homoscedasticity and no functional form misspecification respectively with p-values given in square brackets.

Table 7-26: Estimates of the long run coefficients of the FP model using quarterly data and the SD;

$$vs_t = \alpha_0 + \alpha_1 vm_t + \alpha_2 vy_t + \alpha_3 vr_t + \varepsilon_{1t}$$

Country	α_0	α_1	α_2	α_3	$\overline{R}^2$	$\chi^2_{sc}(4)$	χ^2_H	$\chi^2_{FF}(1)$
Algeria	0.008** (7.07)	-0.04 (-0.52)	-0.08** (-2.09)	0.02** (2.16)	0.91	3.67[0.45]	23.67[0.74]	0.04[0.84]
Canada	0.005** (5.49)	-0.18** (-2.15)	0.34** (2.07)	0.007 (1.25)	0.66	7.91[0.10]	19.89[0.75]	0.01[0.92]
Germany	0.009** (3.75)	-0.18 (-0.90)	0.59** (2.06)	-0.02** (-2.45)	0.76	1.13[0.89]	36.99[0.73]	1.02[0.31]
Japan	0.001 (0.21)	-0.02 (-0.19)	0.93** (2.50)	0.09** (2.59)	0.64	10.80[0.03]	63.97[0.06]	1.35[0.24]
Kuwait	0.007** (2.90)	-0.28** (-2.03)	-0.02** (-2.05)	0.02 (0.93)	0.64	9.49[0.05]	38.64[0.49]	0.006[0.94]
Libya	0.004 (1.62)	-0.03 (-0.55)	0.12 (1.15)	0.02 (0.57)	0.80	1.02[0.91]	32.90[0.0.005]	0.57[0.45]
UK	-0.004 (-0.71)	1.03** (2.38)	1.28** (3.06)	-0.04 (-1.55)	0.64	0.64[0.96]	54.36[0.16]	2.33[0.13]
Venezuela	0.008 (0.89)	-0.56** (-2.33)	0.48** (2.83)	0.13** (3.90)	0.95	5.14[0.27]	17.60[0.86]	1.85[0.17]

Notes: Dummy variables were excluded from regression for Japan because they caused serial correlation, heteroscedasticity and/or misspecification problems.**, * indicate significance at 5% and 10% level respectively. White method was used to correct for heteroscedasticity in Libya and Newey-west method was used to correct for serial correlation and heteroscedasticity in Japan. $\chi^2_{sc}(4)$, χ^2_H and $\chi^2_{FF}(1)$ denote chi-squared statistics to test for no residual serial correlation up to four order, homoscedasticity and no functional form misspecification respectively with p-values given in square brackets.

The results of the estimation of the FP model shown in Table 7-26 reveal that the variability of money supply differential is negatively significant in Canada, Kuwait and Venezuela and positively significant in the UK. The volatility of income difference has a significant inverse relationship with exchange rate volatility in Algeria and Kuwait and a significant direct relationship in Canada, Germany, the UK and Venezuela. The variability of interest differential is positive and significant in Algeria, Japan and Venezuela and negative and significant in Germany. The results obtained from estimating the FP model closely resemble those obtained from estimating the SP model insofar as they differ in only three parameters, namely, the volatility of money supply differential in Algeria, the volatility of income differential in Libya and the volatility of interest differential in Venezuela.

The results of the estimated coefficients of the RER model are shown in Table 7-27. The coefficient of money stock differential volatility is negative and significant only in Algeria, Canada, Kuwait and Venezuela. It is not significant in the rest of the sample. The parameter of consumption differential variability is positive and significant at 5% level in Algeria and Libya only. The *t*-values of the coefficient show that it is insignificant in the rest of countries.

The fit of the three above regressions is good, with a lowest adjusted squared R of 0.55 and they pass all diagnostic tests.

Table 7-27: Estimates of the long run coefficients of the RER model using quarterly data and the SD;

$$vs_t = \gamma_0 + \gamma_1 vm_t + \gamma_2 vc_t + \varepsilon_{2t}$$

Country	γ_0	γ_1	γ_2	$\overline{R}^2$	$\chi^2_{sc}(4)$	χ^2_H	$\chi^2_{FF}(1)$
Algeria	0.007** (7.07)	-0.42** (-3.60)	0.12** (4.59)	0.92	1.67[0.80]	22.78[0.91]	0.55[0.46]
Canada	0.006** (8.04)	-0.11* (-1.88)	0.0008 (0.06)	0.63	5.38[0.25]	21.74[0.75]	0.29[0.59]
Germany	0.009** (3.03)	0.99 (1.22)	-0.14 (-1.24)	0.70	1.79[0.77]	15.23[0.99]	0.02[0.89]
Japan	0.01** (4.35)	-0.07 (-0.66)	0.02 (0.27)	0.55	2.09[0.72]	27.70[0.19]	0.05[0.82]
Kuwait	0.006** (4.39)	-0.18* (-1.95)	-0.02 (-1.62)	0.59	7.47[0.11]	23.87[0.64]	0.02[0.90]
Libya	0.005** (5.30)	-0.04 (-0.87)	0.06** (2.69)	0.82	0.72[0.95]	17.46[0.42]	1.29[0.26]
UK	0.006 (1.06)	1.08 (1.63)	-0.24 (-1.30)	0.56	3.06[0.55]	22.42[0.95]	0.26[0.61]
Venezuela	0.05** (3.79)	-0.89** (-2.43)	-0.13 (-1.52)	0.86	6.47[0.17]	10.65[0.78]	0.41[0.52]

Notes: Dummy variables were excluded from regression for Japan because they caused serial correlation, heteroscedasticity and/or misspecification problems.**, * indicate significance at 5% and 10% level respectively. $\chi^2_{sc}(4)$, χ^2_H and $\chi^2_{FF}(1)$ denote chi-squared statistics to test for no residual serial correlation up to four order, homoscedasticity and no functional form misspecification respectively with p-values given in square brackets.

7.3.4 Results of the bounds testing tests for cointegration using monthly data and a volatility measure based on ARCH models

The following are the results of bounds testing tests for cointegration in the conventional models using monthly data and a volatility proxy based on the best fit ARCH model. The variables for which optimal ARCH models, as introduced in chapter six, were used to test for the existence of long run relationships.

Table 7-28: Results of the bounds *F*-tests for cointegration analysis of the SP model using monthly data and the ARCH volatility proxy

Country	*F*-statistic	*t*-statistic	AIC	Lags
Kuwait	14.87**	-3.88**	-21.94	6
UK	14.43**	-4.29**	-19.05	2

The lower and upper *F*-critical values are 3.79 and 4.85 respectively at %5, 3.17 and 4.14 respectively at %10 level of significance, the lower and upper *t*-critical values are -2.86 and -3.53 respectively at %5, -2.57 and -3.21 respectively at %10 level of significance when k=2. The lower and upper *F*-critical values are 2.86 and 4.01 respectively at %5, 2.45 and 3.52 respectively at %10 level of significance when, and the lower and upper *t*-critical values are -2.86 and -3.99 respectively at %5, -2.57 and -3.66 respectively at %10 level of significance when k=4. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. ** refers to significance at %5 and %10 levels respectively.

For equation (7.4) the only countries for which fitted ARCH models were found are Kuwait and the UK. The regressors included for the Kuwaiti equation were the volatility of money supply differential and inflation differential.

The F cointegration tests indicate the existence of equilibrium relationships in the SP model using estimated volatility from an ARCH model.

For the FP model the only country for which fitted ARCH model was found is the UK.

Table 7-29: Results of the bounds *F*-tests for cointegration analysis of the FP model using monthly data and the ARCH volatility proxy

Country	*F*-statistic	*t*-statistic	AIC	Lags
UK	17.68**	-4.40**	-19.07	2

The lower and upper *F*-critical values are 3.23 and 4.35 respectively at %5, 2.72 and 3.77 respectively at %10 level of significance, the lower and upper *t*-critical values are -2.86 and -3.78 respectively at %5, -2.57 and -3.46 respectively at %10 level of significance when k=3. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. ** refers to significance at %5 and %10 levels respectively.

The results above show the presence of a cointegrating vector in the FP model for the UK data.

For the RER model an ARCH model was found to fit the data from which volatility measures were estimated three countries. Volatility of money supply difference was suppressed in the Japanese case as no ARCH model fits the series.

The findings shown in the table below indicate the existence of long run relationships in the RER model for these three countries.

Table 7-30: Results of the bounds *F*-tests for cointegration analysis of the RER model using monthly data and the ARCH volatility proxy

Country	*F*-statistic	*t*-statistic	AIC	Lags
Japan	6.58**	-4.77**	-17.80	5
Kuwait	13.01**	-4.11**	-21.93	7
UK	29.43**	-4.85**	-19.08	1

The lower and upper *F*-critical values are 3.79 and 4.85 respectively at %5, 3.17 and 4.14 respectively at %10 level of significance, the lower and upper *t*-critical values are -2.86 and -3.53 respectively at %5, -2.57 and -3.21 respectively at %10 level of significance when k=2. The lower and upper *F*-critical values are 4.94 and 5.73 respectively at %5, 4.04 and 4.78 respectively at %10 level of significance, and the lower and upper *t*-critical values are -2.86 and -3.22 respectively at %5, -2.57 and -2.91 respectively at %10 level of significance when k=1. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. ** refers to significance at %5 and %10 levels respectively.

Clearly, these *F*-test results reveal the existence of level relationships in the cases considered, and the *t*-statistics reinforce these findings. Next the absence of feedback from exchange rate volatility to the regressors was tested for. The results of the tests appear in the following tables.

7.3.4.1 Results of bounds testing *t*-tests for long run forcing using monthly data and the ARCH measure

The bounds testing *t*-statistics for the absence of more than one cointegrating relationship in the SP model give results shown in Table 7-31.

Table 7-31: Results of long run forcing *t*-tests of the SP model using monthly data and the ARCH proxy

Country	vm	vy	vr	$v\pi$
Kuwait	-0.72	-	-	-0.42
UK	-2.30	-0.51	-0.90	-0.33

The upper bound of the critical *t*-values at %5 is -3.99 for k=4, and upper bound of the critical *t*-values at %5 is -3.53 for k=2 for unrestricted intercept and no trend. The lower *t* critical value at %5 is -2.86.

The findings show that the assumption of a unique level relationship is accepted at the %5 level in both cases for the SP model.

Table 7-32: Results of long run forcing *t*-tests of the FP model using monthly data and the ARCH proxy

Country	vm	vy	vr
UK	-1.79	0.04	-0.58

The upper bound of the critical *t*-values at %5 level is -3.78 for k=3 for unrestricted intercept and no trend. The lower *t* critical value at %5 is -2.86.

From the above table no variable was found to be significant, which confirms the absence of feedback in the FP for the UK using monthly data and a volatility measure estimated from ARCH model.

The results of long run forcing *t*-statistics reveal the absence of feedback in the RER model for the three considered cases.

Table 7-33: Results of long run forcing *t*-tests of the RER model using monthly data and the ARCH proxy

Country	*vm*	*vc*
Japan	-	-1.08
Kuwait	-0.86	-1.34
UK	-2.22	1.16

The upper bound of the critical *t*-values at %5 level is -3.53 for k=2, and upper bound of the critical *t*-values at %5 is -3.22 for k=1 for an unrestricted intercept and no trend. The lower *t* critical value at %5 is -2.86.

Clearly it can be noted from the tables above that all of the computed *t*-values fall below the lower critical value, confirming the absence of feedback from exchange rate volatility using monthly data and a volatility measure based on estimation by means of ARCH models. Thus we proceed further to calculate the long run coefficients.

7.3.4.2 The estimation of long run coefficients using monthly data and the ARCH measure

Using monthly data and a volatility measure estimated from ARCH models, the long run parameters are shown in Table 7-34.

Table 7-34: Estimates of the long run coefficients of the SP model using monthly data and the ARCH proxy

Country	β_0	β_1	β_2	β_3	β_4	$\overline{R}^2$	$\chi^2_{sc}(4)$	χ^2_H	$\chi^2_{FF}(1)$
Kuwait	0.00001**	-0.012	-	-	0.22	0.57	[0.54]	[0.02]	[0.58]
	(2.23)	(-1.61)			(0.50)				
UK	0.0002**	-0.41	0.99	0.00001	-	0.54	[0.003]	[0.04]	[0.001]
	(4.59)	(-0.45)	(1.57)	(0.86)	9.14**				
					(-2.60)				

Notes: The White method was used to correct for heteroscedasticity in Kuwait and the Newey-West method was used to correct for serial correlation and heteroscedasticity in the UK. **, * indicate significance at %5 and %10 level respectively. $\chi^2_{sc}(4)$, χ^2_H and $\chi^2_{FF}(1)$ denote p-values of the chi-squared statistics to test for no residual serial correlation up to four order, homoscedasticity and no functional form misspecification respectively.

From the table above one can see that the only significant variable is the variability of inflation differential in the UK, which has a negative sign.

Table 7-35: Estimates of the long run coefficients of the FP model using monthly data and the ARCH proxy

Country	α_0	α_1	α_2	α_3	$\overline{R}^2$	$\chi^2_{sc}(4)$	χ^2_H	$\chi^2_{FF}(1)$
UK	0.0002**	0.007	0.36	0.000005	0.51	[0.000]	[0.02]	[0.04]
	(4.38)	(0.008)	(0.79)	(0.35)				

Notes: The Newey-West method was used to correct for serial correlation and heteroscedasticity in the UK. **, * indicate significance at %5 and %10 level respectively. $\chi^2_{sc}(4)$, χ^2_H and $\chi^2_{FF}(1)$ denote p-values of the chi-squared statistics to test for no residual serial correlation up to four order, homoscedasticity and no functional form misspecification respectively.

None of the variables in the above equation seems to be significant. This can also be attributed to the multicollinearity problem.

Table 7-36: Estimates of the long run coefficients of the RER model using monthly data and the ARCH proxy

Country	γ_0	γ_1	γ_2	$\overline{R}^2$	$\chi^2_{sc}(4)$	χ^2_H	$\chi^2_{FF}(1)$
Japan	0.0003**	-	-0.07	0.28	[0.87]	[0.42]	[0.52]
	(6.70)		(-0.85)				
Kuwait	0.00003**	-0.005	-0.004*	0.57	[0.39]	[0.11]	[0.17]
	(4.24)	(-0.68)	(-1.81)				
UK	0.0002**	-1.85**	-0.008	0.49	[0.005]	[0.015]	[0.34]
	(7.43)	(-1.99)	(-0.51)				

Notes: The Newey-West method was used to correct for serial correlation and heteroscedasticity in the UK. **, * indicate significance at %5 and %10 level respectively. $\chi^2_{sc}(4)$, χ^2_H and $\chi^2_{FF}(1)$ denote p-values of the chi-squared statistics to test for no residual serial correlation up to four order, homoscedasticity and no functional form misspecification respectively.

The results in the above table indicate that the volatility of consumption differential in Kuwait is significant and has negative sign. The variability of money supply difference is significant with a negative sign in the UK. The volatility of consumption difference in Japan is insignificant.

7.3.5 Results of the bounds testing tests for cointegration for Devereaux-Lane's model using annual data and the standard deviation measure

Having completed the cointegration tests using the traditional-based volatility models of exchange rate, we turn to investigate the existence of long run relationships using Devereux and Lane (Devereux-Lane) model. First equation (4.9) from chapter four is rewritten in the following simple form:

$$vs_t = b_0 + b_1 trade_t + b_2 cycle_t + b_3 size_t + b_4 finance_t + b_5 extfin_t + u_t \quad (7.7)$$

The ARDL-ECM form of Devereux and Lane model, equation (7.7), can be written as follows:

$$\Delta vs_t = b_0 + b_1 t + \eta_1 vs_{t-1} + \eta_2 trade_{t-1} + \eta_3 cycle_{t-1} + \eta_4 size_{t-1} + \eta_5 finance_{t-1} + \eta_6 extfin_{t-1} + \sum_{j=1}^{m1} \alpha_j \Delta vs_{t-j} + \sum_{j=0}^{m2} \beta_j \Delta trade_{t-j} + \sum_{j=0}^{m3} \gamma_j \Delta cycle_{t-j} + \sum_{j=0}^{m4} \lambda_j \Delta size_{t-sj} + \sum_{j=0}^{m5} \delta_j \Delta finance_{t-j} + \sum_{j=0}^{m6} \phi_j \Delta extfin_{t-j} + \varphi D_t + \xi_t \quad (7.8)$$

where *vs* is the volatility of exchange rate, *trade* is the trade linkages, *cycle* is the asymmetric shock, *size* is the economy size, *finance* is the internal financial development, and *extfin* is the external financial dependence. The rest of the terms are as previously defined. The *extfin* variable, is only included for Algeria and Venezuela since the other countries have no significant amounts of external debt. In addition data could not be found for this variable for the rest of countries.

As we are using annual data for this model, and following Narayan and Smyth (2005) and Tang (2003) the maximum lag length will be set at 2 for all countries. A higher number of lags is usually not feasible given that the present study had 26 annual observations. In fact several previous studies have used the bounds testing approach with relatively small sample sizes. For example, Pattichis (1999) used the ARDL-ECM model to a sample size of 20 observations (1975-1994), Tang (2001) employed the bounds methodology for a period from 1973 to 1997 using annual data, and Tang (2002) used annual data for 1973-1998 for the money demand function in Malaysia.

The null hypothesis of no cointegration is tested by estimating equation (7.8) without the lagged levels in order to test the joint significance of these lagged levels, i.e. $H_0 : \eta_1 = \eta_2 = \eta_3 = \eta_4 = \eta_5 = \eta_6 = 0$ against H_1 : at least one of $\eta's \neq 0$.

All models were estimated by OLS and subjected to a number of diagnostic tests. In these tests we relied on both the Lagrange multiplier (LM) version and the *F*-version of the tests, since Pesaran and Pesaran (1997) noted that both versions have the same distribution for large samples. However, the *F*-version is generally preferable to the LM in small samples on the basis of Monte Carlo results (Pattichis, 1999).

The results of the bounds *F*-statistic tests of cointegration for the Devereaux-Lane model using annual data are shown in Table 7-37. The *F*-tests of cointegration indicate the presence of cointegration relationships only in the Algerian and Venezuelan cases, although the *t*-values in these countries do not support the finding from the *F*-statistic test. The null hypothesis that there exist no level relationships cannot be rejected at the 5% level of significance in the rest of the sample since the computed *F*-statistics fall below the lower critical value of 2.86 at 5% level except in the Libyan case where the computed *F*-value falls within the inconclusive area.

Table 7-37: Results of the bounds *F*-tests for cointegration analysis of the Devereux-Lane model

Country	*F*-statistic	Upper *F*-cv	*t*-statistic	Upper *t*-cv	AIC	Lags
Algeria	3.82**	3.79	-2.26	-4.19	-7.52	1
Canada	0.88	4.01	-4.43**	-3.99	-9.62	1
Germany	1.72	4.57	-7.01**	-4.36	-10.67	2
Japan	1.42	4.01	-1.70	-3.99	-7.18	1
Kuwait	0.54	4.01	-2.70	-3.99	-9.70	1
Libya	3.22inc	4.01	-1.36	-3.99	-7.11	0
UK	0.87	4.01	-1.52	-3.99	-7.42	1
Venezuela	11.40**	3.79	-1.56	-4.19	-5.08	0

The upper F and t critical values shown in the table are for the 5% level of significance. The period of estimation starts in 1973 for all countries except for Libya and Venezuela when it starts in 1986 and 1983 respectively and ends in 1998 for all countries. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. Note that the critical values of F and t-statistics for Algeria, Germany and Venezuela differ from those of other countries as the proxy for the external debt was included only in Algeria and Venezuela, and a linear trend is included for the German case. **, * indicate significance at 5% and 10% levels respectively.

To test for the absence of feedback from the exchange rate volatility in Algeria and Venezuela, equation (7.8) is re-estimated taking each of the variables in turn as a dependent variable. The null hypothesis, that the coefficient of the lagged level exchange rate variability is not different from zero, is then performed depending upon the t-tests. If the null cannot be rejected, then the variables on the right hand side of the model considered are confirmed to be long run forcing regressors for vs_t. The results of the t-tests are displayed in Table 7-38.

As can be seen from Table 7-38, both t-values fall below the lower critical value at 5% level of significance. This implies that the assumption of unique cointegration relationship assumed by the bounds methodology cannot be rejected for Algeria and Venezuela.

Table 7-38: Results of long run forcing *t*-tests of the Devereux-Lane model

Country	*cycle*	*trade*	*finance*	*size*	*extfin*
Algeria	0.68	-0.99	-2.36	-0.98 t	0.37 t
Venezuela	0.58	0.28	-1.06	-0.38	-0.86

The upper bound of the critical *t*-values at 5% is -4.19 for k=5 for unrestricted intercept and no trend. The upper bound of the critical *t*-values at 5% is -4.52 for k=5 for unrestricted intercept and unrestricted trend. The lower *t* critical value at 5% is -2.86 and -3.41 with and without trend respectively. t refers to the inclusion of trend in the regression.

Thus, the bounds testing method is appropriate for use in estimating the long run coefficients.

The results of the estimation of long run coefficients of the relationship between the volatility of exchange rate and the regressors introduced by the Devereux-Lane model using the UECM are shown in Table 7-39 below.

The results reveal that the coefficients of *trade* and *finance* are statistically significantly different from zero in the Algerian case at 10% level of significance or less, although the sign of *trade* does not comply with expectations. The rest of the variables do not seem to be significant, although their signs conform to theoretical expectations.

In the Venezuelan case, however, there are three significant variables: *trade, finance* and *extfin*. *Cycle* and *size* are statistically insignificant in both countries. Interestingly, the signs of the parameters in Venezuela entirely conform to expectations. Also; the signs in the Algerian equation are in line with the theoretical expectations except for the *trade* variable. The model above has a very good fit and does not suffer from autocorrelation, heteroscedasticity or misspecification.

Table 7-39: Estimates of the long run coefficients of the Devereux-Lane model

Country	*constant*	*cycle*	*trade*	*finance*	*size*	*extfin*	$\overline{R}^2$	$\chi^2_{sc}(1)$	χ^2_H	$\chi^2_{FF}(1)$
Algeria	0.02	0.11	0.01*	-0.01**	0.003	-0.0003	0.91	1.60[0.23]	22.99[0.40]	1.98[0.19]
	(0.67)	(1.28)	(2.05)	(-2.41)	(0.40)	(-0.80)				
Venezuela	0.02	3.19	-0.22*	-0.31**	0.04	-9.5×10^{12}*	0.94	4.54[0.09]	14.19[0.44]	0.51[0.51]
	(0.05)	(1.27)	(-1.98)	(-3.39)	(0.43)	(-2.36)				

Notes: A dummy variable was included in the regressions for Algeria to take into account the effect of outlier in 1991. The Newey-West method was used to correct for serial correlation and heteroscedasticity in Venezuela. **, * indicate significance at 5% and 10% levels of significance respectively. The model passes a battery of diagnostic tests. $\chi^2_{sc}(1)$, χ^2_H and $\chi^2_{FF}(1)$ denote the *F* version statistics to test for no residual serial correlation, homoscedasticity and no functional form misspecification respectively with p-values given in square brackets.

These findings indicate that the Devereux-Lane model does not perform well for developed economies. Furthermore, it works only for two out of four cases for developing countries.

7.3.6 Results of the bounds testing tests of cointegration for the hybrid model using annual data and the standard deviation

It may now be appropriate to estimate a new model consisting of the Devereux-Lane model and the traditional-based volatility models. Equations (4.10)-(4.12) in chapter four represent such an augmented model. Equation (4.10) incorporates the Devereux-Lane model with the flexible-price monetary-based volatility model, excluding variables that are likely to cause multicollinearity since similar variables already exist in Devereux-Lane model. Equations (4.11) and (4.12) represent hybrid equations which emerge as a result of incorporating Devereux-Lane model into the sticky-price monetary-based volatility model and Redux-based volatility model respectively.

Performing the bounds testing approach to the hybrid models produced the results shown in Table 7-40. This test was applied to all countries except Libya and Venezuela where it was impossible to run regressions with such large numbers of regressors and small numbers of observations. The results of the bounds *F*-tests do not show the existence of any cointegration relationships in any of the countries considered.

The results of the bounds testing method for the sticky-price-based volatility Devereux-Lane model are illustrated in Table 7-41. No single cointegration can be found in this model according to the bounds tests.

Table 7-40: Results of the bounds F-tests for cointegration analysis of the hybrid flexible-price-based volatility-Devereux-Lane model using annual data

Country	F-statistic	Upper F-cv	t-statistic	Upper t-cv	AIC	Lags
Algeria	2.09	3.50	-6.08	-4.57	-7.75	0
Canada	1.68	3.61	-1.52	-4.38	-9.58	0
Germany	1.27	4.00	-3.88	-4.69	-10.08	1
Japan	0.79	3.61	-2.18	-4.38	-7.99	1
Kuwait	0.22	3.61	-5.42	-4.38	-10.90	1
UK	1.36	3.61	-2.95	-4.38	-10.47	1

The upper F and t critical values are for the 5% level of significance. The period of estimation starts in 1973 for all countries. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. Note that the critical values of F and t-statistics for Algeria and Germany differ from those of other countries as the proxy for the external debt was included for Algeria, and a linear trend is included for the German case. **, * indicate significance at 5% and 10% levels respectively.

Table 7-41: Results of the bounds F-tests for cointegration analysis of the hybrid sticky-price-based volatility-Devereux-Lane model using annual data

Country	F-statistic	Upper F-cv	t-statistic	Upper t-cv	AIC	Lags
Algeria	1.76	3.39	-4.39	-4.72	-7.70	0
Canada	2.00	3.50	-1.91	-4.57	-9.63	0
Germany	2.29	3.83	-4.48	-4.85	-8.07	0
Japan	2.79	3.50	-1.93	-4.57	-7.34	0
Kuwait	0.69	3.50	-2.14	-4.57	-9.21	0
UK	0.38	3.50	-3.42	-4.57	-7.44	0

The upper F and t critical values are for the 5% level of significance. The period of estimation starts in 1973 for all countries. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. Note that the critical values of F and t-statistics for Algeria and Germany differ from those of other countries as the proxy for the external debt was included for Algeria, and a linear trend is included for the German case. **, * indicate significance at 5% and 10% levels respectively.

Tests of cointegration for the Redux-based volatility Devereux-Lane model can be seen in Table 7-42.

Table 7-42: Results of the bounds F-tests for cointegration analysis of the hybrid Redux-based volatility-Devereux-Lane model using annual data

Country	F-statistic	Upper F-cv	t-statistic	Upper t-cv	AIC	Lags
Algeria	2.47	3.50	-10.63	-4.57	-8.79	0
Canada	3.37	3.61	-0.97	-4.38	-10.69	1
Germany	0.80	4.00	-2.23	-4.69	-8.78	1
Japan	1.10	3.61	-3.07	-4.38	-7.10	0
Kuwait	1.12	3.61	-3.60	-4.38	-9.86	0
UK	0.03	3.61	-3.51	-4.38	-7.32	0

The upper F and t critical values are for the 5% level of significance. The period of estimation starts in 1973 for all countries. Diagnostic tests indicate the absence of serial correlation, heteroscedasticity and misspecification problems. Note that the critical values of F and t-statistics for Algeria and Germany differ from those of other countries as the proxy for the external debt was included for Algeria, and a linear trend was included for the German case. **, * indicate significance at 5% and 10% levels respectively.

Again no evidence was found for the existence of cointegration using the Redux-based volatility Devereux-Lane model.

The findings reveal that the hybrid models do not contain cointegrating relationships, as the computed F-values fall below the lower critical value at 5% level of significance in all countries. Note that no long run relationship was found in the case of Algeria, which contrasts with the results obtained from the Devereux-Lane model alone. This can be explained by the over-parameterization problem which may lead to inefficient results. Thus we proceeded no further with these augmented models.

The integration of some variables from the traditional-based volatility models into the Devereux-Lane model may cause an inefficiency problem as a consequence of the low number of observations, which means exhausting degrees of freedom. In order to reduce the effects of this possible difficulty, a single variable can be inserted from the conventional-based volatility models into the Devereux-Lane model each time to save degrees of freedom for all countries including Libya and Venezuela. The results of such experiments are not reported here to save space, but yielded no major changes compared to those shown in Tables (7-40)-(7-42) except for the case of Algeria where evidence was found of a long run relationship when the volatility of consumption differential from the Redux-based volatility model was included in the Devereux-Lane equation. Feedback from exchange rate volatility to the regressors of this hybrid model was tested for and the findings showed its absence. Therefore, the estimated long run coefficients using the UECM for this augmented equation are reported in

Table 7-43. Surprisingly, all regressors, except for *cycle*, are now highly significant and have the correct expected signs, except for *trade*.

The diagnostic tests show the absence of serial correlation, heteroscedasticity and misspecification.

This chapter has presented the results of the empirical experiments with the volatility models developed earlier in chapter four. These empirical results include the bounds testing tests for cointegrating relationships, the bounds testing tests for the presence of more than one long run relationship, and finally the estimated long run coefficients for all volatility models introduced. The following chapter contains comparisons between the results presented in this chapter to investigate weather or not the findings are sensitive to the different models used, variability proxies and data frequencies.

Table 7-43: Estimates of the long run coefficients of the Devereux-Lane equation augmented with the volatility of consumption differential for Algeria

Country	*constant*	*cycle*	*trade*	*finance*	*size*	*extfin*	*vc*	$\bar{R}^2$	$\chi^2_{sc}(1)$	χ^2_H	$\chi^2_{FF}(1)$
Algeria	-0.06**	0.006	0.02**	-0.02**	0.02**	-0.001**	0.27**	0.99	2.26[0.17]	21.28[0.36]	2.14[0.18]
	(-2.98)	(0.17)	(5.16)	(-8.34)	(4.53)	(-7.19)	(6.79)				

Notes: A dummy variable was included in the regression to take into account the effect of outlier in 1991. **, * indicate significance at 5% and 10% levels of significance respectively. The model passes a battery of diagnostic tests. $\chi^2_{sc}(1)$, χ^2_H and $\chi^2_{FF}(1)$ denote the *F* version statistics to test for no residual serial correlation, homoscedasticity and no functional form misspecification respectively with p-values given in square brackets.

Chapter eight

Comparisons

8.1 Introduction

The previous chapter introduced the results of the empirical research which involved a variety of models, volatility measures, data frequencies and countries. In the first stage of the empirical analysis evidence was found of the existence of cointegrating relationships in most cases. The assumption of unique long run relationships was globally accepted in the second stage of the bounds testing procedure. In the third step the long run coefficients have been estimated of the unique long run relationships established. Since one of the aims of this thesis is to compare and contrast the results of the empirical analysis obtained using different models, volatility measures and data frequencies, the current chapter makes comparisons between different tests of results. The comparisons will be particularly assigned for the results of the estimated long run parameters. Firstly, the results obtained using different models are compared and contrasted; i.e. comparing the results of the traditional-based volatility models (SP, FP and RER) for each country in the sample using the same volatility measure and data frequency. The performance of these different conventional-based models is assessed using several criteria for model selection. Thus, we may be able to conclude that a certain model is preferred for a certain country under certain conditions. Moreover, we may find that a specific model performs better for a specific group of countries. Secondly, the findings obtained are compared using different volatility proxies using the same model and data frequency. These differences in the results may be discovered using these measures, and explanations are offered for the sources of any such differences. The aim is to discover whether the results are sensitive to the variability measure used and/or that each proxy measures a different sort of volatility. Thirdly, the results obtained by utilizing different data frequencies under the same model and variability proxy are compared, to discover any effect of changing the frequency of the data used on the results. Finally, using the preferred models chosen in section 8.2, as well as the results of Devereux-Lane model and the augmented Devereux-Lane model, the estimated long run coefficients and their policy implications are given. Accordingly, the following section compares the performance of the different models.

8.2 Comparison of the models

A comparison of the performance of the models used in the study can show which model is preferred for each country. The first comparison involves the traditional-based volatility models when using the same measure of volatility and frequency of data. Therefore, the Devereux-Lane model is excluded as it uses different frequency of data. The comparison is conducted to select the most appropriate model in explaining changes in the dependent variable, which entails the use of some criteria by which to distinguish between models. Before introducing the model selection criteria, nested and non-nested models first need to be distinguished. In order to distinguish these consider our SP and FP models; i.e. equations (5-6) and (5-7) respectively in chapter 5 which are repeated here for convenience.

$$vs_t = \beta_0 + \beta_1 vm_t + \beta_2 vy_t + \beta_3 vr_t + \beta_4 v\pi_t + \varepsilon_{0t}, \qquad \text{(SP)}$$

$$vs_t = \alpha_0 + \alpha_1 vm_t + \alpha_2 vy_t + \alpha_3 vr_t + \varepsilon_{1t} \qquad \text{(FP)}$$

We can say that the FP is nested in the SP model because it is a special case of the SP. More precisely, when $\beta_4 = 0$, SP reduces to FP.

The selection criterion for the nested models is straightforward; that is, the normal *t*-test. In order to choose between these two models one should examine the significance of the inflation differential variability term that is used to test the null hypothesis of $\beta_4 = 0$ in the SP. The SP model is rejected in favour of the FP model, if the coefficient of the inflation differential volatility is insignificant according to the *t*-test. In contrast, the FP hypothesis is rejected in favour of the SP hypothesis, if the parameter of inflation difference volatility is statistically significantly different from zero.

On the other hand in non-nested models one cannot be derived as a special case of the other. Consider, for instance, models (5-7) and (5-8), namely, the FP and RER in our study. For convenience we rewrite equation (5-8) here.

$$vs_t = \gamma_0 + \gamma_1 vm_t + \gamma_2 vc_t + \varepsilon_{2t} \qquad \text{(RER)}$$

The FP and RER represent different theories that explain the behaviour of the same time series. In this example one cannot consider the RER as a special case of the FP although both contain the volatility of money supply differential. Each model reflects the views of different and competing hypotheses.

Several tests for choosing the appropriate non-nested model have been proposed, such as the Cox-test and the j-test. We apply the j-test introduced by Davidson and MacKinnon (1981), because it is intuitively appealing.

To illustrate the j-test, assume that we want to compare the FP and RER. Firstly, we estimate the FP from which we obtain fitted values. Secondly, we add these predicted values as an additional regressor in the RER. Thirdly, we estimate the augmented

RER model and test the null hypothesis that the additional variable in the augmented RER are statistically insignificant. If this hypothesis cannot be rejected, we can accept the RER as the appropriate model over the FP because the fitted values added into the RER representing the influence of variables not included in the RER have no additional explanatory power beyond those contributed by the RER. In other words, the FP model does not contain any additional information that will improve the performance of the RER model. In contrast, if the null hypothesis is rejected, the RER model cannot be the true model.

Now we reverse the roles of the models FP and RER and perform the same steps to find out whether the FP is accepted over the RER or if it is rejected in favour of the RER model.

Despite its attractiveness, the j-test has two main shortcomings. The first problem with the j-test is when it rejects or accepts both models. When both hypotheses are rejected, neither helps in explaining the behaviour of the dependent variable. If both equations are accepted, the data are apparently not rich enough to discriminate between these models (Gujarati, 2003). To overcome this difficulty, some other selection criteria are used such as the adjusted R squared ($\overline{R}^2$) and the Akaike information criterion (AIC) by choosing the model that has highest $\overline{R}^2$ and lowest AIC. The second problem with the j-test is that it may not be very reliable in small samples because it tends to reject the true equation more frequently than it should to. Since this test is conducted mainly on monthly and quarterly data sets, we may face the problem of small sample size in only a few cases. Thus, the j-test seems to be an appropriate choice.

One should bear in mind that there is no conclusive criterion which can give us a clear answer as to the best model. As Griffiths, et al (1993) stated, "*selecting the* *appropriate set of regressors, and an appropriate model, are difficult problems for which no satisfactory solution exists. It is therefore, impossible to give a prescription that should be followed at all times. There is no clear-cut definitive method for deciding on the best set of variables*". Thus, the criteria described above can only be considered as an indicative guide of when one set of regressors is preferable to another set.

In the comparisons between the conventional-based variability models, we begin with a comparison between the nested hypotheses; i.e. the SP and FP models, using the standard *t*-test to select the more appropriate model. Having chosen between the two nested models, that model chosen is then compared with the RER model using the j-test.

8.2.1 Comparing the conventional-based volatility models using monthly data and the simple measure

The results of model selection using monthly data and the simple measure of volatility are shown in Table 8-1. From this table we can see that the FP equation was chosen over the SP equation since the volatility of the inflation differential is not significant in Algeria. Then the j-test was performed to choose between the FP and RER models. The RER hypothesis was accepted and the FP model was rejected by the j-test. Thus, we conclude that the RER equation is the most appropriate model to explain the changes in the exchange rate volatility in Algeria when monthly data and a simple variability measure are used. In the Canadian case, the FP equation was firstly selected over the SP model. The j-test was hence applied to choose between the FP and RER models. Unfortunately the j-test rejected both hypotheses. Therefore, the adjusted R squared and AIC criteria indicated the preference of the FP equation.

Following the same procedure with all countries, the chosen models are shown in column 5 in Table 8-1. Here there are a few cases which need further discussion in terms of choosing the suitable model. In the case of Germany the differenced inflation variability has a positive sign and is significant at the 5% level, which implies that the SP is better than the FP for Germany. However, the FP and RER gave a positive significant parameter of the money supply differential volatility with a magnitude approaching one, which complies with the expectations[17], whereas in the SP this coefficient is positive and significant at the 10% level but with a magnitude of 0.28 which is quite far from that anticipated. Therefore, it seems unclear which model should be considered as better.

Table 8-1: Results of the preferred models by model selection criteria using monthly data and the SM measure

Country	Chosen from nested models	j-test	$\overline{R}^2$ and AIC	Chosen model(s)
Algeria	FP	Accept RER, reject FP	-	RER
Canada	FP	Reject FP, reject RER	FP	FP
Germany	FP/SP	Reject SP, reject RER	SP	SP/FP/RER
Japan	FP	Reject FP, reject RER	FP	FP
Kuwait	-	-	-	-
Libya	FP	Accept FP, reject RER	-	FP
UK	FP/SP	Accept FP, reject RER	-	FP/SP
Venezuela#	SP	Accept SP, reject RER	-	SP

Note: # indicates that the bounds F-tests have found more than one cointegrating relationship in Venezuela for the SP and FP models.

For the UK, the FP was chosen over the SP since the coefficient of the inflation differential volatility is insignificant. However, the SP has a positive and significant

[17] In the traditional models, which are based on the familiar Cagan-style money demand function, it is assumed that money is homogeneous of degree 1 in prices. This, in turn, implies that exchange rate is homogeneous of degree 1 in money supply differential, as indicated in chapter 2.

parameter for the money supply difference with a magnitude of 0.66 which is closer to theoretical expectations than that of 1.83 from the FP model. Thus, the most suitable model cannot definitely be selected.

Although the model selection criteria show that the RER is preferred for Kuwait, none of the alternative models gives any significant coefficient. Thus, one would say that none of the models can be considered as appropriate for the Kuwaiti data when using monthly frequency and the simple measure.

Overall the RER model was selected as the most appropriate model using monthly data and a simple volatility measure in Algeria and Germany only. Choosing the RER for Algeria is not surprising, since the FP and SP models are likely to suffer from multicollinearity problem as indicated in chapter seven. The SP model was found to be the preferred model in Germany, the UK and Venezuela. However, the FP model was found to be the best for Canada, Germany, Japan, Libya and the UK. Moreover, it may be noted that the FP is the most appropriate model for most or all of the developed countries in the sample. Theoretically it is said that the FP model is a long run model as it assumes flexible prices and the PPP holds continuously. Thus, one would expect this model to perform better for low frequency data. However, our empirical results surprisingly show that the volatility models based on the flexible price exchange rate monetary model works better for monthly data which is effectively short term data. In fact our results are in line with those of Sarmas (1996) who found that the flexible-price model has superior performance over the sticky-price model using monthly data for the dollar/pound, dollar/mark and dollar/yen.

The story for the developing countries is a rather different, as the RER was selected for Algeria and Kuwait, the FP for Libya and the SP for Venezuela. No single model seems to fit most of these countries. This may be due to the fact that underdeveloped countries vary widely in terms of their characteristics and each country has its own specific features. Moreover, the exchange rate regimes followed by these countries witnessed changes during the sample period. These differences in such economies may make it difficult to group them into one category or to find a specific model that suits most of them. However, one should bear in mind that the results in Algeria using the FP and SP and in Kuwait using all models suffer from multicollinearity problems. Therefore, one should be cautious when interpreting the results of these two countries. On the other hand, since the developed countries in our sample allow their currencies to freely float in the exchange markets, it was easier to group them and to find one model that fitted most or all of them.

8.2.2 Comparing the conventional-based volatility models using quarterly data and the simple measure

The results of selecting the appropriate model using quarterly data and a simple measure for the volatility are included in Table 8-2 below. Using the FP and SP, the bounds testing approach did not find a long run relationship in Algeria; therefore, the RER model was the only estimated model using quarterly data and a simple volatility measure.

As far as Canada is concerned, the coefficient of inflation differential variability is positive and significant as is the interest rate difference volatility which suggests that the SP is preferred to the FP. However, the FP contains two significant variables as well, namely the volatilities of money supply and interest rate differentials. Thus, either can be selected as the most appropriate compared to the RER model according to adjusted squared R and the AIC criteria.

For Germany, the use of $\overline{R}^2$ and AIC meant that the RER was selected over the SP. However, from the estimated long run coefficients in Table 7-16 and Table 7-18 the SP has three significant parameters at 10% level of significance or less compared to only one significant coefficient in the RER. Thus the SP was considered as the most appropriate model for this country.

Although the RER is rejected by the j-test in favour of the SP in Japan, it contains a positive coefficient for the volatility of money supply differential which is significant at 10% with a magnitude of one, which coincides with expectations. On the other hand, the SP has two significant parameters at 10% or less for the volatility of differenced interest rate and inflation. So the best model for Japan remains undecided.

Again, as in the monthly data, no model has a single significant coefficient for Kuwait, although the FP model preferred according the model selection criteria.

Table 8-2: Results of the preferred models by model selection criteria using quarterly data and the SM measure

Country	Chosen from nested models	j-test	$\overline{R}^2$ and AIC	Chosen model(s)
Algeria	_	_	_	RER
Canada	SP	Reject SP, reject RER	SP	SP/FP
Germany	SP	Reject SP, reject RER	RER	SP
Japan	SP	Reject RER, accept SP	_	SP/RER
Kuwait	-	-	-	-
Libya*	SP	Reject SP, reject RER	RER	SP
UK	FP	Reject FP, reject RER	SP	SP
Venezuela*#	SP	Reject SP, reject RER	SP	SP

Notes: * indicates caution should be taken when interpreting results of j-test for Libya and Venezuela as the sample size is quite small; i.e. 50 and 40 observations included for both countries respectively. # indicates that the bounds *F*-tests found more than one cointegrating relationships in Venezuela for the SP and FP models.

Therefore, using the simple measure for both monthly and quarterly observations, we cannot find any preferred model for the Kuwaiti data among from the competing theories.

Furthermore, with regard to the Libyan case, although the use of $\overline{R}^2$ and AIC led to the selection of the RER over the SP, the latter gave three significant regressors at 10% compared to only one significant variable from the RER. Therefore, the SP is chosen as the most appropriate model. As far as the UK is concerned, the FP was selected over the SP since the coefficient of inflation differential variability is not significantly different from zero. However, the SP has two significant regressors compared to only one in the FP. Moreover, the use of $\overline{R}^2$ and AIC indicated the selection of the SP over the other models. For these reasons, the SP is chosen over the other models.

Overall, the RER model was chosen only in Libya, Japan in addition to Algeria, where there was no other model to be compared with. The SP on the other hand was found to be the preferred model for Canada, Germany, Japan, Libya, the UK and Venezuela. Surprisingly, the SP is now considered to be the most suitable model for the developed countries; i.e. Canada, Germany, Japan and the UK. Note that the FP was chosen as the most appropriate for these countries when using monthly data. Theoretically, one would expect the SP to work better for more frequent data and the FP to perform better for less frequent data, as prices are supposed to be less flexible in the short term and more flexible in the longer term. However, the results discussed thus far reveal that the FP works better for monthly data and the SP works better for less frequent data. In fact one cannot consider these results as a completely contradicting theoretical anticipation, since quarterly observations still represent the short term as prices can still be considered sticky within one quarter. Therefore, choosing the SP model as the most appropriate in developed countries using quarterly observations remain theoretically acceptable. On the other hand, for less developed countries it is difficult to classify a specific model as best for these countries. This is because in Algeria we have only one estimated model, which is the RER. Then in Kuwait, despite the criteria used indicating the selection of the FP, no single significant regressor appears in all models including the FP per se. Apparently this is a consequence of multicollinearity problem. On the other hand, the SP was the most appropriate model for both Libya and Venezuela. Thus, we could generally say that the SP model is the most suitable for both developed and underdeveloped economies. This may be due to the more realistic assumptions on which this model relies, compared to the FP and RER models. In particular, the SP assumes that prices are sticky in the short run, and hence, that the PPP holds in the long run only.

8.2.3 Comparing the conventional-based volatility models using quarterly data and the standard deviation

Now we turn to the investigation of the most appropriate model for quarterly data using the standard deviation as a volatility measure. The findings of the tests are displayed in Table 8-3. For Algeria the coefficient of inflation difference volatility is insignificant in the SP, suggesting that the FP is better. However, the SP does contain three significant variables compared to only two in the FP equation; thus, the SP is chosen over the FP. Moreover, the use of $\overline{R}^2$ and AIC also suggest choosing the SP over the FP. The j-test rejects all models, and, according to the other criteria, the RER is better than the SP. However, the SP contains three significant variables compared to two in the RER. Therefore, either can be selected as the most appropriate model.

Table 8-3: Results of the preferred models by model selection criteria using quarterly data and the SD measure

Country	Chosen from nested models	j-test	$\overline{R}^2$ and AIC	Chosen model(s)
Algeria	SP	Reject SP, reject RER	RER	SP/RER
Canada	FP	Reject FP, reject RER	FP	FP
Germany	SP	Reject SP, reject RER	SP	SP
Japan	SP	Reject SP, reject RER	SP	SP
Kuwait	FP	Reject FP, reject RER	FP	FP
Libya*	SP	Reject SP, reject RER	SP	SP
UK	FP/SP	Reject SP, reject RER	SP	FP/SP
Venezuela*	SP	Accept SP, reject RER	_	SP

Notes: * indicates caution should be taken when interpreting results of j-test for Libya and Venezuela as the sample size is quite small; i.e. 50 and 40 observations included for these countries respectively.

The SP has a significant parameter of the inflation differential variability in the UK case; however, the coefficient of differenced money supply volatility is significant with a value of 1.87. On the other hand, this variable is statistically significant with a magnitude of one in the FP which meets theoretical expectations. Thus, there is no clear answer to the question of which model is better for this country.

The SP was selected to be the best hypothesis when quarterly data was used alongside the standard deviation as a proxy for volatility in Algeria, Germany, Japan, Libya, the UK and Venezuela. For Canada and Kuwait, the FP was first chosen over the SP and then also over the RER using the $\overline{R}^2$ and AIC, since the j-test rejected all models. Therefore, the SP seems to be the most appropriate model for all countries except for Canada and Kuwait. Again these results reinforce the previous findings that the SP is a short term model and thus works better for more frequent data for most of our sample countries.

8.2.4 Comparing the conventional-based volatility models using monthly data and an ARCH-based measure

The findings of the j-test and the other model selection criteria for the monthly data using ARCH-based volatility measure can be seen in Table 8-4. For the Kuwaiti case the comparison is between the SP and RER as there is no FP model using ARCH method. Here the j-test rejects both models and the use of the other criteria results in the choice of the SP as the best model for this country. However, the RER model has a significant consumption differential volatility compared to none in the SP, and hence, the RER model is in our view preferred to the SP for this country. With respect to the UK, j-test accepted the SP and rejected RER. Thus, the SP is considered to be more suitable for UK when using monthly data and an ARCH-based proxy of variability.

Table 8-4: Results of the preferred models by model selection criteria using monthly data and the ARCH measure

Country	Chosen from nested models	j-test	$\bar{R}^2$ and AIC	Chosen model
Kuwait	-	Reject SP, reject RER	SP	RER
UK	SP	Accept SP, reject RER	-	SP

Looking at the overall results of the model selection so far, one can say that the j-test was not able to give a clear-cut answer regarding the appropriateness of the different models in the majority of cases, so that other criteria had to be used. We should remember that such model selection criteria are indicative rather than precise, and thus, one should be careful when deciding the most appropriate model.

It will be useful to summarise the results reached so far regarding the most appropriate model using different variability proxies and frequencies. Table 8-5 summarises these findings. Generally speaking, the RER was found to be the most suitable model for Algeria regardless of the volatility measure and data frequency used.

The FP can be considered as the most appropriate model for Canada, because it was chosen as the preferred model either individually or jointly with the SP when using the simple measure for quarterly observations.

Table 8-5: Summary of the most preferred models for each individual country using different data frequencies and volatility measures

Country	Monthly & SM	Quarterly & SM	Quarterly & SD	Monthly & ARCH	Most frequent model
Algeria	RER	RER	SP-RER	-	RER
Canada	FP	FP-SP	FP	-	FP
Germany	FP-SP-RER	SP	SP	-	SP
Japan	FP	SP-RER	SP	-	SP
Kuwait	-	-	FP	RER	-
Libya	FP	SP	SP	-	SP
UK	FP-SP	SP	FP-SP	SP	SP
Venezuela	SP	SP	SP	-	SP

With regard to Germany, one can conclude that the SP generally works better than the other models, although it gives a level of performance similar to those of the other models when using monthly data. As regards Japan, we can state that the FP is the best model when using monthly data and the SP is preferred when using quarterly observations irrespective of the volatility proxy used, although the RER was selected alongside the SP when using the SM for quarterly data.

The story is more complicated regarding Kuwait, since there is no model gives any significant coefficient using the simple measure for monthly and quarterly frequencies. On the other hand, when using the standard deviation volatility measure for quarterly data, the FP was found to be the preferred model. However, when using the ARCH-based measure for monthly data the RER was found to give the best performance, so it in turn can be considered as preferred over the others. Thus, apart from these two cases, it is very difficult to assign a specific preferred model to the Kuwaiti data.

Considering the Libyan case it was found that the FP works better for monthly data and the SP is preferred using quarterly data regardless of the volatility proxy used. These findings resemble those from the Japanese case.

As for the UK, both the FP and SP can be considered most appropriate models when using monthly data and the SM and when using quarterly data with the SD measure. However, when using quarterly data with the SM measure, the SP alone was proven to be most appropriate. Moreover, when using the ARCH model for monthly data, the SP was also chosen as the best model.

One can clearly say that the SP is the most suitable model for the Venezuelan data, because it was chosen over other models using different variability measures and different data frequencies.

If comparing the most appropriate models for monthly and quarterly data, it seems that the FP model is preferred when using monthly data, being chosen in 5 out of 7 cases. Conversely the SP is preferred when using quarterly data, as it was chosen in 6 out of 7 cases using the simple measure, after excluding Kuwait, and in 6 out of 8 cases using the standard deviation.

If we are to group countries on this basis, we can report that the monetary exchange rate-based volatility models, namely the FP and SP, work better for developed economies using both monthly and quarterly observations; with priority given to the SP using quarterly data. On the other hand the SP can be considered as the most suitable model for both Libya and Venezuela, whereas the RER is preferred for

Algeria. The good overall performance of the SP models can be interpreted by arguing that prices in the short run are indeed sticky in both groups of countries. Therefore, we may conclude that the SP model is generally more appropriate than the FP and RER models for industrial and less-developed economies in explaining the behaviour of exchange rates. Moreover, one could say that the relatively worse performance of the monetary models in the developing economies could be due to the fact that their exchange rates are not entirely flexible, as is the case for developed countries. More precisely, exchange rates of less-developed countries are not entirely determined by market forces. As discussed in chapter four, the Algerian dinar was kept linked to a composite of currencies till mid 1990s. The Kuwaiti dinar was also pegged to a composite of currencies during our sample period. The Libyan dinar has been fixed to the SDR basket since 1986.

8.3 Comparison of the results using different volatility measures

The next a comparison to make is between the results of the bounds approach using different measures of volatility. This comparison can be conducted with the traditional-based volatility models for quarterly data using the simple measure and the standard deviation to approximate volatility. When using an ARCH-based volatility proxy for monthly data few cases were found in which estimated long run parameters were obtained. Therefore, it is concluded that a comparison involving these proxies will be less fruitful in drawing significant conclusions. Moreover, for the comparison to be consistent, the inclusion of the ARCH-based models requires making adjustments to the variables included and to the sample size, because the regressors included have to exhibit an ARCH effect which is decided using tests detailed in chapter six. Therefore, the results using ARCH models are excluded from our comparisons of the results of the different variability proxies.

The focus in what follows falls on the regressors suggested by the traditional-based variability models. Thus, the intercept will be excluded from the comparison. To make the comparison easier to understand, two different colours are used to highlight the similarities and differences of the estimated coefficients using different measures of volatility. Coefficients coloured green denote significant positive coefficients and those coloured yellow indicate significant negative coefficients. By significant parameter we mean any coefficient that is statistically significantly different from zero at the 10% level or less.

Table 8-6 shows the long run parameters obtained for the SP model using different measures of volatility; that is SM and SD. A comparison of the results in Table 8-6 between those using the standard deviation with those using the simple measure for quarterly data reveals many differences between coefficients in terms of sign and level of significance. For instance, the simple measure and the standard deviation gave similar coefficients in terms of significance and sign in only five cases. An example of this similarity is the coefficient of differenced money supply variability in the UK, where the variable is significantly different from zero and has a direct relationship with the exchange rate volatility using both the SM and the SD. Moreover, from the table one can note that the SD has generally produced more significant parameters than that obtained using the SM. For example, the variability of money supply difference was found to be statistically significant in four countries using the SD, compared to in only three countries when using the SM. Furthermore, the volatility of income differential was found to be significant in all seven cases using the SD compared to in only two cases using the SM. The SM generally produces results that comply with the theoretical expectations in terms of sign more frequently than those of the SD.

Table 8-6: A comparison between the estimated long run coefficients of the SP model using the SM vs SD;

$$vs_t = \beta_0 + \beta_1 vm_t + \beta_2 vy_t + \beta_3 vr_t + \beta_4 v\pi_t + \varepsilon_{0t}$$

Country	β_1		β_2		β_3		β_4	
	SM	SD	SM	SD	SM	SD	SM	SD
Canada	0.04 (1.21)	-0.17** (-2.16)	0.09 (0.76)	0.33** (2.24)	0.003** (2.00)	0.008 (1.59)	2.11** (2.14)	-0.32 (-0.86)
Germany	5.43** (5.71)	-0.19 (-1.16)	0.96 (1.52)	0.97** (3.76)	0.004* (1.69)	-0.01** (-2.11)	-11.15** (-2.20)	2.66** (4.42)
Japan	-0.38 (-0.50)	-0.06 (-0.49)	-0.14 (-0.15)	0.73* (1.73)	0.006** (2.13)	0.09** (2.39)	-9.66* (-1.74)	-2.12** (-2.18)
Kuwait	0.0009 (0.15)	-0.31** (-2.08)	0.001 (1.55)	-0.03** (-1.98)	-0.0003 (-0.39)	0.03 (0.96)	-0.56 (-0.94)	-0.41 (-0.89)
Libya	0.11** (2.02)	-0.10 (-1.65)	0.07** (1.99)	0.23** (3.24)	0.005 (1.19)	0.04 (0.72)	-0.008* (-1.83)	3.48 (1.11)
UK	0.12** (4.02)	1.87** (2.41)	1.35* (1.96)	1.55** (3.05)	-0.0006 (-0.42)	0.003 (0.16)	-0.25 (-0.40)	-6.02** (-3.22)
Venezuela#	0.12 (0.45)	-0.76* (-1.99)	-0.61 (-0.82)	1.38** (2.21)	-0.002 (-0.62)	0.25 (1.29)	4.34** (3.39)	1.67 (0.56)

Note: t-values are in parentheses under each coefficient. **, * indicate significance at 5% and 10% levels respectively. # indicates that the bounds F-tests have found more than one cointegrating relationship in Venezuela.

In addition, for the coefficients which were found significant using both proxies of volatility (7 cases), there are two cases in which the sign of these coefficients changes depending on the volatility measure used. These two coefficients are those of the volatility of differenced interest rate and inflation rate in Germany. Such changes in the sign of coefficient may be due to the way by which the volatility proxy was calculated, that is the SD measures shorter term volatility compared to the SM which approximates longer term volatility. Thus, the variability of interest differential in Germany may have a negative relationship with exchange rate variability in the short term but a positive relationship in a longer term. On the other hand, the opposite might be true for the volatility of inflation differential. This difference between the SM and SD proxies may also explain why some variables are significant using one measure and insignificant using the other. This means that some variables may affect exchange rate variability in terms of its short term volatility but may not affect it in terms of long term volatility. In contrast, some variables may not have an impact on exchange rate variability in the short term volatility and may have an impact in the longer term. Therefore, one can state that the results are in fact sensitive to the volatility measure used.

In addition, the SM yielded only 3 negative parameters out of total 13 significant parameters, particularly for the volatility of inflation differentials. Meanwhile the SD gave 7 negative coefficients out of total 16 significant coefficients. Negative signs for the volatility of differenced inflation using the SM measure were found in Germany, Japan and Libya. One possible explanation to this phenomenon is that in the 1980s and 1990s some countries conducted a policy of inflation targeting by controlling short-term interest rates in an environment of floating exchange rates. Alexandre et al. (2002), for example, found that the introduction of inflation target led to a rise in exchange rate volatility. The authors stated that a more stable inflation was achieved at the expense of greater variability of output, interest rate and exchange rate (see also Gali and Monacelli, 2005). This situation is not true for Venezuela, which may be explained by that this country did not follow the policy of inflation targeting. We may attribute the non-negative sign of the differenced inflation volatility term for Canada to the fact that this country has adopted the inflation targeting policy only over a short period. Canada followed this policy in 1991, and since our sample period starts 1973 through 1998, the effect of inflation targeting policy on our estimated coefficient in Canada may be minimal. The other possible explanation of such negative signs can be due to the phenomenon of overshooting. More precisely, within a relatively short term prices are less flexible (that shows low volatility in the inflation differential), and therefore, the exchange rate may overshoot its long run value (that is, showing more exchange rate volatility), hence we have a negative relationship. One may also say that if prices responded to a shock (thus showing higher inflation differential volatility), then there is no need for exchange rate to respond (that is showing less exchange rate volatility) which implies a negative relationship.

The next comparison is between the estimated long run coefficients obtained from the bounds method when the SM and SD are used to approximate variability for the FP model. The results of such coefficients are included in Table 8-7. By comparing the results one can see that the use of the SM gave 7 significant coefficients, all of which are positive, and there were 13 significant coefficients using the SD, 5 of which are negative. Moreover, as with the SP model, the volatility of differenced income was found to be significant in all countries except for Libya using the SD, whereas it is statistically insignificant in all countries using the SM. Again this can be attributed to the idea of the short run volatility measured by the SD and the long run volatility measured by the SM. Thus, the SD is more likely to pick up small fluctuations in the income differential compared to the SM, since it is believed that the income differential is relatively less volatile compared with other fundamentals. In addition, for the parameters which were found to be significant using both measures of volatility in 4 cases, only in one case the coefficient's sign did change with the volatility measure used. This was the variability of money supply difference in Canada. This can also be explained in terms of short and long run variability measures.

Table 8-7: A comparison between the estimated long run coefficients of the FP model using the SM vs SD;

$$vs_t = \alpha_0 + \alpha_1 vm_t + \alpha_2 vy_t + \alpha_3 vr_t + \varepsilon_{1t}$$

Country	α_1		α_2		α_3	
	SM	SD	SM	SD	SM	SD
Canada	0.07** (2.16)	-0.18** (-2.15)	0.06 (0.56)	0.34** (2.07)	0.003** (2.36)	0.007 (1.25)
Germany	4.84** (5.09)	-0.18 (-0.90)	0.96 (1.19)	0.59** (2.06)	0.002 (1.07)	-0.02** (-2.45)
Japan	0.07 (0.10)	-0.02 (-0.19)	-0.20 (-0.21)	0.93** (2.50)	0.006** (1.98)	0.09** (2.59)
Kuwait	-0.0007 (-0.11)	-0.28** (-2.03)	0.001 (1.47)	-0.02** (-2.05)	-0.0002 (-0.28)	0.02 (0.93)
Libya	0.12** (2.11)	-0.03 (-0.55)	0.03 (1.25)	0.12 (1.15)	0.002 (0.41)	0.02 (0.57)
UK	0.08** (4.14)	1.03** (2.38)	0.92 (1.41)	1.28** (3.06)	-0.0005 (-0.36)	-0.04 (-1.55)
Venezuela#	-0.10 (-0.42)	-0.56** (-2.33)	-0.72 (-0.91)	0.48** (2.83)	0.04** (2.45)	0.13** (3.90)

Note: *t*-values are in parentheses under each coefficient. **, * indicate significance at 5% and 10% level respectively. # indicates that the bounds *F*-tests found more than one cointegrating relationship in Venezuela.

A comparison regarding the long run effects of the RER model using the SM and SD can be found in Table 8-8. From the table one can notice that when using the SM 6 significant coefficients were found. This is the same number found using the SD, but they are not the same parameters. All significant parameters with the SM are positive, whereas 4 of those with the SD are negative. As with the FP, the volatility of money stock differential in Canada changes from having a positive sign using the SM to a negative sign using the SD. The variability of consumption difference is significant only in Algeria and Libya with a positive sign.

Table 8-8: A comparison between the estimated long run coefficients of the RER model using the SM vs SD;

$$vs_t = \gamma_0 + \gamma_1 vm_t + \gamma_2 vc_t + \varepsilon_{2t}$$

Country	γ_1		γ_2	
	SM	SD	SM	SD
Algeria	-0.56 (-1.23)	-0.42** (-3.60)	0.75** (4.13)	0.12** (4.59)
Canada	0.09** (2.08)	-0.11* (-1.88)	0.17 (0.51)	0.0008 (0.06)
Germany	5.69** (4.41)	0.99 (1.22)	-0.33 (-1.02)	-0.14 (-1.24)
Japan	1.05* (1.95)	-0.07 (-0.66)	2.10 (1.41)	0.02 (0.27)
Kuwait	-0.002 (-0.32)	-0.18* (-1.95)	0.0001 (0.03)	-0.02 (-1.62)
Libya	0.13** (2.28)	-0.04 (-0.87)	0.002 (0.36)	0.06** (2.69)
UK	0.08** (3.87)	1.08 (1.63)	-0.35 (-0.56)	-0.24 (-1.30)
Venezuela	0.09 (0.25)	-0.89** (-2.43)	0.13 (1.12)	-0.13 (-1.52)

Note: *t*-values are in parentheses under each coefficient. **, * indicate significance at 5% and 10% level respectively.

Overall, from the results discussed so far for the traditional models in terms of the volatility measures, the following conclusions can be drawn. 26 significant coefficients were found in the three conventional models using the SM, whereas 35 significant parameters were found using the SD. Obtaining different results regarding the significance of the regressors using different volatility measures can be attributed to the underlying process of calculating the variability in these measures. As it was explained in chapter three, the SD was computed for a quarterly frequency using monthly observations. The variability of each quarter is computed by the standard deviation of the data series for the three months which combine in that quarter. Therefore this proxy measures the deviation of observations from the mean of every three months; hence, it measures the variation from a short run average. Thus, one can say that the SD proxy used in this study approximates short run volatility rather than a long run one. Consequently, it picks up volatilities more frequently than other proxies would do.

On the other hand, the SM proxy, as shown in chapter three captures deviations from the mean of the whole period; i.e. the average of about 100 quarters. This is effectively a long run volatility measure which picks up only large changes around the long run mean.

Against this background, it is not surprising that the SD gives more significant variables than the SM proxy, since the former captures minor and major variations whereas the SM picks up only major changes in a series. This may also explain why the estimated coefficient of the volatility of income differential is not significant in all cases for the FP, and to some extent, for the SP using the SM measure. This is the case because the income differential is believed to be less volatile compared with the other regressors. Therefore, the SD will pick up both small and large changes in this variable, whereas the SM will pick up only large changes, which are fewer in this variable. Moreover, this may also be true for the coefficient of consumption differential variability in the RER model; because we believe that the real economic fundamentals such as income and consumption are not as volatile as financial time series, such as the interest and inflation rates (see, for example, MacDonald, 1988).

Furthermore, the SD proxy has yielded almost all of the negative parameters (16) compared to the SM (only 3). Thus, one can say that the SD tends to produce inverse relationships between the regressors and the regressand, whereas the SM tends to give direct relationships. Since our theoretical expectations are that all regressors should have direct relationships with the exchange rate volatility, we could argue that the SM is preferred to the SD using quarterly observations as it produces mostly positive parameters, despite the fact that the SD gives significant variables more frequently than the SM. However, giving negative signs does not necessarily rule out the usefulness of using the SD as a measure of variability. Since the SD measures short term variability compared to the SM which measures longer term variability, there might be a negative relationship between a regressor and exchange rate variability in the short run and a positive or no relationship in a longer run. This can be a possible source for the different signs derived from different volatility proxies.

8.4 Comparison of the results using different frequencies of data sets

Regarding the findings obtained from using different frequency of data, this section compares monthly data using the SM and quarterly data using the same measure for the three traditional-based variability models.

The results using monthly (M) and quarterly (Q) data for the estimates of long run parameters using the simple measure for the SP model are shown in Table 8-9.

In comparing the results of monthly and quarterly observations using the SP in Table 8-9, one can notice that when using monthly data 11 significant coefficients were found, one of which is negative; whereas using quarterly data 13 significant coefficients were found, three of which are negative. These three negative parameters are the volatility of inflation rate differentials in Germany, Japan and Libya, and have been attributed to the hypothesis of contradiction between various economic goals as explained in section 8.3. Moreover, while the variability of differenced inflation is negative using quarterly data in Germany, it is positive using monthly data. This may be explained by the phenomenon of overshooting introduced by Dornbusch (1976).

Table 8-9: A comparison between the estimated long run coefficients of the SP model using monthly and quarterly data and the SM proxy;

$$vs_t = \beta_0 + \beta_1 vm_t + \beta_2 vy_t + \beta_3 vr_t + \beta_4 v\pi_t + \varepsilon_{0t}$$

Country	β_1		β_2		β_3		β_4	
	M	Q	M	Q	M	Q	M	Q
Canada	0.02	0.04	0.13**	0.09	0.001**	0.003**	-0.62	2.11**
	(1.11)	(1.21)	(2.08)	(0.76)	(4.82)	(2.00)	(-0.75)	(2.14)
Germany	0.28*	5.43**	-0.04	0.96	-0.00001	0.004*	15.57**	-11.15**
	(1.83)	(5.71)	(-0.37)	(1.52)	(-0.06)	(1.69)	(2.79)	(-2.20)
Japan	-0.12**	-0.38	-0.46	-0.14	0.004*	0.006**	2.86	-9.66*
	(-2.28)	(-0.50)	(-1.54)	(-0.15)	(1.70)	(2.13)	(0.68)	(-1.74)
Kuwait	-0.004	0.0009	-0.00009	0.001	0.0004	-0.0003	-0.06	-0.56
	(-0.62)	(0.15)	(-0.55)	(1.55)	(0.64)	(-0.39)	(-0.39)	(-0.94)
Libya	-0.0006	0.11**	0.06**	0.07**	0.02**	0.005	-0.19	-0.008*
	(-0.03)	(2.02)	(2.56)	(1.99)	(2.31)	(1.19)	(-0.21)	(-1.83)
UK	0.66**	0.12**	-0.18	1.35*	0.002	-0.0006	0.17	-0.25
	(3.39)	(4.02)	(-1.47)	(1.96)	(1.39)	(-0.42)	(0.37)	(-0.40)
Venezuela#	-0.05	0.12	0.003	-0.61	0.02**	-0.002	10.76**	4.34**
	(-1.04)	(0.45)	(0.02)	(-0.82)	(5.92)	(-0.62)	(5.60)	(3.39)

Notes: *t*-values are in parentheses under each coefficient. **, * indicate significance at 5% and 10% level respectively. # indicates that the bounds *F*-tests found more than one cointegrating relationships in Venezuela.

This implies that in the longer term (here quarterly) prices are more flexible compared to over a monthly span, which means that a shock in money supply will be followed by movements in prices, in which case there is no need for changes in interest rates to clear money market, and hence it is less likely that overshooting will occur. Thus, the more flexible the prices are, the higher the volatility in inflation difference will be, which, in turn, leads to less volatility in exchange rate; that is an inverse relationship. This interpretation conforms to the idea of conflicts between various economic goals, as discussed earlier in this chapter. So these may be a contradiction between stable prices and stable exchange rates when quarterly observations are used.

Except with the volatility of money supply differential in Japan using monthly data, and in the other three cases explained above, all parameters have the expected positive signs irrespective of the data frequency used.

As mentioned above, an inverse relationship was noted between exchange rate variability and variability of money supply difference in Japan on a monthly basis. This might be due to the idea that flexible exchange rates are accompanied by more independent monetary policy as opposed to the situation with fixed exchange rates. More precisely, when exchange rates are flexible (which is likely to imply higher volatility), money supply will not be affected by external shocks, and hence there will be fewer changes in the money supply, representing less volatility. However, this is apparently not the case when using quarterly data.

The significant variables found with monthly and quarterly data correspond to each other in only six cases where they gave similar results in terms of significance and sign, although they may differ in terms of magnitude. Thus, about 50% of the total number of significant variables is found in both monthly and quarterly observations. Thus, the frequency of data may have given different results in terms of significance and sometimes sign. In other words, the findings are sensitive to the data frequency used.

Regarding the FP model, the estimates of long run parameters using monthly and quarterly data and the simple measure are shown in Table 8-10.

Clearly, Table 8-10 indicates that only the coefficient of the volatility of the money supply differential in Japan has a negative sign when using monthly data. The other 15 significant parameters have the expected positive sign.

Table 8-10: A comparison between the estimated long run coefficients of the FP model using monthly and quarterly data and the SM proxy;

$$vs_t = \alpha_0 + \alpha_1 vm_t + \alpha_2 vy_t + \alpha_3 vr_t + \varepsilon_{1t}$$

Country	α_1		α_2		α_3	
	M	Q	M	Q	M	Q
Canada	0.01 (0.96)	0.07** (2.16)	0.14** (2.13)	0.06 (0.56)	0.001** (4.87)	0.003** (2.36)
Germany	0.76** (2.20)	4.84** (5.09)	-0.05 (-0.45)	0.96 (1.19)	-0.000005 (-0.02)	0.002 (1.07)
Japan	-0.12** (-2.33)	0.07 (0.10)	-0.45 (-1.56)	-0.20 (-0.21)	0.004* (1.67)	0.006** (1.98)
Kuwait	-0.005 (-0.82)	-0.0007 (-0.11)	-0.00008 (-0.63)	0.001 (1.47)	0.0002 (0.38)	-0.0002 (-0.28)
Libya	-0.0004 (-0.02)	0.12** (2.11)	0.06** (2.57)	0.03 (1.25)	0.03** (2.41)	0.002 (0.41)
UK	1.83** (2.49)	0.08** (4.14)	-0.10 (-0.90)	0.92 (1.41)	0.002 (1.27)	-0.0005 (-0.36)
Venezuela#	-0.04 (-0.76)	-0.10 (-0.42)	0.70 (1.18)	-0.72 (-0.91)	0.05** (4.73)	0.04** (2.45)

Notes: *t*-values are in parentheses under each coefficient. **, * indicate significance at 5% and 10% level respectively. # indicates that the bounds *F*-tests found more than one cointegrating relationships in Venezuela.

In five cases the significant coefficients with monthly data corresponded to those of the parameters with quarterly data. This means there are 6 cases in which the estimated coefficients using monthly data differed from that estimated using quarterly periods (for example the coefficients of money supply differential variability in Canada, Japan and Libya). Again as in the SP model, all the coefficients of the FP in Kuwait are statistically insignificantly different from zero. Moreover, using the quarterly frequency the volatility of income differential is insignificant in all countries for the FP. However, in the SP this variable is positive and significant in Libya and the UK. This may be due to the omission of the variability of the inflation differential from the regression.

The comparison for the results of the RER model is shown in Table 8-11. One noteworthy point is that the volatility of money supply differential in Japan is still negative using the monthly data as it was with the FP and SP models. However, in contrast with the SP and FP equations, this variable is now significant and positive using quarterly data. We may attribute the significance of this parameter in the RER to the exclusion of relevant variables such as the interest differential variability which was included in the SP and FP models.

Table 8-11: A comparison between the estimated long run coefficients of the RER model using monthly and quarterly data and the SM proxy;

$$vs_t = \gamma_0 + \gamma_1 vm_t + \gamma_2 vc_t + \varepsilon_{2t}$$

Country	γ_1		γ_2	
	M	Q	M	Q
Algeria	-0.18 (-1.65)	-0.56 (-1.23)	0.27** (4.25)	0.75** (4.13)
Canada	0.01 (0.71)	0.09** (2.08)	0.007** (2.94)	0.17 (0.51)
Germany	1.07** (2.68)	5.69** (4.41)	0.003 (0.16)	-0.33 (-1.02)
Japan	-0.12** (-2.06)	1.05* (1.95)	-0.02 (-1.60)	2.10 (1.41)
Kuwait	-0.003 (-0.51)	-0.002 (-0.32)	0.0002 (0.26)	0.0001 (0.03)
Libya	0.03 (1.55)	0.13** (2.28)	0.002 (0.41)	0.002 (0.36)
UK	1.66** (2.54)	0.08** (3.87)	0.0008 (0.07)	-0.35 (-0.56)
Venezuela	-1.02 (-1.12)	0.09 (0.25)	0.11 (0.52)	0.13 (1.12)

Note: *t*-values are in parentheses under each coefficient. **, * indicate significance at 5% and 10% level respectively.

The other significant coefficients have positive signs which comply with our theoretical expectations. However, significant coefficients appearing in common with the two frequencies are only three out of seven. This means there are four cases in which there is a difference according to the frequency considered. Thus, one could in general say that the findings for estimated parameters vary according to the data frequency used.

8.5 The estimated coefficients of the conventional-based volatility models and their policy implications

This section comments on and explains the estimated parameters of the preferred models (as summarised in Table 8-5) and their policy implications. The findings are addressed for each individual country in alphabetical order. For convenience the results of the estimated coefficients of the preferred models for each country are shown again.

Algeria

For Algeria the most preferred model is the RER using different frequencies and measures. From Table 8-12 we can note that using the simple volatility measure for both monthly and quarterly data, the RER gives only one significant variable, which is the volatility of consumption differential. This regressor has a positive sign in both data frequencies and its magnitude is 0.27 using monthly data and 0.75 using quarterly data.

Table 8-12: The estimated coefficients of the preferred models in Algeria

Model	Frequency	Proxy	m	c
RER	Monthly	SM	-0.18 (-1.65)	0.27** (4.25)
RER	Quarterly	SM	-0.56 (-1.23)	0.75** (4.13)
RER	Quarterly	SD	-0.42** (-3.60)	0.12** (4.59)

On the other hand, when using the SD for quarterly data it was found that the variability of money supply differential is significant with a negative sign, and there was a positive and significant coefficient for the variability of consumption differential. The magnitude of the parameter of the volatility money supply differential is estimated at -0.42, which means that about 42% of any change in this regressor will be transmitted into exchange rate volatility in the opposite direction. The estimated coefficient of the variability of consumption difference using the SD is 0.12, meaning that a fall in the volatility of consumption differential by 1 in absolute value would be followed by a fall in exchange rate volatility by only 0.12. This estimate differs from the previous ones obtained using the SM proxy, implying that different measures affect the magnitude of the estimated coefficients in addition to their significance and sign. The significance of money supply difference volatility using the SD can be attributed to the fact that the SD picks up volatility more frequently than the SM does, as discussed previously. This variable has a negative sign using all measures, though. Therefore, we can say that the results of the RER model for the Algerian data using different frequencies and proxies are generally consistent. The Algerian authority, therefore, should take into account the fact that exchange rate variability is mainly affected by the volatility of consumption differential. More precisely, if the government wanted to reduce exchange rate variability, it should reduce the variability of consumption differential. The amount of such a reduction depends on the volatility proxy considered and time span used. Finally, since consumption represents income in the Redux model, and since income in oil-exporting countries is largely represented by oil production, we may attribute the significance of consumption in this country to volatility in oil production.

Canada

It has been concluded that the FP is the preferred model for Canada, as shown in Table 8-5. As can be seen from Table 8-13, using monthly data with the SM it was found that the volatility of income and interest differentials are positive and significant with estimated coefficients of 0.14 and 0.001 respectively. Using quarterly observations with the SM the variability of money supply and interest differentials are found to be positive and significant with estimated coefficients of 0.07 and 0.003 respectively. Compared with the results using monthly data changes are observed in the significant variables. The variability of money supply differential was insignificant for monthly data and became significant using quarterly data. It seems that volatility in money supply differential has no impact on exchange rate volatility in a short period, on the basis of monthly data, but it has a significant impact over a longer quarterly period. In contrast, the variability of income differential seems to have significant effect on monthly basis compared to having no effect on a quarterly basis. Moreover changes in significance can also be seen when using quarterly data with the SD measure, where for the volatility of both the money supply and income differentials. However, the sign of the variability of money supply differential is now negative compared to being positive using the SM.

Table 8-13: The estimated coefficients of the preferred models in Canada

Model	Frequency	Proxy	m	y	r
FP	Monthly	SM	0.01 (0.96)	0.14** (2.13)	0.001** (4.87)
FP	Quarterly	SM	0.07** (2.16)	0.06 (0.56)	0.003** (2.36)
FP	Quarterly	SD	-0.18** (-2.15)	0.34** (2.07)	0.007 (1.25)

This difference is discussed in section 8.3. The magnitudes of the estimated coefficients are fairly similar when using the SM but they differ using the SD. Thus, one can conclude that the results of the FP model for the Canadian data using a variety of volatility proxies and data frequencies are not much in consistent, since these different measures and frequencies have affected the results. Therefore, the policy implications of these results depend on the use of each individual measure and frequency. Finally, we may attribute the significance of money supply differential (with a negative sign in the short run and positive sign in the long run) to the monetary aggregates policy targeting conducted until 1991. Canada is the only industrialized country that has a significant variability of income differential with a magnitude of significantly less than one. This may be as a result of this economy is being in a status of less than full employment.

Germany

With respect to the German case, Table 8-5 shows that the SP was chosen as the most appropriate model using different frequencies and proxies. Using monthly data and the SM the volatilities of money supply and inflation differentials were found positive and significantly different from zero as shown in Table 8-14 below. The magnitudes of these variables are 0.28 and 15.57 respectively meaning that a change in these two regressors by 1 leads to changes in the exchange rate volatility by 0.28 and 15.57 respectively in the same direction. Using quarterly observations with the SM positive and significant variability in the money supply differential was found, which resembles the results using monthly data except with a different magnitude coefficient (5.43). This means that a change by 1 in the volatility of money supply differential causes a change by 5.43 in the exchange rate volatility. This is five times the original change in the regressor. Moreover, we found a positive significant parameter in the variability of interest differential, although at 10% this was not significant using monthly data. This variable has a value of 0.004, implying that a large change in the volatility of interest difference would cause only slight change in exchange rate variability. Furthermore, the volatility of inflation differential is again significant but with a negative sign and a magnitude of -11.15. The change in the sign can be explained by the phenomenon of overshooting as discussed in section 8.4. The magnitude, however, is still similar using both frequencies.

Table 8-14: The estimated coefficients of the preferred models in Germany

Model	Frequency	Proxy	m	y	r	π
SP	Monthly	SM	0.28*	-0.04	-0.00001	15.57**
			(1.83)	(-0.37)	(-0.06)	(2.79)
SP	Quarterly	SM	5.43**	0.96	0.004*	-11.15**
			(5.71)	(1.52)	(1.69)	(-2.20)
SP	Quarterly	SD	-0.19	0.97**	-0.01**	2.66**
			(-1.16)	(3.76)	(-2.11)	(4.42)

Using the SD for quarterly observations we obtained an insignificant money supply differential volatility which contrasts with the previous findings. Also, a positive and significant income differential volatility is now found with a magnitude of almost unity. This magnitude may be explained by the phenomenon of full employment in Germany in most of sample period. This result differs from that using monthly data and the SM; although it is not so far from the results using quarterly data and the SM proxy if we consider a 12% level of significance with the same magnitude. In addition, the variability of interest differential is negative and significantly different from zero. Thus, it has an inverse impact on exchange rate volatility, compared to a positive impact when using the SM. This contradiction may be due to the way of calculating volatility measures, or, more precisely, that short run volatility may have an inverse effect whereas long run volatility has a direct effect. The variability of inflation differential is positive and significant, which contrasts with the result using the simple measure for the same frequency. This may also be attributed to the way of calculating variability measures. The value of the estimated coefficient of this regressor is 2.66, which is very different from that using monthly data (15.57). Thus, there are many differences when using the same preferred model for different measures and frequencies. Some of the differences can be explained and some cannot, unless they are attributed to the effect of different measures and frequencies. Therefore, policy implications again vary with each individual proxy and frequency. The significance of inflation differential variability with high magnitude can be explained by the character of inflation aversion among the German public, which has been translated into inflation targeting policy. This characteristic may also explain the over-reaction towards money supply differential volatility compared to other industrialized countries as indicated by the magnitude of this variable when using quarterly data with a long run volatility measure.

Japan

Looking at the Japanese results one can see that the FP model is preferred using monthly data, and the SP is the most appropriate model for quarterly observations as shown in Table 8-5. The FP gives two significant variables which are the volatilities of money supply and interest rate differentials, which can be seen in Table 8-15 below. The first has a negative relationship with the exchange rate variability with a magnitude of -0.12, meaning that an increase in the volatility of money supply differential by 1 leads to a reduction in exchange rate volatility by 0.12. This magnitude is far lower than the theoretical expectations of a one to one relationship. The volatility of interest differential is positive with an estimated value of 0.004, which means that 0.4% of any change in interest rate difference variability will pass through into exchange rate volatility.

Table 8-15: The estimated coefficients of the preferred models in Japan

Model	Frequency	Proxy	m	y	r	π
FP	Monthly	SM	-0.12** (-2.33)	-0.45 (-1.56)	0.004* (1.67)	–
SP	Quarterly	SM	-0.38 (-0.50)	-0.14 (-0.15)	0.006** (2.13)	-9.66* (-1.74)
SP	Quarterly	SD	-0.06 (-0.49)	0.73* (1.73)	0.09** (2.39)	-2.12** (-2.18)

On the other hand the SP, which encompasses the FP, for quarterly data and the simple measure, gave insignificant coefficient for the variability of money stock differential. However, it gave a positive significant parameter for the variability of interest differential, however, with a magnitude of 0.006 which is very similar to the previous result for the FP. Moreover, the variability of inflation differential is statistically significantly different from zero at 10% level. It has a negative sign and an estimated value of -9.66 implying that a change by 1 in this variable would cause a change in opposite direction by about 10 in exchange rate volatility. Thus, we can say that a small change in inflation differential variability leads to a large opposite change in exchange rate variability, which policy makers should take into account. Using quarterly data with the SD three significant variables were found. The first is the volatility of income differential which has a direct relationship with exchange rate volatility with an estimated coefficient of 0.73. This implies that 73% of a change in volatility of income differential will transfer into exchange rate variability in the same direction. Since the SD captures volatility more frequently than the SM, we attribute the significance of income differential volatility to this reason as discussed in section 8.3. The second significant variable is the volatility of interest differential with a positive sign which resembles previous results for both the FP and SP using quarterly data with the SM. The value of the coefficient of 0.09 is similar to previous estimates. Thus, one can say that all of the models discussed give similar results for this variable;

hence, policy makers should know that an increase in the variability of interest differential by 1 would cause a rise in exchange rate volatility by less than 0.1. The volatility of inflation differential is negative and significant as is the case for quarterly data using the SM, but with a different magnitude. Using the SD the coefficient of inflation differential volatility is -2.12, which are considerably far less than that using the SM. Generally, there are some differences and some similarities in the results of the preferred models for Japan. However, the similarities can be considered to be greater than the differences. Therefore, one can conclude that there is consistency in the results for this country. The results clearly show the significance of interest rate differential volatility. This may be explained by the importance of using interest rate as a monetary instrument in keeping the exchange rate stable. Using a short run volatility proxy we can see that the coefficient of income differential volatility is not significantly different from one. This may reflect the full employment status of the Japanese economy at that time.

Kuwait

With respect to the Kuwaiti case, none of the models appear to be able to produce any significant variable using the simple measure for both monthly and quarterly observations. Despite the lack of significant variables in the regression, the *F* and *t*-tests indicate the existence of a long run relationship. This possibly results from the presence of extreme multicollinearity problem as suggested by Pesaran[18]. However, using the SD measure two regressors seem to be significant for the FP model. One possible explanation of this is that the Kuwaiti data has rather low levels of volatility which cannot be picked up by the SM. However, the SD captures small as well as

[18] Correspondence with Hashem Pesaran was through the following email: mhp1@cam.ac.uk.

large volatilities, and thus appears to be able to pick up such low levels of volatility, hence giving significance to some variables.

Table 8-16: The estimated coefficients of the preferred models in Kuwait

Model	Frequency	Proxy	*m*	*y*	*r*	*c*
FP	Quarterly	SD	-0.28** (-2.03)	-0.02** (-2.05)	0.02 (0.93)	_
RER	Monthly	ARCH	-0.005 (-0.68)	_	_	-0.004* (-1.81)

In Table 8-16 shows that there are two significant variables using the SD, which are the volatility of money supply and income differentials in both the FP and SP. The FP was preferred to the other models, and hence, we would comment on its findings. The variability of money supply differential has an inverse impact on exchange rate volatility with a magnitude of 0.28 meaning that 28% of a change in volatility of money stock differential will pass through to exchange rate volatility in the opposite direction. Furthermore, the volatility of income differential is also significant with a negative sign, but its coefficient is only -0.02. This implies that only 2 per cent of a change in this regressor would be transmitted into exchange rate volatility in the opposite direction. Thus, Kuwaiti authorities should take into account that lower exchange rate volatility can be achieved only at the expense of higher money supply differential volatility or higher variability of income differential. Moreover, when using a volatility measure estimated from an ARCH model for monthly data, the RER gives significant negative consumption differential variability with a magnitude of -0.004. This is actually consistent with the results of the FP discussed above, since income and consumption can be considered as alternatives. The negative sign of the coefficients of money supply, income and consumption differentials variability may be due to the fact that the Kuwaiti authorities try to keep the exchange rate against the

US dollar as stable as possible. Thus, even with higher volatility in these fundamentals, the exchange rate volatility would be low.

Libya

For Libya the preferred model for monthly data is the FP and for quarterly data the SP as shown in Table 8-5. Thus, we will comment on the results of the FP for monthly data and the results of the SP for quarterly data. From Table 8-17 we can note that, using monthly data, both the volatilities of income and interest differentials were found to be significant with the expected positive sign, which means that an increase in either of these volatilities will cause an increase in exchange rate volatility. The estimated coefficients of these regressors are 0.06 and 0.03 respectively, implying that only 6% and 3% of a change in the variables would be transmitted to exchange rate volatility in the same direction. For the results from quarterly data, the SP has given three significant variables using the SM and only one significant variable using the SD. For the SM, the volatilities of money stock and income differentials are positive and significant at 5% level implying that they have a direct effect on exchange rate variability. They have coefficient of magnitude of about 0.1, meaning that only 10% of a change in these variables will be transferred into exchange rate volatility in the same direction. The volatility of inflation differential is significant at 10% level with negative sign, meaning that it has an inverse impact on exchange rate volatility. The value of the estimated parameter is -0.008 which implies that an increase in the volatility of inflation differential by 1 would reduce exchange rate variability by only 0.008. Thus, the Libyan government should know that to keep exchange rate volatility low, it must allow high inflation differential variability.

Table 8-17: The estimated coefficients of the preferred models in Libya

Model	Frequency	Proxy	m	y	r	π
FP	Monthly	SM	-0.0004 (-0.02)	0.06** (2.57)	0.03** (2.41)	–
SP	Quarterly	SM	0.11** (2.02)	0.07** (1.99)	0.005 (1.19)	-0.008* (-1.83)
SP	Quarterly	SD	-0.10 (-1.65)	0.23** (3.24)	0.04 (0.72)	3.48 (1.11)

Using quarterly data and the SD, however, we found that only the volatility of income differential was significant at the 5% level with a positive sign. The magnitude of the coefficient indicates that a rise in variability of income differential by 1 would cause a rise in exchange rate volatility by about 0.23, meaning that Libyan authorities would have to lower income differential variability to reduce exchange rate volatility. All three of the models considered find the variability of income differential significant with a positive impact on exchange rate variability. Although there are some differences among the results using these models, they are generally consistent in that there are no contradictory signs. We can see that the volatility of income (proxied by oil production) is significant in all models in this oil-exporting country. This is similar to other oil-exporting developing countries.

United Kingdom

As far as the UK is concerned, both the FP and SP can be considered as preferred models, as can be seen in Table 8-5. We will comment on the results of SP, however, because it contains an extra significant variable and in addition to estimated coefficients found in the FP. Starting with monthly data and the SM, there is only one significant variable, which is the volatility of money supply differential with a positive sign as can be seen from Table 8-18 below. This implies that an increase in this variable will cause a rise in exchange rate volatility. The estimated magnitude of the coefficient is 0.66 which is close to the theoretical expectations, meaning that two-thirds of a change in the volatility of money supply differential would be transferred into exchange rate volatility in the same direction. Considering quarterly data using the SM, the findings for the SP show that there are two significant variables, namely the volatility of money supply differential (at the 5% level) and the volatility of income differential (at the 10% level). Surprisingly, the estimated coefficient of the variability of money supply differential is now quite low (0.12) compared to its previous counterpart, namely that for monthly data using the SM. This means that only 12% of a change in the variability of money supply differential would pass through to exchange rate volatility. The variability of income differential is now significantly different from zero at 10% with a positive sign and a value of 1.35, meaning that a rise in income differential variability by 1 will cause a rise in exchange rate volatility by 1.35.

Table 8-18: The estimated coefficients of the preferred models in UK

Model	Frequency	Proxy	m	y	r	π
SP	Monthly	SM	0.66** (3.39)	-0.18 (-1.47)	0.002 (1.39)	0.17 (0.37)
SP	Quarterly	SM	0.12** (4.02)	1.35* (1.96)	-0.0006 (-0.42)	-0.25 (-0.40)
SP	Quarterly	SD	1.87** (2.41)	1.55** (3.05)	0.003 (0.16)	-6.02** (-3.22)
SP	Monthly	ARCH	-0.41 (-0.45)	0.99 (1.57)	0.00001 (0.86)	-9.14** (-2.60)

Looking at the results of the SP using quarterly observations and the SD , we find a positive significant parameter for the volatility of money supply differential with a magnitude of 1.87, meaning that a small change in this variable will be followed by a larger change in exchange rate variability. The coefficient of volatility of income differential is very similar to its previous counterpart (i.e. quarterly with the SM) in terms of sign and magnitude. Moreover, the volatility of inflation differential is now negative and significant at 5% level with an estimated value of -6.02, which implies that a rise in this variable by 1 would lead to a reduction in exchange rate volatility by 6. Thus, monetary authority should take into account that to reduce exchange rate variability by 1, for instance, it should allow inflation differential volatility to rise by only 0.17. In addition, using the ARCH-based measure for monthly data, the SP shows significant inflation differential variability. It is negative and has an estimated value of -9.14, which is very similar to previous parameters although the magnitude is rather different. Therefore, the results of the SP using different frequencies and volatility proxies are generally consistent except for some differences in terms of the magnitudes. The coefficient of the volatility of the money supply differential in two out of three significant cases is not significantly different from one, which complies with the expectations. The significance of this variable may be explained by the policy of monetary targeting which was conducted till 1992. As in the German and Japanese cases, the coefficient of the variability of the income differential did not significantly differ from one. Again the coefficient's magnitude of the inflation differential variability is large as it was for Germany and Japan. The UK authorities have used interest rates as a monetary policy only recently, this may explain that its coefficient is not significant in all models.

Venezuela

Finally, for Venezuela, Table 8-5 indicates that the most preferred model is the SP for all frequencies and measures considered. Table 8-19 shows that when using monthly data with the SM, the volatilities of interest and inflation rates differentials were found to be significant at 5% level with positive signs, implying that there are direct relationships between exchange rate volatility and these two variables. The estimated magnitude for the first variable is 0.02, meaning that an increase by 1 in this regressor will be followed by increment in exchange rate variability of only 0.02. The magnitude of the second regressor is 10.76 which mean that a rise in the variability of inflation differential by 1 would be followed by as high increase as 11 in exchange rate variability. Therefore, the Venezuelan government should take into account the fact that a large change in exchange rate volatility can be achieved by a small change in the variability of inflation differential. However, using the SM for quarterly data the volatility of inflation differential was found to be the only significant variable, with a positive sign and a magnitude of 4.34. Although this coefficient is now lower than its previous counterpart, it still gives a higher than one to one relationship.

Table 8-19: The estimated coefficients of the preferred models in Venezuela

Model	Frequency	Proxy	m	y	r	π
SP	Monthly	SM	-0.05 (-1.04)	0.003 (0.02)	0.02** (5.92)	10.76** (5.60)
SP	Quarterly	SM	0.12 (0.45)	-0.61 (-0.82)	-0.002 (-0.62)	4.34** (3.39)
SP	Quarterly	SD	-0.76* (-1.99)	1.38** (2.21)	0.25 (1.29)	1.67 (0.56)

Using the SD for quarterly observations quite different variables were found to be significant; i.e. the volatilities of money supply and income differentials. The volatility of money supply differential has a negative sign and significant at 10% level. The magnitude of the coefficient of this regressor is -0.76, implying that a rise in this variable by 1 would cause almost the same amount of a change in the volatility of exchange rate, but in the opposite direction. The variability of income differential is positive and significant at 5% level with an estimated value of 1.38. This implies that an increase in the volatility of income differential by1 will be followed by an increase of more than 1 in exchange rate volatility. Generally speaking, the first two models gave similar results; however, the third gave quite different findings in terms of the variables found significant. There is no contradiction among the results for these models, though. Thus, one can state that the preferred model in Venezuela in general gives consistent results using different measures and frequencies. The high magnitude of inflation differential variability may reflect the sensitivity to inflation which has been increasing steadily since the mid 1980s. As in the rest of developing country sample, the volatility of income differential is significant using the SD proxy but with a relatively larger value.

Overall, one can state that different volatility proxies, frequencies and models have generally affected the significance, sign and magnitude of the estimated coefficients. Thus, policy implications would depend on each of these individual factors. It should also be noted that the volatility of inflation differential has a relatively large impact on the variability of exchange rates in most cases for which it was found significant.

Given that we are talking about the volatility of differential fundamentals, it is not straightforward for a government to be able to control exchange rate volatility by controlling its own fundamentals. This is because exchange rate variability depends also upon the fundamentals of the base country, which is the US in our case. Therefore, reducing exchange rate instability is not an easy task for a country since it has no power over the variables of its partners. As a consequence, the policy implications described here are merely indicative of possible ways of controlling exchange rate variability.

8.6 The estimated coefficients of Devereux-Lane model and their policy implications

In contrast to the results of previous studies which have found a significant impact of the variables of the OCA theory on the volatility of exchange rates in developed and developing countries, such a relationship has been found only in two developing countries in this study. Using the hybrid model no improvements were gained on the findings obtained from the original Devereux-Lane model. Using Devereux-Lane model long run relationships were found only in the cases of Algeria and Venezuela. In the Algerian case, there are only two significant variables, namely trade and finance as shown in Table 8-20. This means that cycle, size and external finance have no long run effect on exchange rate volatility. For the significant variables, trade has a positive sign implying that an increase in trade with the US leads to an increase in exchange rate volatility. This conflicts with theoretical expectations. This direct relationship can be explained by that the US is not a major trading partner of Algeria. The average total trade between Algeria and US as a percentage to the Algerian GDP for the period 1973-1998 is 8%. Thus, the Algerian authority might not care much about their exchange rate against the US dollar; and hence exchange rate volatility may rise as the trade rises. The value of the estimated coefficient is 0.01 which means that an increase by 1 in absolute value in trade would cause an increase in exchange rate variability by only 0.01. On the other hand, the finance variable has the correct expected negative sign, implying that a rise in internal finance leads into a decrease in exchange rate volatility. This conforms to the argument that when domestic financial sector is more developed, the risk premium is likely to be less important. The magnitude of the coefficient is -0.01, which means that a rise in finance by 1 would cause a reduction in exchange rate variability by only 0.01. The monetary authority, thus, should know that developing the internal financial sector lowers exchange rate variability, although this influence is small. The non-significance of external debt in Algeria may be explained by the fact that Algeria has no major debt denominated in US dollars.

Table 8-20: The estimated coefficients of the Devereux-Lane model for Algeria and Venezuela

Country	*cycle*	*trade*	*finance*	*size*	*extfin*
Algeria	0.11	0.01*	-0.01**	0.003	-0.0003
	(1.28)	(2.05)	(-2.41)	(0.40)	(-0.80)
Venezuela	3.19	-0.22*	-0.31**	0.04	-9.5×10^{12}*
	(1.27)	(-1.98)	(-3.39)	(0.43)	(-2.36)

In the Venezuelan case, we found three significant parameters were found using Devereux-Lane model; i.e. trade, finance and external finance. All had the correct expected signs. The US is the main trading partner for Venezuela with an average total trade of 20% of the Venezuelan GDP for the period 1973-1998 is 20%. Thus, an increase in trade with the US means a decrease in exchange rate volatility. This implies that the government should pay an attention to keeping exchange rate volatility as low as possible with its trading partner to protect importers and exporters from the inverse effects of unstable exchange rate. The magnitude of the parameter is estimated at -0.22, which means a rise in trade by 1 cause a decline in exchange rate variability of 0.22 which is quite reasonable. The finance term is significant and negative, meaning that a rise in internal finance leads to a decrease in exchange rate volatility, as theoretically expected. About 31% of an absolute change in finance will be conveyed into exchange rate volatility, which is fairly high compared to that of Algeria. Finally, the external debt seems to have a significant impact on exchange rate volatility in Venezuela, reflecting the sensitivity towards foreign debt. A rise in the external financial dependence would cause a decline in the regressand, implying that countries that cannot issue debt in their own currencies would not prefer an adjustable exchange rate in response to external shocks. This corresponds to the theoretical expectations of Devereux and Lane. Therefore, the Venezuelan government should prefer to keep exchange rate volatility as low as possible as its foreign debt gets larger. The value of the estimated coefficient, however, is very low, which means that any absolute change in external debt will cause a very tiny change in exchange rate variability.

Finally, the results of the augmented Devereux-Lane model, which are shown in Table 8-21, require some comments. Only one result was found using this model for the case of Algeria. This model has given quite interesting results in terms of the number of significant variables. Similar results can be seen for the variables of trade and finance to those discussed above using the original Devereux-Lane model. However, in the augmented model three other significant variables were found. Size of the economy is significant and has the correct positive sign, implying that a direct relationship exists between the size of the economy and exchange rate variability in Algeria. The magnitude of the coefficient means that only 2% of any change in the size of the economy will pass through to exchange rate volatility. The external financial dependence is now significant with the right sign, which indicates an inverse relationship between external debt and exchange rate volatility in Algeria.

Table 8-21: The estimated coefficients of the Devereux-Lane equation augmented with the volatility of the consumption differential for Algeria

Country	*cycle*	*trade*	*finance*	*size*	*extfin*	*vc*
Algeria	0.006	0.02**	-0.02**	0.02**	-0.001**	0.27**
	(0.17)	(5.16)	(-8.34)	(4.53)	(-7.19)	(6.79)

The results show that an increase by 1 in the external debt will cause a reduction in exchange rate variability by 0.001. As with the RER model, the volatility of consumption differential has a positive impact on exchange rate fluctuations. If the variability of consumption differential rises by 1, the volatility of exchange rate would go up by about 0.27. Thus, one can say that exchange rate instability is more responsive to changes in the consumption differential variability compared to the other regressors in the augmented Devereux-Lane model. The difference in the results obtained for Algeria using Devereux-Lane model and the augmented Devereux-Lane model can be attributed to small sample size. Adding another regressor to the equation reduced degrees of freedom, and hence may have affected the results.

Surprisingly, the Devereux-Lane model does not work well for any of the developed countries in the sample and works for only two of the developing country sample: Algeria and Venezuela. This partly matches our expectations that the Devereux-Lane model would perform better for developing countries than for developed ones. Kuwait and Libya have some similar characteristics which may prevent this model from working well for these countries. Firstly, both economies are small in terms of economy size compared to Algeria and Venezuela. Secondly, they do not have significant amounts of external debt as is the case for Algeria and Venezuela. Thirdly, neither Kuwait nor Libya those have a major trading relationships with the US. So these factors may cause the Devereux-Lane model to badly perform for these two countries. Venezuela is an outstanding case as it has major trading links with the US, suffers from heavy foreign debt and is geographically close to the US. Algeria on the other hand suffers from heavy external debt and has relatively quite a large economy. These factors may play a role in leading the Devereux-Lane model to work quite well for such countries.

8.7 Conclusion

A variety of comparisons have been made involving different models, measures and data spans. In section 8.2 it was found that the SP model performs better using quarterly data and that the FP outperforms other models using monthly observations. This is generally more applicable to developed countries, whereas it was difficult to find a preferable model for developing countries. In section 8.3 it was concluded that using the SD as a volatility measure generally gives significant variables more frequently than are obtained when using the SM. Moreover, the SD gives more inverse relationships than are obtained by the SM. The sort of volatility measured by the two proxies could be a possible reason behind such negative coefficients. On the other hand the SM has mostly given positive and significant variables which conform to theoretical expectations, despite giving fewer significant parameters. Section 8.4 involves a comparison between the results using different frequencies of observations. It was found that most estimated coefficients are positive, except for a few cases using monthly or quarterly data; however, about 50% of the significant parameters found were different when using monthly and quarterly data are not the same parameters. All in all, we may say that the results in general are quite sensitive to the models, measures and frequencies used in the regressions. In section 8.5 it was found that the results from the preferred models usually differed from each other, suggesting that their policy implications depend on the model, variability proxy and data frequency used. Moreover, for most cases in which was found significant inflation differential variability, this regressor had a large impact on exchange rate volatility as indicated by its large magnitude. It was found that the Devereux-Lane model works better for developing economies, and particularly Algeria and Venezuela as reported in section 8.6, and does not work for any developed country in our sample. In addition, the augmented Devereux-Lane model does not improve (or may worsen) the results obtained by the original Devereux-Lane model.

Chapter nine

Summary and Conclusions

The main purpose of this thesis is to find out the fundamental causes of exchange rate volatility in a sample of countries representing two different types of economies. The previous literature in the context of international macroeconomic exchange rate theory has concentrated on the underlying factors that determine exchange rate levels. The empirical examination conducted in this literature has been unable to reach conclusive results regarding the elements that drive exchange rate behaviour. Our focus in this thesis, on the other hand, is on the underlying macroeconomic fundamentals that determine exchange rate volatility rather than its level. This is because it is widely believed that exchange rate variability creates uncertainty about future rates which in turn affects decisions made by economic agents, particularly under the assumption of risk aversion. Therefore, we assume that exchange rate volatility is driven by the volatility of certain economic fundamentals. These fundamentals are the same ones introduced by economic theories of exchange rate to be the determinants of exchange rate levels.

Chapter two presented the monetary exchange rate models which are the most popular models of exchange rates used in the literature. The frameworks of two versions of the monetary models of exchange rates are presented, namely the flexible price and sticky price Dornbusch-Frankel exchange rate models. The empirical estimation of these models has generally produced contradictory findings with respect to the underlying determinants of exchange rates. This chapter also presented a new macroeconomics exchange rate model for open economy which is known as "Redux model". Empirical research on this model has been scarce. This model and the previous two monetary models assume the validity of the PPP hypothesis in the long run at least. Thus, we have tested this hypothesis for our sample countries, and the results globally support it. Since it is generally accepted that exchange rate volatility creates uncertainty regarding future costs or proceeds among participants in foreign exchange markets, this issue was given closer attention in chapter three. The importance of addressing exchange rate variability was discussed and the issue of approximating exchange rate volatility was dealt with. In particular, the advantages and disadvantages of most variability proxies used in the literature were discussed, and three measures were chosen to approximate volatility in our study. These proxies were: a simple measure which mainly captures long run volatility; the standard deviation which mainly captures short term volatility; and an ARCH model-based volatility proxy which can capture some important features of financial time series. Different proxies of exchange rate variability can give different results as they vary in the methods by which they are computed. Concerning the impact of exchange rate volatility on international trade, for instance, the various proxies of exchange rate variability can be blamed for ambiguous results (see, for example, Abbott, 1999). Therefore, it is assumed that the variability proxies chosen may produce different empirical results for the estimations.

The empirical models in this study link exchange rate variability to the volatilities of some macroeconomic fundamental differentials, as originally proposed by previous economic literature. These were introduced in chapter four. Our models explicitly address exchange rate volatility, instead of its level by considering the variance of both sides of the conventional the monetary exchange rate models and Redux model. Bayoumi and Eichengreen (1997a;b) have, in fact, introduced another model which also explicitly addresses exchange rate volatility. This model was derived from the theory of optimum currency areas, which provides some factors that can help in explaining the choice of exchange rate arrangements. Devereux and Lane (2003) extended such an OCA theory-based model of exchange rate volatility to include a set of financial variables in addition to those introduced by the OCA theory. The Devereux-Lane model was also introduced in chapter four and was compared with the traditional exchange rates models. Moreover, an ad hoc exchange rate volatility model was developed which augments the Devereux-Lane model with our transformed volatility models.

Previous studies in different fields of economics including that of exchange rates determination, in particular, have since the late 1980s used the residual-based cointegration method of Engle and Granger (1987) and/or the multivariate cointegration method of Johansen (1988). Recent developments in econometric methodology suggested an appropriate way to estimate cointegrating relationships. The bounds testing approach to cointegration proposed by Pesaran et al. (2001) allows the estimation of long run structural relationships encompassing both I(1) and I(0) variables simultaneously. This method, therefore, represents an advance over the two previous cointegration approaches. Furthermore, the bounds method is robust for small samples compared to the former methods. Chapter five outlined both the Johansen and the bounds approaches in addition to the unit root tests which were to be used in the empirical analysis.

In chapter six the results of the unit root tests were provided, and they show that our time series data are either I(0) or I(1), meaning that the bounds testing approach is the only suitable method for cointegration analysis with our data sets. Moreover, we attempted to estimate a volatility proxy using ARCH models for monthly and quarterly data. Apparently as a result of using low frequency data, variability estimates were obtained for only a few cases using monthly data and almost none using quarterly data. This in fact has reduced our ability to investigate the effects of using different volatility measures on the findings from the empirical work.

The empirical analysis presented in chapter seven tested the influence of the volatility of some economic fundamentals on exchange rate volatility for eight countries over the sample period 1973-1998, using monthly and quarterly observations for the traditional exchange rates models-based volatility models with different variability proxies. Annual data were used for the Devereux-Lane model with the standard deviation measure. The main results obtained are as follows: (1) significant long run relationships were found between exchange rate variability and its regressors using different models, frequencies and measures for most cases considered; (2) unique cointegrating relationships were confirmed in most cases in which level relationships were found, implying that the bounds testing method is eligible to be used for estimating the long run parameters; (3) a modest number of significant coefficients was obtained; (4) the monetary approach-based volatility models generally work better for developed countries than for the less-developed economies.

Comparisons between the results obtained in chapter seven using different models, data frequencies and volatility proxies were performed in chapter eight. The major findings from these comparisons can be outlined as follows: (1) the results of long run coefficients were sensitive to the chosen volatility proxy in that different variability measures have generally given different results in terms of the significant coefficients, their magnitudes and signs; (2) the results of long run coefficients were also sensitive to the data frequencies used; (3) some models were found to be more appropriate for the data for each country under some particular conditions; (4) the exchange rate volatility models based on the sticky-price real interest exchange rate monetary model generally, performed better for both industrial and less-developed countries than the other competing models considered in this study. This may be due to the more realistic assumptions made by this model compared to the other models. Moreover, within the SP model the variability of inflation differential was found to have larger effect on the variability of exchange rates in most cases, compared to the other regressors which were found to be significant. Explanations of the estimated long run coefficients of the preferred models were then given alongside their policy implications at the end of chapter 8. The main findings of that section are as follows: (1) the coefficients of income differential volatility were not significantly different from a value of one in the industrialized countries, except for Canada, which may be due to the high level of employment in these economies; (2) unlike the inflation which has large coefficients in most cases, the coefficient of interest rate differential variability is quite low in all significant cases, which imply a low response of exchange rate volatility to this variable; (3) the negative signs, especially for the inflation differential volatility, may be explained by the fact that when a shock occurs and prices respond there is no need for exchange rates to respond or when exchange rates respond there is no need for prices to respond, so that there is an inverse relationship between them.

In the light of the above summary the main contributions of this thesis can be highlighted as follows: (1) it has been established that exchange rate volatility is, indeed, at least partially driven by the volatility of some fundamentals which were built upon fundamental differentials originally proposed by previous exchange rate theories; (2) the results of long run relationships are sensitive to the volatility proxy used and the chosen frequency of data.

Avenues for future research are therefore opened by this work. Investigating the causes of exchange rate variability can be extended in various ways; For example, extending the sample period to incorporate recent years; i.e. after introducing the Euro; expanding the sample of countries to include a larger range of different types of economies; and using more volatility measures could be useful in finding out more about the issue of exchange rate volatility. Moreover, the bounds testing approach assumes the presence of a single level relationship. Pesaran et al. (2001) indicated that future developments with the bounds method could encompass more than one level relationship. If such developments became available, it would be interesting to apply them to cases in which more than one cointegrating equation is found. Furthermore, future research with our models can be expanded to involve other models of exchange rates, such as the portfolio balance model. In addition, one could also proceed to investigate the predictive power of our exchange rate volatility models.

Bibliography

Abbott, A., Darnell, A. C. and Evans, L. (2001) The Influence of Exchange Rate Variability on UK Exports *Applied Economics Letters,* **8,** 1 47-49.

Abbott, A. and De Vita, G. (2002) Testing the Long-Run Structural Validity of the Monetary Exchange Rate Model *Economics Letters,* **75,** 2 157-64.

Abbott, A. J. (1999) *An investigation into the influence of exchange rate variability on UK export volumes and prices.* PhD thesis. Durham University.

Abbott, A. J. and De Vita, G. (2001) A Reassessment of the Long-Run Validity of the Flexible Price Monetary Exchange Rate Model *Economic Issues,* **6,** 1 47-57.

Abbott, A. J. and De Vita, G. (2002) Long-Run Price and Income Elasticities of Demand for Hong Kong Exports: A Structural Cointegrating VAR Approach *Applied Economics,* **34,** 8 1025-32.

Abbott, A. J. and De Vita, G. (2003) Another Piece in the Feldstein-Horioka Puzzle *Scottish Journal of Political Economy,* **50,** 1 69-89.

Abrishami, H. and Mehrara, M. (2002) ARDL Approach to the Demand for Disaggregate Imports: The Case of Iran *Iranian Economic Review,* **7,** 7 87-110.

Aghevli, B. B., Khan, M. S. and Montiel, P. J. 'Exchange Rate Policy in Developing Countries: Some Analytical Issues', in International Monetary Fund, IMF Occasional Papers: 78.

Aguirre, A., Ferreira, A. and Notini, H. (2003) 'The Impact of Exchange Rate Volatility on Brazilian Manufactured Exports', in CENTRE FOR RESEARCH IN INTERNATIONAL ECONOMICS, pp. 1-12.

Ahn, B. C. and Oh, K.-Y. (2001) Is the Monetary Model Useful in Explaining Exchange Rates?--Panel Cointegration Evidence *Seoul Journal of Economics,* **14,** 2 169-82.

Akhtar, M. A. and Hilton, R. S. (1984) Effects of Exchange Rate Uncertainy on German and U.S. Trade *Federal Reserve Bank of New York Quarterly Review,* **9,** 1 7-16.

Akhtar, M. A. and Hilton, R. S. (1984) 'Exchange Rate Uncertainty and International Trade: Some Conceptual Issues and New Estimates for Germany and The United States', in Federal Reserve Bank of New York Research Papers: 8403, pp. 78.

Alam, M. I. and Quazi, R. M. (2003) Determinants of Capital Flight: An Econometric Case Study of Bangladesh *International Review of Applied Economics,* **17,** 1 85-103.

Aleisa, E. A. and Dibooglu, S. (2002) Sources of Real Exchange Rate Movements in Saudi Arabia *Journal of Economics and Finance,* **26,** 1 101-10.

Alexandre, F., Driffill, J. and Spagnolo, F. (2002) Inflation Targeting, Exchange Rate Volatility and International Policy Coordination *Manchester School,* **70,** 4 546-69.

Alexius, A. (2001) Sources of Real Exchange Rate Fluctuations in the Nordic Countries *Scandinavian Journal of Economics,* **103,** 2 317-31.

Al-qudair, K. (2005) The Relationship between Government Expenditure and

Revenues in the Kingdom of Saudi Arabia: Testing for Cointegration and Causality *Economics and Administration,* **19,** 1 31-43.
Altinay, G. and Karagol, E. (2004) Structural Break, Unit Root, and the Causality between Energy Consumption and GDP in Turkey *Energy Economics,* **26,** 6 985-94.
Amuedo-Dorantes, C. and Pozo, S. (2001) Exchange-Rate Uncertainty and Economic Performance *Review of Development Economics,* **5,** 3 363-74.
Anaman, K. A. and Mahmod, T. H. (2003) Determinants of Supply of Non-oil Exports in Brunei Darussalam *ASEAN Economic Bulletin,* **20,** 2 144-57.
Andersen, T. G., Bollerslev, T. and Diebold, F. X. (2002) 'Parametric and Nonparametric Volatility Measurement', in National Bureau of Economic Research, Inc, NBER Technical Working.
Arndt, S. W., Thursby, M. and Willett, T. D. (1985) 'Flexible Exchange Rates and International Trade: An Overview', in Arndt, S. W. e., Sweeney, R. J. e. and Willett, T. D. e.(Eds) *Exchange Rates, Trade, and the U.S. Economy.*
Artis, M. and Lewis, M. (1981) *Monetary Control in the United Kingdom.* Philip Allan: Oxford.
Asafu-Adjaye, J. (1999) 'Exchange rate variability and export growth in Fiji', in Asia Pacific Press, pp. 1-13.
Asseery, A. and Peel, D. A. (1991) The Effects of Exchange Rate Volatility on Exports: Some New Estimates *Economics Letters,* **37,** 2 173-77.
Atkins, F. J. and Coe, P. J. (2002) An ARDL Bounds Test of the Long-Run Fisher Effect in the United States and Canada *Journal of Macroeconomics,* **24,** 2 255-66.
Atkins, F. J. and Serletis, A. (2003) Bounds Tests of the Gibson Paradox and the Fisher Effect: Evidence from Low-Frequency International Data *Manchester School,* **71,** 6 673-79.
Backus, D. (1984) Empirical Models of the Exchange Rate: Separating the Wheat from the Chaff *Canadian Journal of Economics,* **17,** 4 824-46.
Bahmani-Oskooee, M. (1991) Is There a Long-Run Relation between the Trade Balance and the Real Effective Exchange Rate of LDCs? *Economics Letters,* **36,** 4 403-07.
Bahmani-Oskooee, M. (1998) Do Exchange Rates Follow a Random Walk Process in Middle Eastern Countries? *Economics Letters,* **58,** 3 339-44.
Bahmani-Oskooee, M. and Mirzai, A. (2000) Real and Nominal Effective Exchange Rates for Developing Countries: 1973:1-1997:3 *Applied Economics,* **32,** 4 411-28.
Bahmani-Oskooee, M. and Nasir, A. B. M. (2004) ARDL Approach to Test the Productivity Bias Hypothesis *Review of Development Economics,* **8,** 3 483-88.
Bahmani-Oskooee, M. and Ng, R. C. W. (2002) Long-Run Demand for Money in Hong Kong: An Application of the ARDL Model *International Journal of Business and Economics,* **1,** 2 147-55.
Bailey, M. J., Tavlas, G. S. and Ulan, M. (1986) Exchange-Rate Variability and Trade Performance: Evidence for the Big Seven Industrial Countries *Weltwirtschaftliches Archiv,* **122,** 3 466-77.

Bailey, M. J., Tavlas, G. S. and Ulan, M. (1992) 'The Impact of Exchange-Rate Volatility on Export Growth: Some Theoretical Considerations and Empirical Results', in Bailey, M. J.(Ed), *Studies in positive and normative economics.*

Baillie, R. T. and Selover, D. D. (1987) Cointegration and Models of Exchange Rate Determination *International Journal of Forecasting,* **3,** 1 43-51.

Bardsen, G. (1989) Estimation of Long Run Coefficients in Error Correction Models *Oxford Bulletin of Economics and Statistics,* **51,** 3 345-50.

Batchelor, R. A. and Wood, G. E. (1982) *Exchange rate policy.* Macmillan Press: London.

Bayoumi, T. (1994) 'A Formal Model of Optimum Currency Areas', in International Monetary Fund, IMF Working Papers: 94/42.

Bayoumi, T. and Eichengreen, B. (1997a) 'Ever Closer to Heaven? An Optimum Currency Area Index for European Countries', in Eichengreen, B.(Ed), *European monetary unification: Theory, practice, and analysis.*

Bayoumi, T. and Eichengreen, B. (1997b) 'Optimum Currency Areas and Exchange Rate Volatility: Theory and Evidence Compared', in Cohen, B. J. e.(Ed), *International trade and finance: New frontiers for research: Essays in honor of Peter B. Kenen.*

Bayoumi, T. and Eichengreen, B. (1998) Exchange Rate Volatility and Intervention: Implications of the Theory of Optimum Currency Areas *Journal of International Economics,* **45,** 2 191-209.

Beaudry, P. and Devereux, M. B. (1995) Money and the Real Exchange Rate with Sticky Prices and Increasing Returns *Carnegie-Rochester Conference Series on Public Policy,* **43,** 55-101.

Belke, A. (2001) Exchange Rate Uncertainty and the German Labour Market: A Cointegration Application of the ARDL Approach *Ekonomia,* **5,** 1 8-46.

Belke, A., Kaas, L. and Setzer, R. (2004) 'Exchange Rate Volatility and Labour Markets in the CEE Countries', in C.E.P.R. Discussion Papers, CEPR Discussion Papers: 4802.

Belke, A. and Setzer, R. (2003) Costs of Exchange Rate Volatility for Labour Markets: Empirical Evidence from the CEE Economies *Economic and Social Review,* **34,** 3 267-92.

Belke, A. and Setzer, R. (2003) *Exchange rate volatility and employment growth: Empirical evidence from the CEE economies.* CESifo

Ben-David, D., Lumsdaine, R. L. and Papell, D. H. (2003) Unit Roots, Postwar Slowdowns and Long-Run Growth: Evidence from Two Structural Breaks *Empirical Economics,* **28,** 2 303-19.

Bera, A. K. and Higgins, M. L. (1993) ARCH Models: Properties, Estimation and Testing *Journal of Economic Surveys,* **7,** 4 305-66.

Berger, H., Jensen, H. and Schjelderup, G. (2001) To Peg or Not to Peg? A Simple Model of Exchange Rate Regime Choice in Small Economies *Economics Letters,* **73,** 2 161-67.

Bergin, P. (2004) Measuring the Costs of Exchange Rate Volatility *FRBSF Economic Letter,* **2004,** 22 1-3.

Bergin, P. R. (2003) Putting the 'New Open Economy Macroeconomics' to a Test *Journal of International Economics,* **60,** 1 3-34.

Bergin, P. R. and Feenstra, R. C. (2000) Staggered Price Setting, Translog Preferences, and Endogenous Persistence *Journal of Monetary Economics,* **45,** 3 657-80.
Bergin, P. R. and Feenstra, R. C. (2001) Pricing-to-Market, Staggered Contracts, and Real Exchange Rate Persistence *Journal of International Economics,* **54,** 2 333-59.
Berkowitz, J. and Giorgianni, L. (2001) Long-Horizon Exchange Rate Predictability? *Review of Economics and Statistics,* **83,** 1 81-91.
Bernanke, B. S., Gertler, M. and Gilchrist, S. (1999) 'The Financial Accelerator in a Quantitative Business Cycle Framework', in Taylor, J. B. and Woodford, M. e.(Eds) *Handbook of macroeconomics. Volume 1C.*
Betts, C. and Devereux, M. B. (1996) The Exchange Rate in a Model of Pricing-to-Market *European Economic Review,* **40,** 3-5 1007-21.
Betts, C. and Devereux, M. B. (2000) Exchange Rate Dynamics in a Model of Pricing-to-Market *Journal of International Economics,* **50,** 1 215-44.
Bhandari, J. S. (1982) *Exchange rate determination and adjustment.* Praeger: New York, N.Y.
Bhandari, J. S. (1985) *Exchange rate management under uncertainty.* MIT Press: Cambridge, Mass.
Bharucha, N. and Kent, C. (1998) 'Inflation Targeting in a Small Open Economy', in Reserve Bank of Australia Research Discussion Paper: 9807, pp. 28.
Bhawnani, V. and Kadiyala, K. R. (1997) Forecasting Foreign Exchange Rates in Developing Economies *Applied Economics,* **29,** 1 51-62.
Bilson, J. F. O. Macroeconomics Stability and Flexible Exchange Rates *American Economic Review,* **75,** 2 62.
Bilson, J. F. O. (1978a) 'Rational Expectations and the Exchange Rate', in Frenkel, J. A. and Johnson, H.(Eds) *The Economics of Exchange Rates: Selected Studies.* Addison-Wesley: Reading.
Bilson, J. F. O. (1978b) The Monetary Approach to the Exchange Rate: Some Empirical Evidence *International Monetary Fund Staff Papers,* **25,** 1 48-75.
Bilson, J. F. O., Marston, R. C. and National Bureau of Economic Research. (1984) *Exchange rate theory and practice.* University of Chicago Press: Chicago ; London.
Bini-Smaghi, L. (1988) Exchange Rate Variability, Trade and Capital Mobility.
Bini-Smaghi, L. (1991) Exchange Rate Variability and Trade: Why Is It So Difficult to Find Any Empirical Relationship? *Applied Economics,* **23,** 5 927-35.
Bleaney, M. (1992) Comparisons of Real Exchange Rate Volatility across Exchange Rate Systems *Oxford Bulletin of Economics and Statistics,* **54,** 4 557-65.
Bollerslev, T. and Hodrick, R. J. (1992) 'Financial Market Efficiency Tests', in National Bureau of Economic Research, Inc, NBER Working Papers: 4108.
Bollerslev, T. and Wooldridge, J. M. (1992) Quasi-maximum Likelihood Estimation and Inference in Dynamic Models with Time-Varying Covariances *Econometric Reviews,* **11,** 2 143-72.
Bouoiyour, J. and Rey, S. (2005) Exchange Rate Regime, Real Exchange Rate, Trade Flows and Foreign Direct Investments: The Case of Morocco *African*

Development Review/Revue Africaine de Developpement, **17,** 2 302-34.
Broda, C. (2004) Terms of Trade and Exchange Rate Regimes in Developing Countries *Journal of International Economics,* **63,** 1 31-58.
Brodsky, D. A. (1984) Fixed versus Flexible Exchange Rates and the Measurement of Exchange Rate Instability *Journal of International Economics,* **16,** 3-4 295-306.
Brooks, S., Cuthbertson, K. and Mayes, D. G. (1986) *The exchange rate environment.* Groom Helm: London ; Wolfeboro, N.H.
Cavaglia, S. M. F. G. and Wolff, C. C. P. (1996) A Note on the Determinants of Unexpected Exchange Rate Movements *Journal of Banking and Finance,* **20,** 1 179-88.
Cavanagh, C. L., Elliott, G. and Stock, J. (1995) 'Inference in Models with Nearly Integrated Regressors', in Department of Economics, UC San Diego, University of California at.
central bank of kuwait (2006) *market data: exchange rate.* Available at: http://www.cbk.gov.kw/WWW/index.html (Accessed: 12 july 2003).
Chadha, J. S. and Nolan, C. (2002) Inflation and Price Level Targeting in a New Keynesian Model *Manchester School,* **70,** 4 570-95.
Chen, S. (1999) Complex Dynamics of the Real Exchange Rate in an Open Macroeconomic Model *Journal of Macroeconomics,* **21,** 3 493-508.
Cheung, Y.-W. and Lai, K. S. (1993) Long-Run Purchasing Power Parity during the Recent Float *Journal of International Economics,* **34,** 1-2 181-92.
Cheung, Y.-W. and Lai, K. S. (1993) A Fractional Cointegration Analysis of Purchasing Power Parity *Journal of Business and Economic Statistics,* **11,** 1 103-12.
Chinn, M. D. and Meredith, G. (2005) 'Testing Uncovered Interest Parity at Short and Long Horizons during the Post-Bretton Woods Era', in National Bureau of Economic Research, Inc, NBER Working Papers: 11077.
Chipman, J. S. and Kindleberger, C. P. (1980) *Flexible exchange rates and the balance of payments : essays im memory of Egon Sohmen.* North-Holland Pub. Co.: Amsterdam ; New York.
Choudhry, T. and Lawler, P. (1997) The Monetary Models of Exchange Rates: Evidence from the Canadian Float of the 1950s *Journal of Macroeconomics,* **19,** 2 349-62.
Chowdhury, A. R. (1993) Does Exchange Rate Volatility Depress Trade Flows? Evidence from Error-Correction Models *Review of Economics and Statistics,* **75,** 4 700-706.
Clarida, R. and Gali, J. (1994) Sources of Real Exchange-Rate Fluctuations: How Important Are Nominal Shocks? *Carnegie-Rochester Conference Series on Public Policy,* **41,** 1-56.
Cochrane, J. H. (1994) Shocks *Carnegie-Rochester Conference Series on Public Policy,* **41,** 295-364.
Coe, P. J. and Serletis, A. (2002) Bounds Tests of the Theory of Purchasing Power Parity *Journal of Banking and Finance,* **26,** 1 179-99.
Collard, F. and Dellas, H. (2002) Exchange Rate Systems and Macroeconomic Stability *Journal of Monetary Economics,* **49,** 3 571-99.

Collignon, S. (1999) 'Bloc Floating and Exchange Rate Volatility: The Causes and Consequences of Currency Blocs', in Collignon, S., Pisani-Ferry, J. and Park, Y. C. e.(Eds) *Exchange rate policies in emerging Asian countries.*
Conejo, C. and Shields, M. P. (1993) Relative PPP and the Long-Run Terms of Trade for Five Latin American Countries: A Cointegration Approach *Applied Economics,* **25,** 12 1511-15.
Cooray, A. (2003) The Fisher Effect: A Survey *Singapore Economic Review,* **48,** 2 135-50.
Copeland, L. S. (1989) *Exchange rates and international finance.* Addison-Wesley: Wokingham, England ; Reading, Mass.
Copeland, L. S. (2000) *Exchange rates and international finance.* Third edition. Hawlow, U.K.; New York and Toronto: Pearson Education, Financial Times Prentice Hall.
Corbae, D. and Ouliaris, S. (1988) Cointegration and Tests of Purchasing Power Parity *Review of Economics and Statistics,* **70,** 3 508-11.
Corbin, A. (2004) Capital Mobility and Adjustment of the Current Account Imbalances: A Bounds Testing Approach to Cointegration in 12 Countries (1880-2001) *International Journal of Finance and Economics,* **9,** 3 257-76.
Cote, A. (1994) *Exchange rate volatility and trade: A survey,* in
Crosby, M. and Voss, G. (1999) Theoretical Issues in Exchange Rate Determination *Australian Economic Review,* **32,** 2 175-79.
Crystal, A. K. and Sedgwick, R. (1987) *Exchange rates and the open economy.* Wheatsheaf: Brighton.
Cushman, D. O. (1986) Has Exchange Risk Depressed International Trade? The Impact of Third-Country Exchange Risk *Journal of International Money and Finance,* **5,** 3 361-79.
Cushman, D. O. (1988) U.S. Bilateral Trade Flows and Exchange Risk during the Floating Period *Journal of International Economics,* **24,** 3/4 317-30.
Da Silva, S. (1998) 'Overshooting and Foreign Exchange Intervention', in Department of Economics, University of Birmingham, Discussion Papers, pp. 38 pages.
Da Silva, S. (2002) A Brief Overview of the Current State of Exchange Rate Modeling *Revista Brasileira de Economia de Empresas/Brazilian Journal of Business Economics,* **2,** 1 25-32.
Davidson, R. and MacKinnon, J. G. (1993) *Estimation and inference in econometrics.* Oxford; New York; Toronto and Melbourne: Oxford University Press.
Davis, G. K. and Miller, N. C. (1996) Exchange Rate Mean Reversion from Real Shocks within an Intertemporal Equilibrium Model *Journal of International Money and Finance,* **15,** 6 947-67.
De Grauwe, P. (1988) Exchange rate variability and the slowdown in the growth of international trade *IMF Staff Papers,* **35,** 63-84.
De Vita, G. and Abbott, A. (2002) Are Saving and Investment Cointegrated? An ARDL Bounds Testing Approach *Economics Letters,* **77,** 2 293-99.
De Vita, G. and Abbott, A. (2004) Real Exchange Rate Volatility and US Exports: An ARDL Bounds Testing Approach *Economic Issues,* **9,** 1 69-78.

De Vita, G. and Abbott, A. (2004) The Impact of Exchange Rate Volatility on UK Exports to EU Countries *Scottish Journal of Political Economy,* **51,** 1 62-81.
DeJong, D. N. and et al. (1992) The Power Problems of Unit Root Tests in Time Series with Autoregressive Errors *Journal of Econometrics,* **53,** 1-3 323-43.
DeJong, D. N. and Husted, S. (1993) Towards a Reconciliation of the Empirical Evidence on the Monetary Approach to Exchange Rate Determination *Review of Economics and Statistics,* **75,** 1 123-28.
Devereux, M. B. and Engel, C. (1998) 'Fixed vs. Floating Exchange Rates: How Price Setting Affects the Optimal Choice of Exchange-Rate Regime', in National Bureau of Economic Research Working Paper: 6867, pp. 39.
Devereux, M. B. and Lane, P. (2001) 'Exchange Rates and Monetary Policy in Emerging Market Economies', in CEPR Discussion Paper: 2874, pp. not available.
Devereux, M. B. and Lane, P. R. (2003) Understanding Bilateral Exchange Rate Volatility *Journal of International Economics,* **60,** 1 109-32.
Diamandis, P. F., Georgoutsos, D. A. and Kouretas, G. P. (1996) Cointegration Tests of the Monetary Exchange Rate Model: The Canadian-U.S. Dollar, 1970-1994 *International Economic Journal,* **10,** 4 83-97.
Diamandis, P. F., Georgoutsos, D. A. and Kouretas, G. P. (1998) The Monetary Approach to the Exchange Rate: Long-Run Relationships, Identification and Temporal Stability *Journal of Macroeconomics,* **20,** 4 741-66.
Diamandis, P. F. and Kouretas, G. P. (1996) Exchange Rate Determination: Empirical Evidence for the Greek Drachma *Managerial and Decision Economics,* **17,** 3 277-90.
Diamandis, P. F. and Kouretas, G. P. (1996) The Monetary Approach to the Exchange Rate: Long-Run Relationships, Coefficient Restrictions and Temporal Stability of the Greek Drachma *Applied Financial Economics,* **6,** 4 351-62.
Diba, B. T. (1987) A Critique of Variance Bounds Tests for Monetary Exchange Rate Models: A Note *Journal of Money, Credit and Banking,* **19,** 1 104-11.
Dibooglu, S. (1998) The Incidence and Effects of Macroeconomic Disturbances under Alternative Exchange Rate Systems: Evidence since the Classical Gold Standard *International Review of Economics and Finance,* **7,** 2 225-41.
Dibooglu, S. and Kutan, A. M. (2001) Sources of Real Exchange Rate Fluctuations in Transition Economies: The Case of Poland and Hungary *Journal of Comparative Economics,* **29,** 2 257-75.
Dickey, D. A., Bell, W. R. and Miller, R. B. (1986) Unit Roots in Time Series Models: Tests and Implications *American Statistician,* **40,** 1 12.
Dickey, D. A. and Fuller, W. (1979) Distribution of the Estimators for Autoregressive Time Series with a Unit Root *Journal of the American Statistical Association,* **74,** 366 427-31.
Dickey, D. A. and Fuller, W. A. (1981) Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root *Econometrica,* **49,** 4 1057-72.
Dickey, D. A., Hasza, D. P. and Fuller, W. A. (1984) Testing for Unit Roots in Seasonal Time Series *Journal of the American Statistical Association,* **79,**

386 355-67.
Dickey, D. A. and Rossana, R. J. (1994) Cointegrated Time Series: A Guide to Estimation and Hypothesis Testing *Oxford Bulletin of Economics and Statistics,* **56,** 3 325-53.
Dornbusch, R. (1976) Expectations and Exchange Rate Dynamics *Journal of Political Economy,* **84,** 6 1161-76.
Dornbusch, R. (1987) 'Exchange Rate Economics: 1986', in National Bureau of Economic Research, Inc, NBER Working Papers: 2071.
Dornbusch, R. (1988) *Exchange rates and inflation.* MIT Press: Cambridge, Mass.
Dornbusch, R. (1999) 'Bloc Floating and Exchange Rate Volatility: The Causes and Consequences of Currency Blocs: Discussion', in Collignon, S., Pisani-Ferry, J. and Park, Y. C. e.(Eds) *Exchange rate policies in emerging Asian countries.*
Doroodian, K. (1999) Does Exchange Rate Volatility Deter International Trade in Developing Countries? *Journal of Asian Economics,* **10,** 3 465-74.
Driskill, R. (1981) Exchange Rate Overshooting, the Trade Balance, and Rational Expectations *Journal of International Economics,* **11,** 3 361-77.
Driskill, R. A. (1981) Exchange-Rate Dynamics: An Empirical Investigation *Journal of Political Economy,* **89,** 2 357-71.
Duarte, M. (2003) Why Don't Macroeconomic Quantities Respond to Exchange Rate Variability? *Journal of Monetary Economics,* **50,** 4 889-913.
Edwards, S. (1983) Floating Exchange Rates in Less-Developed Countries: A Monetary Analysis of the Peruvian Experience, 1950-54 *Journal of Money, Credit, and Banking,* **15,** 1 73-81.
Edwards, S. (1995) *Capital controls, exchange rates, and monetary policy in the world economy.* Cambridge University Press: Cambridge England.
Edwards, S. (1996) Exchange Rates and the Political Economy of Macroeconomic Discipline *American Economic Review,* **86,** 2 159-63.
Eichengreen, B. and Hausmann, R. (1999) 'Exchange Rates and Financial Fragility', in National Bureau of Economic Research, Inc, NBER Working Papers: 7418.
Eichengreen, B. and Masson, P. (1998) *Exit Strategies: Policy Optionsf or Countries Seeking Greater Exchange Rate Flexibility.* IMF
Elliott, G., Rothenberg, T. J. and Stock, J. H. (1996) Efficient Tests for an Autoregressive Unit Root *Econometrica,* **64,** 4 813-36.
Eltis, W. A. and Sinclair, P. J. N. (1981) *The Money supply and the exchange rate.* Clarendon Press ; Oxford University Press: Oxford, New York.
Enders, W. (1988) ARIMA and Cointegration Tests of PPP under Fixed and Flexible Exchange Rate Regimes *Review of Economics and Statistics,* **70,** 3 504-08.
Enders, W. (1995) *Applied econometric time series.* Series in Probability and Mathematical Statistics. New York; Chichester, U.K. and Toronto: Wiley.
Engel, C. (1999) Accounting for U.S. Real Exchange Rate Changes *Journal of Political Economy,* **107,** 3 507-38.
Engel, C. (2002) 'Expenditure Switching and Exchange Rate Policy', in National Bureau of Economic Research, Inc, NBER Working Papers: 9016.

Engel, C. and Hakkio, C. S. (1993) Exchange Rate Regimes and Volatility *Federal Reserve Bank of Kansas City Economic Review,* **78,** 3 43-58.

Engel, C. and West, K. D. (2004) 'Exchange Rates and Fundamentals', in National Bureau of Economic Research, Inc, NBER Working Papers: 10723.

Engle, R. F. (1982) Autoregressive Conditional Heteroscedasticity with Estimates of the Variance of United Kingdom Inflation *Econometrica,* **50,** 4 987-1007.

Engle, R. F. (1995) *ARCH : selected readings.* Oxford University Press: Oxford England.

Engle, R. F. and Granger, C. W. J. (1987) Co-integration and Error Correction: Representation, Estimation, and Testing *Econometrica,* **55,** 2 251-76.

Engle, R. F., Lilien, D. M. and Robins, R. P. (1987) Estimating Time Varying Risk Premia in the Term Structure: The Arch-M Model *Econometrica,* **55,** 2 391-407.

Engle, R. F., Lilien, D. M. and Robins, R. P. (1995) 'Estimating Time-Varying Risk Premia in the Term Structure: The ARCH-M Model', in Engle, R. F. e.(Ed), *ARCH: Selected readings.*

Engle, R. F. and Patton, A. J. (2001) What Good Is a Volatility Model? *Quantitative Finance,* **1,** 2 237-45.

Engle, R. F. and Patton, A. J. (2004) Impacts of Trades in an Error-Correction Model of Quote Prices *Journal of Financial Markets,* **7,** 1 1-25.

Erlat, H. and Erlat, G. (1998) Permanent and Transitory Shocks on Real and Nominal Exchange Rates in Turkey during the Post-1980 Period *Atlantic Economic Journal,* **26,** 4 379-96.

Fama, E. F. and Roll, R. (1971) Parameter Estimates for Symmetric Stable Distributions *Journal of the American Statistical Association,* **66,** 334 331-38.

Felipe, J. (2002) Unemployment and Profitability in Spain *Journal of Income Distribution,* **11,** 1-2 7-33.

Fendel, R. (2002) Open Economy Macroeconomics in the Post Mundell-Fleming Era *Jahrbuc,* **53,** 1 53-87.

Fernandez, D. G. (1997) Breaking Trends and the Money-Output Correlation *Review of Economics and Statistics,* **79,** 4 674-79.

Fisher, E. O. N. and Park, J. Y. (1991) Testing Purchasing Power Parity under the Null Hypothesis of Co-integration *Economic Journal,* **101,** 409 1476-84.

Flood, R. P. and Rose, A. K. (1995) Fixing Exchange Rates: A Virtual Quest for Fundamentals *Journal of Monetary Economics,* **36,** 1 3-37.

Flood, R. P. and Rose, A. K. (1999) Understanding Exchange Rate Volatility without the Contrivance of Macroeconomics *Economic Journal,* **109,** 459 F660-72.

Flood, R. P. and Rose, A. K. (2002) Uncovered Interest Parity in Crisis *IMF Staff Papers,* **49,** 2 252-66.

Francis, B., Hasan, I. and Lothian, J. R. (2001) The Monetary Approach to Exchange Rates and the Behaviour of the Canadian Dollar over the Long Run *Applied Financial Economics,* **11,** 5 475-81.

Frankel, J. A. (1979) On the Mark: A Theory of Floating Exchange Rates Based on Real Interest Differentials *American Economic Review,* **69,** 4 610-22.

Frankel, J. A. and Froot, K. A. (1990) Chartists, Fundamentalists, and Trading in the Foreign Exchange Market *American Economic Review,* **80,** 2 181-85.
Frankel, J. A. and Rose, A. K. (1996) A Panel Project on Purchasing Power Parity: Mean Reversion within and between Countries *Journal of International Economics,* **40,** 1-2 209-24.
Franses, P. H. and McAleer, M. (2002) Financial Volatility: An Introduction *Journal of Applied Econometrics,* **17,** 5 419-24.
Frenkel, J. A. (1976) A Monetary Approach to the Exchange Rate: Doctrinal Aspects and Empirical Evidence *Scandinavian Journal of Economics,* **78,** 2 200-224.
Frenkel, J. A. and Johnson, H. G. (1978) *The Economics of exchange rates : selected studies.* Addison-Wesley Pub. Co.: Reading, Mass.
Frenkel, J. A. and National Bureau of Economic Research. (1983) *Exchange rates and international macroeconomics.* University of Chicago Press: Chicago ; London.
Frenkel, J. A. and Razin, A. (1996) *Fiscal policies and growth in the world economy.* Third edition. With the collaboration of Chi-Wa Yuen. Cambridge and London: MIT Press.
Friedman, M. (1953) 'The Case for Flexible Exchange Rates', in *Essays in positive economics.* University of Chicago Press: Chicago, pp. 157-203.
Frommel, M., MacDonald, R. and Menkhoff, L. (2005) Do Fundamentals Matter for the D-Mark/Euro-Dollar? A Regime Switching Approach *Global Finance Journal,* **15,** 3 321-35.
Frommel, M., MacDonald, R. and Menkhoff, L. (2005) Markov Switching Regimes in a Monetary Exchange Rate Model *Economic Modelling,* **22,** 3 485-502.
Frommel, M. and Menkhoff, L. (2003) Increasing Exchange Rate Volatility during the Recent Float *Applied Financial Economics,* **13,** 12 857-63.
Froot, K. A. and Thaler, R. H. (1990) Foreign Exchange *Journal of Economic Perspectives,* **4,** 3 179-92.
Fry, M. J. (1976) A Monetary Approach to Afghanistan's Flexible Exchange Rate *Journal of Money, Credit, and Banking,* **8,** 2 219-25.
Gali, J. and Monacelli, T. (2005) Monetary Policy and Exchange Rate Volatility in a Small Open Economy *Review of Economic Studies,* **72,** 3 707-34.
Garcia-Cebro, J. A. and Varela-Santamaria, R. (2004) Vertically Integrated North-South Trade and the Redux Model *Review of International Economics,* **12,** 4 643-61.
Ghatak, A. (1997) Unit Roots and Structural Breaks: The Case of India 1900-1988 *Journal of Applied Statistics,* **24,** 3 289-300.
Ghatak, S. and Siddiki, J. U. (2001) The Use of the ARDL Approach in Estimating Virtual Exchange Rates in India *Journal of Applied Statistics,* **28,** 5 573-83.
Godfrey, L. G. (1978) Testing against General Autoregressive and Moving Average Error Models When the Regressors Include Lagged Dependent Variables *Econometrica,* **46,** 6 1293-1301.
Goldberg, M. D. and Frydman, R. (2001) Macroeconomic Fundamentals and the DM/$ Exchange Rate: Temporal Instability and the Monetary Model *International Journal of Finance and Economics,* **6,** 4 421-35.

Gotur, P. (1985) Effects of Exchange Rate Volatility on Trade: Some Further Evidence *International Monetary Fund Staff Papers,* **32,** 3 475-512.
Granger, C. W. J. and Newbold, P. (1974) Spurious Regressions in Econometrics *Journal of Econometrics,* **2,** 2 111-20.
Granger, C. W. J. and Newbold, P. (1986) *Forecasting economic time series.* Second edition Economic Theory, Econometrics, and Mathematical Economics series. Orlando, Fla.; London; Sydney and Toronto: Harcourt, Brace, Jovanovich; Academic Press.
Grauwe, P. d., Dewachter, H. and Embrechts, M. (1993) *Exchange rate theory : chaotic models of foreign exchange markets.* Blackwell: Oxford, UK ; Cambridge, Mass.
Grauwe, P. d., Fratianni, M. and Nabli, M. K. (1985) *Exchange rates, money, and output : the European experience.* Macmillan: London.
Greene, W. H. (2003) *Econometric analysis.* Prentice Hall: Upper Saddle River, N.J., Great Britain.
Gujarati, D. N. (2003) *Basic econometrics.* McGraw-Hill: Boston, Mass. ; London.
Gurney, A. (1988) The Exchange Rate, Interest Rates and the Current Balance in a Forward-Looking Model *National Institute Economic Review,* **0,** 125 40-55.
Gylfason, T. (2002) The Real Exchange Rate Always Floats *Australian Economic Papers,* **41,** 4 369-81.
Hairault, J.-O. and Sopraseuth, T. (2004) *Exchange rate dynamics : a new open economy macroeconomics perspective.* Routledge: London.
Hall, S. G. (1989) Maximum Likelihood Estimation of Cointegration Vectors: An Example of the Johansen Procedure *Oxford Bulletin of Economics and Statistics,* **51,** 2 213-18.
Hamilton, J. D. (1994) *Time series analysis.* Princeton: Princeton University Press.
Harris, R. I. D. (1995) *Using cointegration analysis in econometric modelling.* Harvester Wheatsheaf Prentice Hall: Hemel Hempstead, England.
Harris, R. I. D. and Sollis, R. (2003) *Applied time series modelling and forecasting.* J. Wiley: Chichester, West Sussex, England ; Hoboken, NJ.
Harvey, D. I. and Leybourne, S. J. (2005) On Testing for Unit Roots and the Initial Observation *Econometrics Journal,* **8,** 1 97-111.
Harvey, D. I. and Mills, T. C. (2004) Tests for Stationarity in Series with Endogenously Determined Structural Change *Oxford Bulletin of Economics and Statistics,* **66,** 5 863-94.
Harvey, J. T. (1996) Long-Term Exchange Rate Movements: The Role of the Fundamentals in Neoclassical Models of Exchange Rates *Journal of Economic Issues,* **30,** 2 509-16.
Harvey, J. T. (1996) Orthodox Approaches to Exchange Rate Determination: A Survey *Journal of Post Keynesian Economics,* **18,** 4 567-83.
Harvey, J. T. (2001) Exchange Rate Theory and "the Fundamentals." *Journal of Post Keynesian Economics,* **24,** 1 3-15.
Hau, H. (2000) Exchange Rate Determination: The Role of Factor Price Rigidities and Nontradeables *Journal of International Economics,* **50,** 2 421-47.
Hau, H. (2002) Real Exchange Rate Volatility and Economic Openness: Theory

and Evidence *Journal of Money, Credit, and Banking,* **34,** 3 611-30.
Hausmann, R., Panizza, U. and Stein, E. (2001) Why Do Countries Float the Way They Float? *Journal of Development Economics,* **66,** 2 387-414.
Haynes, S. E. and Stone, J. A. (1981) On the Mark: Comment [On the Mark: A Theory of Floating Exchange Rates Based on Real Interest Differential] *American Economic Review,* **71,** 5 1060-67.
Haynes, S. E. and Stone, J. A. (1982) Spurious Tests and Sign Reversals in International Economics *Southern Economic Journal,* **48,** 4 868-76.
Hendry, D. F. (1995) *Dynamic econometrics.* Oxford University Press: Oxford.
Hodgman, D. R. and Wood, G. E. (1987) *Monetary and exchange rate policy.* Macmillan in association with Centre for Banking and International Finance the City University Business School: London.
Hodrick, R. J. (1978) 'An Empirical Analysis of the Monetary Approach to the Determination of the Exchange Rate', in Frenkel, J. A. and Johnson, H.(Eds) *The Economics of Exchange Rates: Selected Studies.* Addison-Wesley: Reading.
Hook, L. S. and Boon, T. H. (2000) 'Real Exchange Rate Volatility and the Malaysian Exports to Its Major Trading Partners', in Habibullah, M. S. e.(Ed), *ASEAN in an interdependent world: Studies on trade and finance.*
Hooper, P. and Kohlhagen, S. W. (1978) The Effect of Exchange Rate Uncertainty on the Prices and Volume of International Trade *Journal of International Economics,* **8,** 4 483-511.
Hooper, P. and Morton, J. (1992) 'Fluctuations in the Dollar: A Model of Nominal and Real Exchange Rate Determination', in MacDonald, R. and Taylor, M. P. e.(Eds) *Exchange rate economics. Volume 1.*
Horvath, J. (2003) *Optimum currency area theory : a selective review.* Bofit: Helsinki.
Horvath, R. (2005) Exchange rate variability, pressures and optimum currency area criteria: some empirical evidence from the 1990s *Applied Economics Letters,* **12,** 15 919-922.
Horvath, R. and Komarek, L. (2003) 'Optimum currency area indices: Evidence from 1990s', in, pp. 1-17.
Hsieh, D. A. (1988) The Statistical Properties of Daily Foreign Exchange Rates: 1974-1983 *Journal of International Economics,* **24,** 1/2 129-45.
Huang, R. D. (1981) The Monetary Approach to Exchange Rate in an Efficient Foreign Exchange Market: Tests Based on Volatility *Journal of Finance,* **36,** 1 31-41.
Hwang, J.-K. (2001) Dynamic Forecasting of Monetary Exchange Rate Models: Evidence from Cointegration *International Advances in Economic Research,* **7,** 1 51-64.
Hwang, J.-K. (2003) Dynamic Forecasting of Sticky-Price Monetary Exchange Rate Model *Atlantic Economic Journal,* **31,** 1 103-14.
In, F. and Sugema, I. (1995) Testing Purchasing Power Parity in a Multivariate Cointegrating Framework *Applied Economics,* **27,** 9 891-99.
Iqbal, Z. and Erbas, S. N. (1997) 'External Stability Under Alternative Nominal Exchange Rate Anchors: An Application to the GCC Countries', in

International Monetary Fund, IMF Working Papers: 97/8.
Isard, P. (1995) *Exchange rate economics.* Surveys of Economic Literature. Cambridge; New York and Melbourne: Cambridge University Press.
Islam, M. Q. (2001) Structural Break, Unit Root, and the Causality between Government Expenditures and Revenues *Applied Economics Letters,* **8,** 8 565-67.
Islam, M. Q. (2004) The Long Run Relationship between Openness and Government Size: Evidence from Bounds Test *Applied Economics,* **36,** 9 995-1000.
Jadresic, E., Masson, P. and Mauro, P. (2001) Exchange Rate Regimes of Developing Countries: Global Context and Individual Choices *Journal of the Japanese and International Economies,* **15,** 1 68-101.
Jansen, D. W. (1989) Does Inflation Uncertainty Affect Output Growth? Further Evidence *Federal Reserve Bank of St. Louis Review,* **71,** 4 43-54.
jeong, J.-G. (2000) What Drives Exchange Rates?: The Case of the Yen/Dolllar Rate *Multinational Business Review,* 31-6.
Johansen, S. (1991) Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models *Econometrica,* **59,** 6 1551-80.
Johansen, S. and Juselius, K. (1990) Maximum Likelihood Estimation and Inference on Cointegration--With Applications to the Demand for Money *Oxford Bulletin of Economics and Statistics,* **52,** 2 169-210.
Johansen, S. r. (1995) *Likelihood-based inference in cointegrated vector autoregressive models.* Oxford University Press: Oxford.
Johnson, C. (1990) *Changing exchange rate systems.* Pinter Publishers: London ; New York.
Johnston, J. and DiNardo, J. (1997) *Econometric methods.* McGraw-Hill: New York ; London.
Kanas, A. (1998) Testing for a Unit Root in ERM Exchange Rates in the Presence of Structural Breaks: Evidence from the Bootstrap *Applied Economics Letters,* **5,** 7 407-10.
Karras, G., Lee, J. M. and Stokes, H. (2005) Sources of Exchange-Rate Volatility: Impulses or Propagation? *International Review of Economics and Finance,* **14,** 2 213-26.
Kearney, C. and MacDonald, R. (1990) Rational Expectations, Bubbles and Monetary Models of the Exchange Rate: The Australian/U.S. Dollar Rate during the Recent Float *Australian Economic Papers,* **29,** 54 1-20.
Kempa, B. (2005) An Oversimplified Inquiry into the Sources of Exchange Rate Variability *Economic Modelling,* **22,** 3 439-58.
Kenen, P. B. (1994) *Exchange rates and the monetary system: Selected essays of Peter B. Kenen.* Economists of the Twentieth Century series. Aldershot, U.K.: Elgar; distributed in the U.S. by Ashgate, Brookfield, Vt.
Kenen, P. B. and Rodrik, D. (1986) Measuring and Analyzing the Effects of Short-term Volatility in Real Exchange Rates *Review of Economics and Statistics,* **68,** 2 311-15.
Kenen, P. B. and Rodrik, D. (1994) 'Measuring and Analyzing the Effects of Short-Term Volatility in Real Exchange Rates', in Kenen, P. B.(Ed),

Exchange rates and the monetary system: Selected essays of Peter B. Kenen.
Khan, H. (2001) Price Stickiness, Inflation, and Persistence in Real Exchange Rate Fluctuations: Cross-Country Results *Economics Letters,* **71,** 2 247-53.
Khan, M. A. and Sajjid, M. Z. (2005) The Exchange Rates and Monetary Dynamics in Pakistan: An Autoregressive Distributed Lag (ARDL) Approach *Lahore Journal of Economics,* **10,** 2 87-99.
Kholdy, S. and Sohrabian, A. (1995) Testing for the Relationship Between Nominal Exchange Rates and Economic Fundamentals *Global Finance Journal,* **6,** 2 121-134.
Kim, Y. (1990) Purchasing Power Parity in the Long Run: A Cointegration Approach *Journal of Money, Credit, and Banking,* **22,** 4 491-503.
Kimball, M. S. (1995) The Quantitative Analytics of the Basic Neomonetarist Model *Journal of Money, Credit, and Banking,* **27,** 4 1241-77.
Klein, M. W. (1990) Sectoral Efforts of Exchange Rate Volatility on United States Exports *Journal of International Money and Finance,* **9,** 3 299-308.
Knight, F. (1921) *Risk, uncertainty and profit.* Mifflin: Boston.
Koop, G. (2003) *Bayesian econometrics.* J. Wiley: Hoboken, N.J.
Koop, G. (2005) *Analysis of economic data.* John Wiley & Sons: Hoboken, NJ.
Koray, F. and Lastrapes, W. D. (1989) Real Exchange Rate Volatility and U.S. Bilateral Trade: A VAR Approach *Review of Economics and Statistics,* **71,** 4 708-12.
Koronowski, A. (2001) Monetary Impact on Exchange Rate Volatility *Economic and Financial Modelling,* **8,** 3 103-25.
Kumar Narayan, P. (2005) The Relationship between Saving and Investment for Japan *Japan and the World Economy,* **17,** 3 293-309.
laabas, B. and Limam, I. (2002) Are GCC Countries Ready for Currency Union? *Arab Planning Institute,* 1-34.
Lafrance, R. and van Norden, S. (1995) Exchange Rate Fundamentals and the Canadian Dollar *Bank of Canada Review,* 17-34.
Lan, Y. (2003) *The Long-Term Behaviour of Exchange Rates.* thesis. The University of Western Australia.
Lane, P. R. (2001) The New Open Economy Macroeconomics: A Survey *Journal of International Economics,* **54,** 2 235-66.
Lanyi, A. and Suss, E. C. (1982) Exchange Rate Variability: Alternative Measures and Interpretation *International Monetary Fund Staff Papers,* **29,** 4 527-60.
Larrain B, F. and Tavares, J. (2003) Regional Currencies versus Dollarization: Options for Asia and the Americas *Journal of Policy Reform,* **6,** 1 35-49.
Lastrapes, W. D. (1989) Exchange Rate Volatility and U.S. Monetary Policy: An ARCH Application *Journal of Money, Credit, and Banking,* **21,** 1 66-77.
Lastrapes, W. D. (1992) Sources of Fluctuations in Real and Nominal Exchange Rates *Review of Economics and Statistics,* **74,** 3 530-39.
Lastrapes, W. D. and Koray, F. (1990) Exchange Rate Volatility and U.S. Multilateral Trade Flows *Journal of Macroeconomics,* **12,** 3 341-62.
Leitemo, K., Roisland, O. and Torvik, R. (2002) Time Inconsistency and the Exchange Rate Channel of Monetary Policy *Scandinavian Journal of*

Economics, **104**, 3 391-97.
LeRoy, S. F. and Singell, L. D., Jr. (1987) Knight on Risk and Uncertainty *Journal of Political Economy*, **95**, 2 394-406.
Levine, R., Beck, T. and Loayza, N. (1999) 'Financial Intermediation and Growth: Causality and Causes', in The World Bank, Policy Research Working Paper Series: 2059.
Lopez, J. H. (1997) The Power of the ADF Test *Economics Letters*, **57**, 1 5-10.
Lutkepohl, H. (1993) *Introduction to multiple time series analysis.* Springer-Verlag: Berlin ; New York.
MacDonald, R. (1984) The Monetary Approach to the Exchange Rate and the 1920s' Experience with Floating Exchange Rates *Journal of Economic Studies*, **11**, 3 14-25.
MacDonald, R. (1984) The Monetary Approach to the Exchange Rate Revisited *Applied Economics*, **16**, 5 771-82.
MacDonald, R. (1988) *Floating exchange rates: Theories and evidence.* London; Boston; Sydney and Wellington, New Zealand: Unwin Hyman.
MacDonald, R. (1999) Exchange Rate Behaviour: Are Fundamentals Important? *Economic Journal*, **109**, 459 F673-91.
MacDonald, R. and Taylor, M. P. (1991) The Monetary Approach to the Exchange Rate: Long-Run Relationships and Coefficient Restrictions *Economics Letters*, **37**, 2 179-85.
MacDonald, R. and Taylor, M. P. (1992) Exchange Rate Economics: A Survey *International Monetary Fund Staff Papers*, **39**, 1 1-57.
MacDonald, R. and Taylor, M. P. (1994) Reexamining the Monetary Approach to the Exchange Rate: The Dollar-Franc, 1976-90 *Applied Financial Economics*, **4**, 6 423-29.
Maddala, G. S. (2001) *Introduction to econometrics.* John Wiley: New York ; Chichester.
Maddala, G. S., Rao, C. R. and Vinod, H. D. (1993) *Econometrics.* North-Holland: Amsterdam.
Maghayereh, A. (2003) Causal Relations among Stock Prices and Macroeconomic Variables in the Small , Open Economy of Jordan *Economics and Administration*, **17**, 2 3-12.
Mah, J. S. (2000) An Empirical Examination of the Disaggregated Import Demand of Korea--The Case of Information Technology Products *Journal of Asian Economics*, **11**, 2 237-44.
Makrydakis, S. (1998) Testing the Long-Run Validity of the Monetary Approach to the Exchange Rate: The Won-US Dollar Case *Applied Economics Letters*, **5**, 8 507-11.
Mandelbrot, B. (1997) 'The Variation of Certain Speculative Prices', in Malliaris, A. G. e.(Ed), *Futures markets. Volume 2.*
Mankiw, N. G. and Summers, L. H. (1986) Money Demand and the Effects of Fiscal Policies *Journal of Money, Credit, and Banking*, **18**, 4 415-29.
Manning, L. M. and Andrianacos, D. (1993) Dollar Movements and Inflation: A Cointegration Analysis *Applied Economics*, **25**, 12 1483-88.
Mark, N. C. (1995) Exchange Rates and Fundamentals: Evidence on

Long-Horizon Predictability *American Economic Review,* **85,** 1 201-18.
Mark, N. C. and Sul, D. (2001) Nominal Exchange Rates and Monetary Fundamentals: Evidence from a Small Post-Bretton Woods Panel *Journal of International Economics,* **53,** 1 29-52.
McKenzie, M. D. (1997) ARCH Modelling of Australian Bilateral Exchange Rate Data *Applied Financial Economics,* **7,** 2 147-64.
McKenzie, M. D. (1998) The Impact of Exchange Rate Volatility on Australian Trade Flows *Journal of International Financial Markets, Institutions and Money,* **8,** 1 21-38.
McKenzie, M. D. (1999) The Impact of Exchange Rate Volatility on International Trade Flows *Journal of Economic Surveys,* **13,** 1 71-106.
McKenzie, M. D. and Brooks, R. D. (1997) The Impact of Exchange Rate Volatility on German-US Trade Flows *Journal of International Financial Markets, Institutions and Money,* **7,** 1 73-87.
McKinnon, R. I. (1963) Optimum Currency Areas *American Economic Review,* **53,** 4 717.
McKinnon, R. I. (1976) Floating foreign exchange rates 1973-1974: The emperor's new clothes *Carnegie-Rochester Conference Series on Public Policy,* **3,** 79-114.
McKinnon, R. I. (2002) Optimum Currency Areas and the European Experience *Economics of Transition,* **10,** 2 343-64.
McNown, R. and Wallace, M. (1989) Co-integration Tests for Long Run Equilibrium in the Monetary Exchange Rate Model *Economics Letters,* **31,** 3 263-67.
Medhora, R. (1990) The Effect of Exchange Rate Variability on Trade: The Case of the West African Monetary Union's Imports *World Development,* **18,** 2 313-24.
Meese, R. A. and Rogoff, K. (1983) Empirical Exchange Rate Models of the Seventies: Do They Fit Out of Sample? *Journal of International Economics,* **14,** 1-2 3-24.
Meese, R. A. and Rogoff, K. (1988) Was It Real? The Exchange Rate-Interest Differential Relation over the Modern Floating-Rate Period *Journal of Finance,* **43,** 4 933-48.
Miyakoshi, T. (2000) The Monetary Approach to the Exchange Rate: Empirical Observations from Korea *Applied Economics Letters,* **7,** 12 791-94.
Moersch, M. and Nautz, D. (2001) A Note on Testing the Monetary Model of the Exchange Rate *Applied Financial Economics,* **11,** 3 261-68.
Mongelli, F. P. (2002) '"New" views on the optimum currency area theory : what is EMU telling us?' in European Central Bank, Working Paper Series: 138.
Moosa, I. A. (1994) The Monetary Model of Exchange Rates Revisited *Applied Financial Economics,* **4,** 4 279-87.
Morana, C. and Beltratti, A. (2002) The Effects of the Introduction of the Euro on the Volatility of European Stock Markets *Journal of Banking and Finance,* **26,** 10 2047-64.
Mundell, R. A. (1961) A Theory of Optimum Currency Areas *American Economic Review,* **51,** 4 657.
Mundell, R. A. (1963) Capital Mobility and Stabilization Policy under Fixed and

Flexible Exchange Rates *The Canadian Journal of Economics and Political Science,* **29,** 4 475-485.
Mussa, M. (1976) The Exchange Rate, the Balance of Payments and Monetary and Fiscal Policy under a Regime of Controlled Floating *Scandinavian Journal of Economics,* **78,** 2 229-48.
Mussa, M. (1992) 'Empirical Regularities in the Behavior of Exchange Rates and Theories of the Foreign Exchange Market', in MacDonald, R. and Taylor, M. P. e.(Eds) *Exchange rate economics. Volume 1.*
Mutoh, T. (1987) On the Relation between Monetary Interdependence and Exchange Rate Volatility *Journal of the Japanese and International Economies,* **1,** 4 351-72.
Najand, M. and Bond, C. (2000) Structural Models of Exchange Rate Determination *Journal of Multinational Financial Management,* **10,** 1 15-27.
Narayan, P. K. (2004) Fiji's Tourism Demand: The ARDL Approach to Cointegration *Tourism Economics,* **10,** 2 193-206.
Narayan, P. K. and Narayan, S. (2005) Estimating Income and Price Elasticities of Imports for Fiji in a Cointegration Framework *Economic Modelling,* **22,** 3 423-38.
Narayan, P. K. and Narayan, S. (2005) Are Exports and Imports Cointegrated? Evidence from 22 Least Developed Countries *Applied Economics Letters,* **12,** 6 375-78.
Narayan, P. K. and Smyth, R. (2005) Trade Liberalization and Economic Growth in Fiji: An Empirical Assessment Using the ARDL Approach *Journal of the Asia Pacific Economy,* **10,** 1 96-115.
Narayan, P. K. and Smyth, R. (2005) The Residential Demand for Electricity in Australia: an Application of the Bounds Testing Approach to Cointegration *Energy Policy,* **33,** 467-74.
Neely, C. J. and Sarno, L. (2002) How Well Do Monetary Fundamentals Forecast Exchange Rates? *Federal Reserve Bank of St. Louis Review,* **84,** 5 51-74.
Nelson, D. B. (1991) Conditional Heteroskedasticity in Asset Returns: A New Approach *Econometrica,* **59,** 2 347-70.
Nelson, D. B. (1995) 'Stationarity and Persistence in the GARCH (1,1) Model', in Engle, R. F. e.(Ed), *ARCH: Selected readings.*
Newbold, P. and Granger, C. W. J. (1974) Experience with Forecasting Univariate Time Series and the Combination of Forecasts *Journal of the Royal Statistical Society, Series A,* **137*PD,** 131-46.
Ng, S. and Perron, P. (2001) Lag Length Selection and the Construction of Unit Root Tests with Good Size and Power *Econometrica,* **69,** 6 1519-54.
Nieh, C.-C. and Wang, Y.-S. (2005) ARDL Approach to the Exchange Rate Overshooting in Taiwan *Review of Quantitative Finance and Accounting,* **25,** 1 55-71.
Obstfeld, M. (1982) Aggregate Spending and the Terms of Trade: Is There a Laursen-Metzler Effect? *Quarterly Journal of Economics,* **97,** 2 251-70.
Obstfeld, M. (2002) 'Exchange Rates and Adjustment: Perspectives from the New Open Economy Macroeconomics', in National Bureau of Economic Research, Inc, NBER Working Papers: 9118.

Obstfeld, M. and Rogoff, K. (1995) Exchange Rate Dynamics Redux *Journal of Political Economy,* **103,** 3 624-60.
Obstfeld, M. and Rogoff, K. (1998) 'Risk and Exchange Rates', in National Bureau of Economic Research, Inc, NBER Working Papers: 6694.
Obstfeld, M. and Rogoff, K. (2000) New Directions for Stochastic Open Economy Models *Journal of International Economics,* **50,** 1 117-53.
Odedokun, M. O. (1996) Monetary Model of Black Market Exchange Rate Determination: Evidence from African Countries *Journal of Economic Studies,* **23,** 4 31-49.
Odedokun, M. O. (1997) The Monetary Approach to Analysing Floating Exchange Rate Behaviour in Developing Countries: Evidence from Sub-Saharan African Countries, 1986-1992 *Journal of Development Economics,* **52,** 2 463-81.
Officer, L. H. (1978) The Relationship between Absolute and Relative Purchasing Power Parity *Review of Economics and Statistics,* **6,** 4 562-68.
Officer, L. H. (1982) *Purchasing power parity and exchange rates : theory, evidence and relevance.* JAI Press: Greenwich, Conn.
Oh, K.-Y. (1996) Purchasing Power Parity and Unit Root Tests Using Panel Data *Journal of International Money and Finance,* **15,** 3 405-18.
Pagan, A. R. and Schwert, G. W. (1990) Alternative Models for Conditional Stock Volatility *Journal of Econometrics,* **45,** 1-2 267-90.
Pagan, A. R. and Schwert, G. W. (1990) Testing for Covariance Stationarity in Stock Market Data *Economics Letters,* **33,** 2 165-70.
Papell, D. H. (1997) Searching for Stationarity: Purchasing Power Parity under the Current Float *Journal of International Economics,* **43,** 3-4 313-32.
Papke, L. E. and Wooldridge, J. M. (2005) A Computational Trick for Delta-Method Standard Errors *Economics Letters,* **86,** 3 413-17.
Patterson, K. (2000) *An introduction to applied econometrics: A time series approach.* New York: St. Martin's Press.
Patterson, K. D. (2000) *An introduction to applied econometrics : a time series approach.* Macmillan: Basingstoke, England.
Pattichis, C. A. (1999) Price and Income Elasticities of Disaggregated Import Demand: Results from UECMs and an Application *Applied Economics,* **31,** 9 1061-71.
Payne, J. E. (2003) Post Stabilization Estimates of Money Demand in Croatia: Error Correction Model Using the Bounds Testing Approach *Applied Economics,* **35,** 16 1723-27.
Perron, P. (1989) The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis *Econometrica,* **57,** 6 1361-1401.
Perron, P. (1997) Further Evidence on Breaking Trend Functions in Macroeconomic Variables *Journal of Econometrics,* **80,** 2 355-85.
Perron, P. and Ng, S. (1996) Useful Modifications to Some Unit Root Tests with Dependent Errors and Their Local Asymptotic Properties *Review of Economic Studies,* **63,** 3 435-63.
Perron, P. and Rodriguez, G. (2003) GLS Detrending, Efficient Unit Root Tests and Structural Change *Journal of Econometrics,* **115,** 1 1-27.

Pesaran, M. H. and Shin, Y. (1998) 'An Autoregressive Distributed-Lag Modelling Approach to Cointegration Analysis', in Strom, S. e.(Ed), *Econometrics and economic theory in the twentieth century: The Ragnar Frisch Centennial Symposium.*

Pesaran, M. H., Shin, Y. and Smith, R. J. (2001) Bounds Testing Approaches to the Analysis of Level Relationships *Journal of Applied Econometrics,* **16,** 3 289-326.

Pesaran, M. H. and Wickens, M. (1995) *Handbook of applied econometrics.* Blackwell: Oxford, UK ; Cambridge, Mass.

Phillips, P. C. B. (1986) Understanding Spurious Regressions in Econometrics *Journal of Econometrics,* **33,** 3 311-40.

Pilbeam, K. (1990) *Exchange rate management : theory and evidence.* Macmillan: Basingstoke.

Pilbeam, K. (1995) Exchange Rate Models and Exchange Rate Expectations: An Empirical Investigation *Applied Economics,* **27,** 11 1009-15.

Poirson, H. (2001) 'How Do Countries Choose Their Exchange Rate Regime?' in International Monetary Fund, IMF Working Papers: 01/46.

Pozo, S. (1992) Are Flexible Exchange Rates Really More Volatile? Evidence from the Early 1900s *Applied Economics,* **24,** 11 1213-18.

Pozo, S. (1992) Conditional Exchange-Rate Volatility and the Volume of International Trade: Evidence from the Early 1900s *Review of Economics and Statistics,* **74,** 2 325-29.

Putnam, B. H. and Woodbury, J. R. (1980) Exchange Rate Stability and Monetary Policy *Review of Business and Economic Research,* **15,** 2 1-10.

Rahmatsyah, T., Rajaguru, G. and Siregar, R. Y. (2002) Exchange-Rate Volatility, Trade and "Fixing for Life" in Thailand *Japan and the World Economy,* **14,** 4 445-70.

Ran, J. (2002) Volatility of Fundamental Variables under Different Exchange Rate System--A Simple Test for Hong Kong *Applied Economics,* **34,** 2 253-58.

Rana, P. B. (1981) Exchange Rate Risk under Generalized Floating: Eight Asian Countries *Journal of International Economics,* **11,** 4 459-66.

Rana, P. B. (1984) Fixed versus Flexible Exchange Rates and the Measurement of Exchange Rate Instability: Comment *Journal of International Economics,* **16,** 3-4 307-10.

Rapach, D. E. and Weber, C. E. (2004) Are Real Interest Rates Really Nonstationary? New Evidence from Tests with Good Size and Power *Journal of Macroeconomics,* **26,** 3 409-30.

Rapach, D. E. and Wohar, M. E. (2002) Testing the Monetary Model of Exchange Rate Determination: New Evidence from a Century of Data *Journal of International Economics,* **58,** 2 359-85.

Rapach, D. E. and Wohar, M. E. (2004) Testing the Monetary Model of Exchange Rate Determination: A Closer Look at Panels *Journal of International Money and Finance,* **23,** 6 867-95.

Richardson, D. H. and Wu, M. T. C. (1988) A Short-run Monetary Model of Exchange Rate Determination: Stability Tests and Forecasting *Empirical Economics,* **13,** 1 1-15.

Rizzo, J.-M. (1998) The Economic Determinants of the Choice of an Exchange Rate Regime: A Probit Analysis *Economics Letters,* **59,** 3 283-87.
Rogers, J. H. (1999) Monetary Shocks and Real Exchange Rates *Journal of International Economics,* **49,** 2 269-88.
Rogers, J. H. and Jenkins, M. (1995) Haircuts or Hysteresis? Sources of Movements in Real Exchange Rates *Journal of International Economics,* **38,** 3-4 339-60.
Rose, A. K. (1994) Are Exchange Rates Macroeconomic Phenomena? *Federal Reserve Bank of San Francisco Economic Review,* 1 19-30.
Rose, A. K. and Engel, C. (2002) Currency Unions and International Integration *Journal of Money, Credit, and Banking,* **34,** 4 1067-89.
Sachs, J. D. (1981) The Current Account and Macroeconomic Adjustment in the 1970s *Brookings Papers on Economic Activity,* **1,** 201-68.
Said, S. E. and Dickey, D. A. (1984) Testing for Unit Roots in Autoregressive-Moving Average Models of Unknown Order *Biometrika,* **71,** 3 599-607.
Sarantis, N. (1994) The Monetary Exchange Rate Model in the Long Run: An Empirical Investigation *Weltwirtschaftliches Archiv,* **130,** 4 698-711.
Sarmas, P. (1996) The Observed vs Implied Dollar Movements: A 1973-1990 Case Study *Finance, Real Estate and Law,* 123-131.
Sarno, L. (2001) Toward a New Paradigm in Open Economy Modeling: Where Do We Stand? *Federal Reserve Bank of St. Louis Review,* **83,** 3 21-36.
Sarno, L. and Taylor, M. P. (2002) *The economics of exchange rates.* Cambridge University Press: New York.
Sauer, C. and Bohara, A. K. (2001) Exchange Rate Volatility and Exports: Regional Differences between Developing and Industrialized Countries *Review of International Economics,* **9,** 1 133-52.
Schwert, G. W. (1987) Effects of Model Specification on Tests for Unit Roots in Macroeconomic Data *Journal of Monetary Economics,* **20,** 1 73-103.
Schwert, G. W. (1989) Tests for Unit Roots: A Monte Carlo Investigation *Journal of Business and Economic Statistics,* **7,** 2 147-59.
Serletis, A. and Zimonopoulos, G. (1997) Breaking Trend Functions in Real Exchange Rates: Evidence from Seventeen OECD Countries *Journal of Macroeconomics,* **19,** 4 781-802.
Singh, T. (2002) On the GARCH Estimates of Exchange Rate Volatility in India *Applied Economics Letters,* **9,** 6 391-95.
Singh, T. (2004) Testing J-Curve Hypothesis and Analysing the Effect of Exchange Rate Volatility on the Balance of Trade in India *Empirical Economics,* **29,** 2 227-45.
Smith, P. N. and Wickens, M. R. (1986) An Empirical Investigation into the Causes of Failure of the Monetary Model of the Exchange Rate *Journal of Applied Econometrics,* **1,** 2 143-62.
Sohmen, E. (1961) *Flexible exchange rates : theory and controversy.* University of Chicago Press: Chicago.
Stevenson, A., Muscatelli, V. and Gregory, M. (1988) *Macroeconomic theory and stabilization policy.* Oxford: Allan; Totowa, N.J.: Barnes and Noble.

Stock, J. H. (1999) 'A Class of Tests for Integration and Cointegration', in Engle, R. F. and White, H. e.(Eds) *Cointegration, causality, and forecasting: A festschrift in honour of Clive W. J. Granger.*
Stock, J. H. and Watson, M. W. (1987) 'Interpreting Evidence on Money-Income Causality', in National Bureau of Economic Research, Inc, NBER Working Papers: 2228.
Stulz, R. M. (1987) An Equilibrium Model of Exchange Rate Determination and Asset Pricing with Nontraded Goods and Imperfect Information *Journal of Political Economy,* **95,** 5 1024-40.
Svensson, L. E. O. and van Wijnbergen, S. (1989) Excess Capacity, Monopolistic Competition, and International Transmission of Monetary Disturbances *Economic Journal,* **99,** 397 785-805.
Tang, T. C. (2002) Demand for M3 and Expenditure Components in Malaysia: Assessment from Bounds Testing Approach *Applied Economics Letters,* **9,** 11 721-25.
Tang, T. C. (2002) Aggregate Import Demand Behavior for Indonesia: Evidence from the Bounds Testing Approach *IIUM Journal of Economics and Management,* **10,** 2 179-98.
Tang, T. C. (2002) An Aggregate Import Demand Function for Hong Kong, China:New Evidence from the Bounds Test *International Journal of Management,* **19,** 4 561-67.
Tang, T. C. (2003) Japanese Aggregate Import Demand Function: Reassessment from the 'Bounds' Testing Approach *Japan and the World Economy,* **15,** 4 419-36.
Tang, T. C. and Nair, M. (2002) A Cointegration Analysis of Malaysian Import Demand Function: Reassessment from the Bonds Test *Applied Economics Letters,* **9,** 5 293-96.
Tawadros, G. B. (2001) The Predictive Power of the Monetary Model of Exchange Rate Determination *Applied Financial Economics,* **11,** 3 279-86.
Taylor, M. P. (1995) The Economics of Exchange Rates *Journal of Economic Literature,* **33,** 1 13-47.
Taylor, M. P. and Sarno, L. (1998) The Behavior of Real Exchange Rates during the Post-Bretton Woods Period *Journal of International Economics,* **46,** 2 281-312.
Thursby, J. G. and Thursby, M. C. (1987) Bilateral Trade Flows, the Linder Hypothesis, and Exchange Risk *Review of Economics and Statistics,* **69,** 3 488-95.
Thursby, M. C. and Thursby, J. G. (1985) 'The Uncertainty Effects of Floating Exchange Rates: Empirical Evidence on International Trade Flows', in Arndt, S. W. e., Sweeney, R. J. e. and Willett, T. D. e.(Eds) *Exchange Rates, Trade, and the U.S. Economy.*
Toda, H. Y. and Phillips, P. C. B. (1991) 'The Spurious Effect of Unit Roots on Exogeneity Tests in Autoregressions: An Analytical Study', in Yale Cowles Foundation Discussion Paper: 978, pp. 32.
Toda, H. Y. and Phillips, P. C. B. (1993) The Spurious Effect of Unit Roots on Vector Autoregressions *Journal of Econometrics,* **59,** 3 229-55.

Towe, C. M. (1989) Exchange Rate "Fundamentals" versus Speculation in a Developing Economy: An Illustrative Example Using Lebanese Data *International Monetary Fund Staff Papers,* **36,** 3 678-707.

Uribe, M. (2003) Real Exchange Rate Targeting and Macroeconomic Instability *Journal of International Economics,* **59,** 1 137-59.

Visser, H. (1989) Exchange Rate Theories *De Economist,* **137,** 1 16-46.

Wang, K. and Barrett, C. (2002) *A new look at the trade volume effects of real exchange rate risk.* Department of Applied Economics and Management Cornell University

Westerfield, J. M. (1977) An Examination of Foreign Exchange Risk under Fixed and Floating Rate Regimes *Journal of International Economics,* **7,** 2 181-200.

White, H. (1980) A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity *Econometrica,* **48,** 4 817-38.

Wickens, M. R. (1996) Interpreting Cointegrating Vectors and Common Stochastic Trends *Journal of Econometrics,* **74,** 2 255-71.

Williams, G., Parikh, A. and Bailey, D. (1998) Are Exchange Rates Determined by Macroeconoimc Factors? *Applied Economics,* **30,** 4 553-67.

Wise, C. and Roett, R. e. (2000) *Exchange rate politics in Latin America.* Washington, D.C.: Brookings Institution Press.

Wolff, C. C. P. (1987) Time-Varying Parameters and the Out-of-Sample Forecasting Performance of Structural Exchange Rate Models *Journal of Business and Economic Statistics,* **5,** 1 87-97.

Wu, J.-L. (1996) The Empirical Investigation of Long-Run Purchasing Power Parity: The Case of Taiwan Exchange Rates *International Economic Journal,* **10,** 4 59-69.

Wu, Y. (1996) Are Real Exchange Rates Nonstationary? Evidence from a Panel-Data Test *Journal of Money, Credit, and Banking,* **28,** 1 54-63.

Yoon, G. (2003) The Time Series Behaviour of Brazilian Inflation Rate: New Evidence from Unit Root Tests with Good Size and Power *Applied Economics Letters,* **10,** 10 627-31.