ESSAYS IN EXCHANGE RATES AND INTERNATIONAL FINANCE

A thesis submitted for the Degree of Doctor of Philosophy

by

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ABSTRACT

This thesis is based on four essays in exchange rates and international finance. The first essay, examined in the second chapter, considers the long-run performance of the flexible-price monetary model as well as the real interest differential monetary model to explain the dollar–yen exchange rate during a period of high international capital mobility. We apply the Johansen methodology to quarterly data over the period 1980:01–2009:04 and show that the inadequacy of the two monetary models is due to the breakdown of their underlying building-blocks, money demand stability and purchasing power parity. In particular, modifying the monetary models by adjusting them for real stock prices to capture the stability of money demands on one hand and also for real economic variables such as productivity differential, relative government spending, and real oil price to explain the persistence in the real exchange rate on the other provide long-run relationships that appear consistent with the monetary models. Our findings of long-run weak exogeneity tests also emphasise the importance of the extended models employed here.

The second essay, examined in the third chapter, is on the nature of the linkages between stock market prices and exchange rates in six advanced economies, namely the US, the UK, Canada, Japan, the euro area, and Switzerland, using data on the banking crisis between 2007 and 2010. Bivariate GARCH-BEKK models are estimated to produce evidence of unidirectional Granger causality from stock returns to exchange rate changes in the US and the UK, in the opposite direction in Canada, and of bidirectional causality in the euro area and Switzerland. Furthermore, causality-in-variance from stock returns to exchange rate changes is found in Japan and in the opposite direction in the euro area and Switzerland, whilst there is evidence of bidirectional causality-in-variance in the US and Canada. These findings imply limited opportunities for investors to diversify their assets during this period.

The third essay, examined in the fourth chapter, considers the impact of net bond and net equity portfolio flows on exchange rate changes. Two-state Markov-switching models are estimated for the exchange rate of the US *vis-a-vis* Canada, the euro area, Japan and the UK. Our results suggest that the relationship between net portfolio flows

and exchange rate changes is nonlinear for all cases considered, except that of the US dollar against the Canadian dollar.

The fourth essay, examined in the fifth chapter, considers the impact of exchange rate uncertainty on different components of net portfolio flows, namely net equity and net bond flows, as well as the dynamic linkages between exchange rate volatility and the variability of these two types of flows. Specifically, a bivariate GARCH-BEKK-in mean model is estimated using bilateral data for the US *vis-à-vis* Australia, the UK, Japan, Canada, the euro area, and Sweden over the period 1988:01-2011:12. The results indicate that the effect of exchange rate uncertainty on net equity flows is negative in the euro area, the UK and Sweden, and positive in Australia, whilst two countries (Canada and Japan) showed insignificant responses. With regard to the impact of uncertainty on net bond flows, it is shown to be negative in all countries, except Canada (where it is positive). Under the assumption of risk aversion, this suggests that exchange rate uncertainty induces investors, especially those of the counterpart countries to the US, to reduce their financing activities to maximise returns and minimise exposure to uncertainty. This evidence is strong for the UK, the euro area and Sweden as opposed to Canada, Australia and Japan. Furthermore, since exchange rate volatility and the variability of flows are interlinked, exchange rate or credit controls on these flows can be used to pursue economic and financial stability.

Dedicated to my family

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DECLARATION

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PUBLICATIONS AND CONFERENCES

I have presented material from **Chapter 2** titled ‘‘*The Monetary Models of the US Dollar-Japanese Yen Exchange Rate: An Empirical Investigation’’* at *Econometrics and Macroeconomics Conference* (in a poster session), 2-3 May 2012, the University of Birmingham, UK, at the *11th Annual Conference of the European Economics and Finance Society (EEFS)*, 14-17 June 2012, Koç University, Istanbul, Turkey, and at an internal *Seminar at Economics and Finance Department*, 3 October 2012, School of Social Sciences, Brunel University, UK. A journal paper is drawn from the chapter and is published in *Economic Modelling* with the following details: Hunter, J. and Menla Ali, F. 2014. Money demand instability and real exchange rate persistence in the monetary model of USD–JPY exchange rate. *Economic Modelling*, 40, pp. 42-51.

I have presented **Chapter 3** titled ‘‘*On the Linkages between Stock Prices and Exchange Rates: Evidence from the Banking Crisis of 2007-2010*’’ at *the BMRC-QASS Conference on Macro and Financial Economics*, 24 May 2012, School of Social Sciences, Brunel University, UK, and at *the International Conference on the Global Financial Crisis*, April 25-26, 2013, the University of Southampton, UK. A paper drawn from the chapter is published in *International Review of Financial Analysis* with the following details: Caporale, G.M., Hunter, J. and Menla Ali, F. 2014. On the linkages between stock prices and exchange rates: Evidence from the banking crisis of 2007-2010. *International Review of Financial Analysis, forthcoming*.

Material drawn from **Chapter 4** titled ‘‘*Exchange Rates and Net Portfolio Flows: A Markov-Switching Approach*’’ has been published as a book chapter with the following details: Menla Ali, F., Spagnolo, F. and Spagnolo, N. 2014. Exchange rates and net portfolio flows: A Markov-switching approach. In: Mamon, R. and Elliott, R.

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I have presented **Chapter 5** titled ‘‘*Exchange Rate Uncertainty and International Portfolio Flows*’’ at *the 9th BMRC-QASS Conference on Macro and Financial Economics*, 30 May 2013, Brunel University, UK and at *the 3rd GPEN-CGR Annual Conference*, 24-25 June 2013, Queen Mary, University of London. A paper drawn from the chapter is to be revised and resubmitted for *Journal of International Money and Finance* with the following details: Caporale, G.M., Menla Ali, F. and Spagnolo, N. 2013. Exchange rate uncertainty and international portfolio flows. Economics and Finance Working Paper, No. 13-10. Brunel University.

Table *of* Contents

CHAPTER THREE

ON THE LINKAGES BETWEEN STOCK PRICES AND EXCHANGE RATES: EVIDENCE FROM THE BANKING CRISIS OF 2007–2010

CHAPTER FOUR

EXCHANGE RATES AND NET PORTFOLIO FLOWS: A MARKOV-SWITCHING APPROACH

CHAPTER FIVE

EXCHANGE RATE UNCERTAINTY AND INTERNATIONAL PORTFOLIO FLOWS

CHAPTER SIX

List *of* Tables

List *of* Figures

CHAPTER ONE

INRODUCTION

Modelling exchange rate movements and analysing its volatility have been the matter of much attention in the field of international macroeconomics and finance, ever since the collapse of the Bretton Woods system in 1971 and the inception of the floating exchange rates in March, 1973. The impact of the exchange rate volatility has been particularly examined on the macroeconomic variables, especially trade flows and economic growth. With regard to modelling the dynamics of exchange rates, a number of exchange rate specifications have been proposed producing subsequently a large body of literature in the field. The most celebrated specifications include the Uncovered Interest Parity (UIP), the flexible-price monetary model (e.g., Frenkel, 1976; Bilson, 1978), the sticky-price monetary model (Dornbusch, 1976), the real interest rate differential monetary model (Frankel, 1979), the portfolio balance models (e.g., Branson, 1976; Dooley and Isard, 1979) and the general equilibrium models of Stockman (1980) and Lucas (1982). Some previous specifications have also been thoroughly examined such as the Purchasing Power Parity (PPP), developed by Cassel.

However, the poor performance of these specifications in out-of-sample forecasting (see Meese and Rogoff, 1983; Cheung et al., 2005) and the increased capital mobility across-borders ever since the removal of capital controls and the deregulation of international financial markets in the late of 1970s and the early of 1980s led scholars to examine the underlying microstructure actions of the exchange rates. Indeed, the recent microstructure approach of exchange rate determination has been rather successful. It is found that the currency order flow (buyer initiated trades minus seller initiated trades in the foreign exchange market) explains a substantial proportion of exchange rate variations (e.g., Evans and Lyons, 2002; 2005; 2008; Rime et al., 2010; Payne, 2003; Chinn and Moore, 2011; and Duffuor et al., 2012; among others).

Furthermore, Hau and Rey (2006) developed a portfolio choice model of exchange rates on the basis that order flows and net capital flows, primarily net portfolio flows, are intimately aligned since the two flows represent the investors' behaviour. While Hau and Rey (2006) found that the portfolio rebalancing motive is strongly supported for 17 OECD countries, the subsequent studies showed that their hypothesis is weak in commodity-exporting countries (see Chaban, 2009; Ferreira Filipe, 2012; among others).

This thesis contains *four essays* in the field of exchange rates and international finance. Specifically, the first three essays are centred on the exchange rate determination issue based on two approaches: the monetary approach and the portfolio choice approach. The fourth essay, on the other hand, is based on the impact of exchange rate uncertainty. That is, the impact of such an uncertainty on the financing activities associated with equity and bond portfolio flows across-borders is analysed. To accomplish the essays, a wide range of time series econometric models is employed such as cointegration tests, multivariate GARCH model, multivariate GARCH-in mean model, and the Markov-switching specifications.

Chapter 2 examines two versions of the standard monetary model of exchange rates: the flexible-price monetary model of Bilson (1978) and the real interest rate differential model of Frankel (1979). The chosen models are among the most thoroughly investigated models in the empirical literature. Despite the models have important theoretical appeal, they have limited success on the empirical side. The empirical failure has been specifically made with regard to the US dollar-Japanese yen exchange rate in which there is no academic consensus on the factors that drive the dynamics. For example, Caporale and Pittis (2001) find that the monetary model of the yen-dollar exchange rate is unstable, whilst the recent study of Chinn and Moore (2011) find that there is limited evidence of a long-run relation between the nominal dollar-yen exchange rate and its monetary fundamentals, even when the cumulative order flow is included in the model. In a recent study, Maurice Obstfeld (2009, p. 1) discusses that 'the determinants of the yen's short- and even longer-term movements remain mysterious in light of the development of Japan's macro economy'.

The aim of the chapter is to investigate the empirical failure of the two models (the flexible-price and real interest rate differential models), applied to the US dollar-Japanese yen exchange rate. We employ the Johansen cointegration technique, using quarterly data over a period characterised by high international capital mobility between the US and Japan, 1980:01-2009:04. In particular, we authenticate that the limited success of the two monetary models considered here is due to the breakdown of their underlying building blocks: money demand stability and PPP. Modifying the models by adjusting for factors affecting the stability of money demands and the persistence of the real exchange rate provides supportive results. The modified models are devised on the basis that domestic and foreign money demand equations include broader asset classes such as real stock prices, whereas the persistence of the real exchange rate is accounted for by the inclusion of real economic shocks (productivity differential, relative government spending, and the real oil price). The findings of long-run weak exogeneity tests also emphasise the importance of the modified models employed here. Considering the real interest rate differential model and its modified version, it is shown that the seemingly cumulated shocks to the nominal exchange rate are originated from the shocks of factors affecting the conventional monetary model's building blocks, especially relative real equity prices and productivity differential. Finally, the results demonstrate that the modified models outperform the random walk benchmark in out-of-sample forecasting in the medium- and long-term, but not the short-term.

Chapter 3 examines the dynamic linkages in terms of the first and second moments between stock market prices and exchange rates during the recent financial crisis. The collapse of the Lehman Brothers on September 15^{th} 2008, which occurred after the crisis with mortgage-backed securities and the failure of Fannie Mae and Freddie Mac in the US that had started to emerge in late 2007, sent a wave of panic across international financial markets. As a consequence, not only international stock markets exhibited severe downturns across developed economies, but also major foreign exchange rates were hit by significant changes during the period.

While the extant studies have investigated the depth of such crisis in terms of its causes and consequences (e.g., Poole, 2010; Yeager, 2011), correlations, volatility spillovers, and contagion effects across international stock markets (e.g., Aloui et al., 2011; Kenourgios et al., 2011; Samarakoon, 2011; Dufrénot et al., 2011; among others) and across foreign exchange markets (e.g., Coudert et al., 2011; Bubák et al., 2011), thus far the relationship between the two financial markets (stock and foreign exchange markets) during the recent financial crisis has drawn less attention.

The only exceptions, to the best of knowledge, are the studies of Wong and Li (2010), Tsai (2012), Tsagkanos and Siriopoulos (2013), and Chkili and Nguyen (2014), though they have some limitations in terms of the data frequency, the sample period, and the adopted econometric techniques. Specifically, Wong and Li (2010) and Tsai (2012) use monthly data which cannot capture the timing of events and the evolution of capital across the financial markets. Also, their data are collated till the end of 2008 and 2009, respectively, thereby not capturing the turbulent periods ensued the collapse of the Lehman Brothers and the European sovereign debt crisis. Tsagkanos and Siriopoulos (2013) and Chkili and Nguyen (2014), on the other hand, use higher frequency data and longer sample periods to cover the recent financial crisis. However, their short-run dynamics results are characterised by significant deviations from normality and conditional heteroscedasticity (see Engle, 1982) that are not captured by their setup.

We use weekly data from six advanced economies, namely the US, the UK, Canada, Japan, the euro area and Switzerland and consider two sub-periods: the pre-crisis period (August 6, 2003-August 8, 2007) and the crisis period (August 15, 2007-December 28, 2011). Furthermore, we conduct the analysis by using a bivariate VAR-GARCH (1, 1) in the BEKK specification of Engle and Kroner (1995). The adopted framework allows for the time-varying conditional correlation and also for interactions in the variances in a lead-lag manner. To avoid potential missing variable bias in the conditional mean, the model is also augmented by the underlying short-run deviations between stock market prices and exchange rates in the conditional mean in the event that both variables share a common stochastic trend; the Engle and Granger (1987), the Johansen (1995), and the Gregory and Hansen (1996) cointegration tests are employed.

The empirical findings proved the existence of unidirectional Granger causality from stock returns to exchange rate changes in the US and the UK, in the opposite direction in Canada, and of bidirectional causality in the euro area and Switzerland during the recent financial crisis. Furthermore, causality-in-variance from stock returns to exchange rate changes is found in Japan and in the opposite direction in the euro area and Switzerland, whilst there is evidence of bidirectional causality-in-variance in the US and Canada during the period. These findings are broadly consistent with those of Tsagkanos and Siriopoulos (2013) in examining the linkages between stock prices and exchange rates during the recent financial crisis and also with those of Granger et al. (2000) and Caporale et al. (2002), who investigated the linkages between the two variables during the 1997 Asian financial crisis.

The results indicate that the heterogeneous strength of the considered economies' currencies against each other throughout the financial crisis may have played a major role in generating capital inflows and outflows, thereby resulting in different results when examining the dynamic linkages between stock returns and exchange rate changes within these economies. Furthermore, as stock and foreign exchange markets are shown to be linked during the crisis period, this implies the existence of limited opportunities for investors for portfolio diversifitcaion during the period.

Chapter 4 examines the extent to what equity and bond portfolio flows across-borders drive the dynamics of exchange rates. With the deregulation of financial markets and the increase in the flow of capital across-borders, it is widely believed that the excess volatility of major currencies is likely to be explained by financial market responses which are, in turn, driven by the resulting increase in such flows of capital. Indeed, while the US dollar-based exchange rates were relatively less volatile in the 1970s, these exchange rates have been characterised by high volatility ever since the removal of capital controls and the deregulation of international financial markets in the early of the 1980s, as stated earlier. With regard to the evolution of capital across-borders, notwithstanding gross cross-border portfolio investments in equities and bonds in the US were accounting for only 4% of GDP in 1975, this proportion surged to 100% in the early of 1990s and has evolved towards 245% by 2000 (Hau and Rey, 2006).

The chapter specifically examines the nonlinear dependence between exchange rate changes and net equity and net bond portfolio flows for the US against the UK, the euro area, Japan and Canada, using quarterly data over the period 1990:01-2011:04. To the best of knowledge, the *regime-switching* in the relationship between net portfolio investment flows and exchange rate changes is yet to be explored in the literature. Most existing empirical studies only document short-run dynamic interactions using linear dependence techniques such as OLS and VAR models (see, for examples, Brooks et al., 2004; Hau and Rey, 2006; Chaban, 2009; Kodongo and Ojah, 2012; among others).

However, such studies have assumed linear dependence and constant parameters, which is not the case since both exchange rates and portfolio flows underwent structural changes in the light of the financial crises observed over last few decades. Therefore the Markov regime-switching specification is particularly appropriate to examine the impact of net portfolio flows on exchange rate changes in two states of such changes by allowing the data themselves to identify these states. That is, when the exchange rate appreciates and depreciates *and* when it exhibits high volatility and low volatility, thereby relaxing the linear constraint associated with earlier studies.

The empirical results show that the relationship between exchange rate changes and net portfolio flows is state-dependent for all cases, except Canada. Specifically, it is shown that net equity and net bond inflows from the UK towards the US result in an appreciation of the US dollar against the British pound in the appreciation regime. Furthermore, net bond inflows from Japan towards the US imply an appreciation of the US dollar against the Japanese yen in the less volatile regime. The results of the euro area, by contrast, suggest that net bond inflows from the euro area towards the US result in a US dollar appreciation (depreciation) against the euro in the low (high) volatility regime. The insignificant effects of net equity and net bond portfolio flows in the case of Canada, though, are in line with previous studies on commodity-exporting countries (see Chaban, 2009; Ferreira Filipe, 2012).

Chapter 5 examines the impact of exchange rate uncertainty on net portfolio flows across borders. The underlying intuition is that exchange rate volatility affects adversely portfolio flows across borders by increasing transaction costs and reducing potential gains from international diversification, hence making the acquisition of foreign securities such as bonds and equities more risky (Eun and Resnick, 1988). The chapter examines the impact of the uncertainty on different components of net portfolio flows, namely net equity and net bond flows, as well as the dynamic linkages between exchange rate volatility and the variability of these two types of flows. Specifically, a bivariate VAR GARCH-BEKK-in mean model is estimated using monthly bilateral data for the US *vis-à-vis* Australia, the UK, Japan, Canada, the euro area, and Sweden over the period 1988:01-2011:12.

The results indicate that the effect of exchange rate uncertainty on net equity flows is negative in the euro area, the UK and Sweden, and positive in Australia, whilst it is negative in all countries except Canada (where it is positive) in the case of net bond flows. These findings suggest that exchange rate uncertainty induces investors, especially those of the counterpart countries to the US, to reduce their international financing activities to maximise returns and minimise exposure to uncertainty. This evidence is strong in the cases of the UK, the euro area and Sweden as opposed to Canada, Australia and Japan. The findings of the latter countries may be due to these countries' specific characteristics which have been documented by previous studies in the literature (e.g., Hau and Rey, 2006; Chaban, 2009; and Ferreira Filipe, 2012). Furthermore, since exchange rate volatility and the variability of flows are interlinked, exchange rate or credit controls on these flows can be used to pursue economic and financial stability.

Chapter 6 offers conclusions and suggestions as to how to develop further the research in this thesis in ways beyond the current scope of the work.

CHAPTER TWO

THE MONETARY MODELS OF THE US DOLLAR-JAPANESE YEN EXCHANGE RATE: AN EMPIRICAL INVESTIGATION

2.1. Introduction

Since the collapse of the Bretton Woods fixed exchange rate system in 1971, much attention has been paid towards finding a meaningful explanation of exchange rates. At a later date, a wide range of models has been proposed to understand movements in the exchange rate. Among the variants which have been scrutinised thoroughly is the monetary approach as it has an important policy relevance. At first sight such an approach has an intuitive hypothetical appeal by linking the nominal exchange rate to its monetary fundamentals, stimulated by the notion that the exchange rate is the relative price of two currencies and that the national price levels are determined by their supply and demand in their corresponding national money markets. Despite having rigorous theoretical underpinnings, the model has empirically had limited success until now.

Strictly speaking, while the empirical examination in the 1970s provided favourable results for such a model (e.g., Frenkel, 1976; Hodrick, 1978; Bilson, 1978; and Frankel, 1979), the model had been characterised by limited success when it had been subjected to data of the 1980s. For instance, Haynes and Stone (1981) showed the collapse of Frankel (1979)'s model once his data were extended to 1980:04, with relative money supply negatively signed, consistent with the evidence of Dornbusch

(1980). Backus (1984) also provided unsupportive evidence of the monetary approach, using data on the Canadian dollar-US dollar exchange rate.

Furthermore, the empirical studies over the later stages of the development of this literature have revealed how difficult it is to detect a cointegrating relationship between the nominal exchange rate and its monetary fundamentals using the Engle and Granger (1987) two-step procedure (e.g., Baillie and Selover, 1987; Boothe and Glassman, 1987; and McNown and Wallace, 1989). More importantly, Meese and Rogoff (1983) showed in a seminal work that such structural exchange rate models, including those based on monetary fundamentals, are incapable of outperforming the naïve random walk in out-of-sample forecasting.

Employing the Johansen (1988; 1995) cointegration technique by MacDonald and Taylor (1991; 1994), on the other hand, enlivened such a model by not only validating it as a long-run equilibrium foundation, but also providing improvement in its predictive power over the random walk model in out-of-sample forecasting after a decade of the gloomy outlook. Studies of MacDonald and Taylor (1991; 1994) have stimulated a strand of empirical literature that utilises the Johansen cointegration technique in examining the monetary model in a multivariate framework and providing evidence of a long-run relationship among its variables (e.g., McNown and Wallace, 1994; Moosa, 1994; Choudhry and Lawler, 1997; Diamandis et al., 1998; Kouretas, 1997; Cushman, 2000; Tawadros, 2001; Francis et al., 2001; among many others). Inspite of this revival of its empirical appeal; however, the broad conclusions emerged from such studies, including those of MacDonald and Taylor (1991; 1994), stress that the signs and magnitudes of the estimated coefficients lend limited support whatsoever to the monetary mainstream theory, thereby inducing its controversy as a valid framework for exchange rate determination once more.

Husted and MacDonald (1998), Groen (2000), Mark and Sul (2001), and Rapach and Wohar (2004) among others found some evidence of the monetary model in a panel context, but this was under the assumption of a high order of heterogeneity across all country models. Similarly, Rapach and Wohar (2002) found some support for the theory using long time series, but this was related to different exchange rates and macro regimes. Taylor and Peel (2000) and Kilian and Taylor (2003) applied nonlinear methods to model a nominal exchange rate and monetary fundamentals, but such results are often sensitive to a small number of observations and become less robust as the sample evolves. Frömmel et al. (2005) estimated the real interest diffential (RID) model of Frankel (1979) with the Markov switching approach. However, the monetary model was shown to be related to only one regime.[1]

Furthermore, the empirical failure of this model has been specifically found in regard of the US dollar-Japanese yen exchange rate (see, for examples, Lizardo and Mollick, 2010; Chinn and Moore, 2011). The evolution of this exchange rate has been much debated over the recent years with no consensus on the factors that drive the dynamics. For instance, while Caporale and Pittis (2001) found an unstable relation based on the yen-dollar exchange rate monetary model, the recent study by Chinn and Moore (2011) failed to uncover evidence of cointegration considering the conventional flexible-price monetary model and using monthly data over eight years on the dollar-yen besides of the dollar-euro exchange rates. Although extending such a model by cumulative order flow provided strong evidence of cointegration for the dollar-euro exchange rate, this was not the case for the dollar-yen exchange rate. By contrast, MacDonald and Nagayasu (1998) only found that a simplified version of the real interest differential model of Frankel (1979), which excludes money demand functions,

[1] For a comprehensive overview of the literature, see MacDonald (2007, Ch6) or Moosa and Bhatti (2010, Ch12).

holds for the yen-dollar exchange rate over the period 1975:Q3-1994:Q3. In addition, Rogoff (2001, p. 6) puts it ‘explaining the yen, dollar and euro exchange rates is still a very difficult task, even ex-post’. Obstfeld (2009, p. 1) also disputes that ‘the determinants of the yen’s short- and even longer-term movements remain mysterious in light of the development of Japan’s macro economy’. Instead, Hamada and Okada (2009), by analysing the evolution of the yen real exchange rate, argue that monetary and global factors were as important as non-monetary and domestic factors in causing the stagnation of the Japanese economy. In a recent paper, Ruelke et al. (2010), by using the Wall Street Journal poll, find that forecasters can be regarded heterogeneous in the expectation formation process for the yen against the US dollar over the period 1989–2007.

A more authentic explanation for the limited success of the monetary model is perhaps due to the breakdown of its underlying building blocks: stable money demands and purchasing power parity (PPP). In other words, since the stability of money demands and PPP are the primary influencing data dynamics in producing a sensible long-run monetary model of the exchange rate, it is rather unlikely to find such an intuitive model in terms of both providing a long-run equilibrium relationship between the nominal exchange rate and its monetary fundamentals and producing signs predicted by the monetary mainstream theory when domestic and foreign money demand equations are unstable or/and the real exchange rate is persistent.

Indeed, Hendry and Ericsson (1991) found that the conventional money demand equation for the US was not stable.[2] In a related vein, Friedman (1988) and McCornac

[2] Barnett (1980) and Barnett et al. (1984) also showed that divisia monetary aggregates measure captures the traditional transaction motive for holding money and tends to be more closely related to the general price level in the economy than the simple sum money. In the context of the monetary model, Chrystal and MacDonald (1995) and Chin et al. (2009) used divisia money rather than simple sum money to take into account the instability of the money demand for the UK and five Asian countries, respectively.

(1991), using data from the United States and Japan, respectively, confirmed the need for real stock prices to stabilise money demand equations. In the context of the monetary model, both Morley (2007) and Baharumshah et al. (2002) provided successful results respectively for the UK and Malaysia when the monetary model has been augmented by real stock prices.[3] Another motivation for including real stock prices into the monetary model via money demand equations, *per se*, is that financial press and financial market analysts advocate that there exists a relationship between stock prices and exchange rates.[4]

Rogoff (1996) and Sarno and Taylor (2002), on the other hand, found little support for the conventional PPP by surveying a range of empirical studies. This corresponds well with the classic findings of Balassa (1964) and Samuelson (1964), which indicate that the persistent deviations from PPP arise from productivity differentials. The real economic shocks that have been found to explain the persistent deviations from the PPP also include government spending and the real oil price; the latter mainly to capture the terms of trade shocks (for a recent survey of the empirical literature, see Tica and Družić, 2006). Lastrapes (1992), Enders and Lee (1997), Chen and Wu (1997), Chinn (1997; 2000), Wang and Dunne (2003), and Tsen (2011) altogether showed that fluctuations in the nominal and real exchange rates are due to the impact of differentials in productivity and government expenditure along with the real oil price.

This chapter contributes to the existing literature by proposing a modified monetary model of the dollar-yen exchange rate that takes into account the breakdown

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[3] To the best of our knowledge, examining the impact of real stock prices on the dollar-yen exchange rate in the context of the monetary model has not been examined in the literature yet.

[4] For instance, titled '*What's next for stocks, M&A, and the Dollar?*' Business Week magazine linked an increase of 2.5% in the S&P 500 stock index on September 21, 2009 and its recording of total rise of 58% since its growing in March 2009 with the dwindling in the dollar against other major currencies (see Bloomberg Businessweek, 2009).

of the aforementioned building blocks. That is, the proposed model captures both the monetary and the real aspects of the economy, thereby circumventing some of the potential pitfalls associated with earlier studies. More specifically, we examine the empirical performance of the standard flexible-price monetary model of Bilson (1978), as well as the standard RID model of Frankel (1979) against their corresponding modified versions, proposed herein, by employing the Johansen (1995) methodology and quarterly data from 1980:01 to 2009:04, a period characterised by high international capital mobility and volatility.

While the RID model is a realistic description when variation in the inflation differential is moderate as is the case between the US and Japan over the period under examination,[5] we also use the flexible-price model for robustness and comparison purposes of this study. The proposed versions of the two conventional monetary models, by contrast, are devised by using domestic and foreign money demand equations based on broader asset classes and also accounting for the factors that cause PPP to fail. That is, we incorporate real stock prices in the money demand equations, while we use the productivity differential, relative government spending, and real oil price to explain the persistence in the real dollar-yen exchange rate.

The remainder of this chapter is organised as follows. Section 2.2 reviews the theoretical monetary models of exchange rates and then alternative versions of the monetary models are introduced. Section 2.3 describes the data and outlines the research econometric technique which has been conducted in the study. Section 2.4 introduces the empirical results and the analysis for both the standard and modified monetary models of exchange rates, and finally Section 2.5 provides conclusions.

[5] Bernanke (2000) and Taylor (2001) argued that the different inflationary environments in the US and Japan are due to the differences in the monetary policies in the two countries.

2.2. The monetary models of exchange rate determination

2.2.1. The standard monetary models of exchange rates

The monetary models of exchange rates are primarily based on two building blocks: stable money demand functions and the PPP (for an extensive discussion of the monetary models, see Pilbeam, 2006; MacDonald, 2007). In this chapter, we scrutinise four forms of the monetary models of exchange rates: the standard flexible-price monetary model (FPM) (Frenkel, 1976; Bilson, 1978), the standard real interest rate differential monetary model (RID) (Frankel, 1979), the modified flexible-price monetary model (MFPM), and the modified real interest rate differential monetary model (MRID). The modified monetary models are in fact the standard ones; however, they adjust for factors that affect the underlying building blocks of the standard models.

The flexible-price monetary model starts with money demand functions in the domestic and foreign country in a Cagan-style as follows:

$$m_t - p_t = a_1 y_t - a_2 i_t \,, \tag{2.1}$$

$$m_t^* - p_t^* = a_1 y_t^* - a_2 i_t^* \,, \tag{2.2}$$

where m_t is the money supply, p_t is the price level (CPI), i_t is the nominal interest rate, and y_t is real income. Apart from the nominal interest rates, all variables are expressed in log terms. Asterisks denote the foreign country variables. Note that, for simplicity, the income elasticity of money demand a_1 and interest rate semi-elasticity of money demand a_2 are assumed to be identical across both domestic and foreign countries. Finally, in our empirical study, the US economy is considered to be the domestic or home country.

Rearranging Eqs. (2.1) and (2.2) for the domestic and foreign price levels, first, and then for relative prices of the domestic and foreign country, we obtain:

$$p_t - p_t^* = (m_t - m_t^*) - a_1(y_t - y_t^*) + a_2(i_t - i_t^*). \qquad (2.3)$$

It is also assumed that PPP, in its absolute form, holds continuously:

$$e_t = p_t - p_t^*, \qquad (2.4)$$

where e_t is the spot exchange rate (domestic currency per unit of foreign currency, \$/yen in our case). Substituting Eq. (2.4) into Eq. (2.3) for the relative prices, the flexible-price monetary model is obtained and can be estimated as follows:

$$e_t = \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*) + \beta_3(i_t - i_t^*) + \varepsilon_t, \qquad (2.5)$$

where ε_t is an error term assumed to be a white noise process. Eq. (2.5) represents the flexible-price monetary model of Frenkel (1976), Mussa (1976), and Bilson (1978), who state that the nominal exchange rate is driven by relative money supply, relative real income and interest rate differential.[6] Specifically, an increase in the domestic money supply relative to the foreign counterpart results in a one for one depreciation of the nominal exchange rate, hence it implies that $\beta_1 = 1$. An increase in the domestic real income relative to that of the foreign economy, on the other hand, boosts the domestic real money demand which results in a reduction in the domestic price levels, and hence an appreciation of the domestic currency is induced for the PPP to be

[6] Considering the variables in relative forms implies that the coefficients across the domestic and foreign country are restricted to be the same in terms of size, but of opposite sign. Such an assumption is made for econometric convenience and is based on the assumptions that money demand functions are identical across the domestic and foreign country (see Boothe and Glassman, 1987).

maintained ($\beta_2 < 0$). A rise in the domestic interest rate relative to its foreign counterpart results in a reduction in the demand for real money balances which leads to higher prices and nominal exchange rate depreciation $(\beta_3 > 0)$, in turn, via the PPP.

The nominal interest rate in Eq. (2.5) can be decomposed into both the real interest rate r_t and the expected inflation rate Δp^e:

$$i_t = r_t + \Delta p^e, \tag{2.6}$$

$$i_t^* = r_t^* + \Delta p^{*e} . \tag{2.7}$$

Assuming that the real interest rates are identical across the domestic and foreign country, it yields:

$$i_t - i_t^* = \Delta p^e - \Delta p^{*e}. \tag{2.8}$$

Substituting Eq. (2.8) into Eq. (2.5), the model can be written as follows:

$$e_t = \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*) + \beta_3(\Delta p^e - \Delta p^{*e}) + \varepsilon_t. \tag{2.9}$$

Frankel (1979), on the other hand, developed a monetary model which combines aspects of both the flexible-price (New Classical) and sticky-price (Neo Keynesian) monetary models of exchange rates by incorporating short-term interest rates to capture the stance of monetary policy. More specifically, the model asserts that the expected rate of depreciation of the exchange rate is a function of not only the gap between the current rate e and the long-run equilibrium rate $\bar{e}$, as in Dornbusch (1976)'s sticky-price model, but also the expected long-run inflation differential between the domestic and foreign country. It is as follows:

$$E(\Delta e) = \delta(\bar{e} - e) + (\Delta p^e - \Delta p^{*e}), \tag{2.10}$$

where δ is the speed of adjustment towards the equilibrium level, Δp^e and Δp^{*e} denote the domestic and foreign expected long-run inflation rates, respectively. The latter equation, Eq. (2.10), highlights the short-run and long-run dynamics of exchange rate changes. In the short-run, the spot exchange rate e is expected to return to its long-run equilibrium value $\bar{e}$ at a rate equal to δ. In the long-run (since $\bar{e} = e$), changes in the exchange rate will be proportional to the expected long-run inflation differential $(\Delta p^e - \Delta p^{*e})$.

Assuming the uncovered interest parity (UIP) condition, $E(\Delta e)= i_t - i_t^*$, that postulates domestic and foreign bonds are perfect substitutes, then combining such a condition with Eq. (2.10) and rearranging for the spot exchange rate, we obtain:

$$e = \bar{e} - \frac{1}{\delta}[(i - \Delta p^e) - (i^* - \Delta p^{*e})]. \tag{2.11}$$

It is usually implied that $\bar{e}$ in Eq. (2.11) is determined by the flexible-price monetary model derived in Eq. (2.9) (see MacDonald, 2007, Ch6), which is the reduced form of $e_t = (m_t - m_t^*) - a_1(y_t - y_t^*) + a_2(\Delta p^e - \Delta p^{*e})$. By combining the corresponding expressions in the latter reduced form and Eq. (2.11), it yields:

$$e = (m - m^*) - a_1(y - y^*) + a_2(\Delta p^e - \Delta p^{*e}) - \frac{1}{\delta}[(i - \Delta p^e) - (i^* - \Delta p^{*e})]. \tag{2.12}$$

It is common practice to estimate this equation empirically on the basis that short term interest rates represent real interest rates (i.e., liquidity effects of monetary policy) and long term interest rates capture the long-run expected inflation rates (see Frankel, 1979; Macdonald, 2007, Ch6). Thus, the baseline model is in the reduced form written as follows:

$$e_t = \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*) + \beta_3(i_t^s - i_t^{s*}) + \beta_4(i_t^l - i_t^{l*}) + \varepsilon_t, \tag{2.13}$$

where i_t^s denotes the short-term interest rate and i_t^l is the long-term interest rate used to capture the expected inflation. The asterisk, as stated earlier, denotes the foreign country (Japan), and the domestic country is the United States. Otherwise, the RID model related to (2.13) hypothesises that an increase in the domestic money supply relative to the counterpart foreign one increases domestic prices and thus causes a one for one depreciation in the exchange rate $(\beta_1 = 1)$. An increase in domestic income or a decline in the expected rate of domestic inflation (proxied by the long-term interest rate) relative to the foreign one raises the demand for money and thus causes an appreciation in the exchange rate $(\beta_2 < 0, \beta_4 > 0)$. An increase in the domestic nominal interest rate relative to the foreign one induces capital inflows towards the domestic economy and thus causes an appreciation in the exchange rate $(\beta_3 < 0)$. For further details the reader is directed to Frankel (1979).

2.2.2. Alternative versions of the monetary models of exchange rates

Friedman (1988) and subsequently McCornac (1991) showed that the stability of money demand functions, Eq. (2.1) and Eq. (2.2), which are utilised to formulate the monetary models, appeal for including real stock prices. The conclusion of Friedman's seminal work has also been confirmed by Choudhry (1996) for the US and Canada, Thornton (1998) for Germany, Caruso (2001) for Japan, the UK, Switzerland and Italy as well as for a panel of 25 (19 industrial and 6 developing) countries, and Baharumshah et al. (2009) for China, among others. Friedman's theoretical interpretation of the relationship between money demand and stock prices has primarily taken two kinds of effects: 'wealth effect' which posits a positive correlation between stock prices and the money demand and 'substitution effect' which suggests a negative correlation.[7]

Another motivation for encompassing stock prices themselves into the monetary model of exchange rates is that there exists a relationship between stock prices and exchange rates. Empirical examples of the dynamic relationship between stock prices and exchange rates are given by Aggarwal (1981), Granger et al. (2000), Kanas (2000), Nieh and Lee (2001), Caporale et al. (2002), and Phylaktis and Ravazzola (2005), among others. See also **Chapter 3,** which examines the dynamic linkages between stock prices and exchange rates during the recent financial crisis. Figure 2.1 displays quarterly relative stock prices, deflated by the corresponding CPIs, and the movements of the dollar-yen exchange rate over the period 1980:01-2009:04. The graphical analysis signals that international stock price differential movements in real terms are likely to provide information for detecting trends in the US dollar against the Japanese yen.

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[7] The wealth effect, according to Friedman, is due to the following three different factors: first, an increase in the stock prices results in an increase in the nominal wealth as well as the wealth to income ratio, thereby causing an increase in money to income ratio. Second, an increase in the stock prices implies an increase in the expected returns on risky assets in comparison to safe assets. This increase in the risk of a portfolio could be offset and eliminated by diversifying the portfolio through decreasing risky assets such as long-term bonds and increasing safer assets such as short-term assets and money in such a portfolio. Third, an increase in the stock prices implies an increase in the volume of financial transactions, thereby increasing the money demanded to undertake such transactions. The negative substitution effect, on the other hand, implies that as real stock prices rise, equities become more attractive for investors, thereby inducing a substitution from other assets, including money, to stocks in a portfolio.

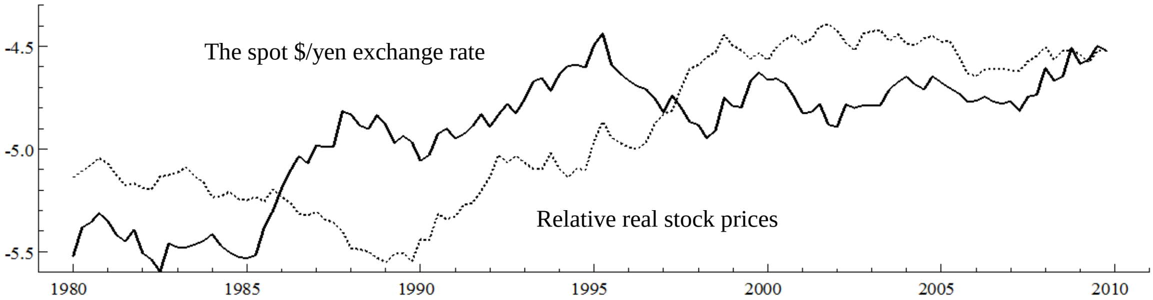

Figure 2.1. The movements of the spot $/yen exchange rate and relative real stock prices between the US and Japan from 1980:Q1 to 2009:Q4.

There appear clear signs of negative correlation between relative stock prices in real terms and the US dollar-Japanese yen exchange rate albeit such correlation is segmented. The negative relationship indicates the pre-dominance of the wealth effect, in the sense of Friedman (1988), of real stock prices in these countries.

Thus, based on the aforementioned motivation, money demand equations, given by Eqs. (2.1) and (2.2), in the previous subsection are modified as follows:

$$m_t - p_t = a_1 y_t - a_2 i_t + a_3 s_t, \tag{2.1$'$}$$

$$m_t^* - p_t^* = a_1 y_t^* - a_2 i_t^* + a_3 s_t^*, \tag{2.2$'$}$$

where s_t and s_t^* denote the log of domestic and foreign real stock price indices, respectively. The parameter on real stock price, a_3, is an empirical matter and depends on whether a substitution effect (negative) or wealth effect (positive) dominates the money demand equation.

Having specified the stability of money demand equations - the first building block of the monetary approach- we also adjust the monetary models for factors that are influencing the validity of PPP- the second building block of the model. The empirical studies on the PPP in the post-Bretton Woods period have analysed its time series properties by either investigating the long-run PPP by applying cointegration techniques between the nominal exchange rate and the domestic and foreign price levels (or national price ratio) or examining the stationarity of real exchange rate series by employing unit root tests, variance ratio tests, and fractional integration techniques.

In a broad sense, the applied single cointegration tests based on the Engle and Granger (1987) two-step procedure revealed that the link between the nominal exchange rate and national price levels is flimsy. Even though the multivariate Johansen (1988; 1995) cointegration technique provided evidence of cointegration for the PPP (mainly the relative version), the proportionality and symmetry restrictions among the nominal exchange rate and national price levels produced mixed results (e.g., Kugler and Lenz, 1993; Cheung and Lai, 1993). However, using time-varying coefficient approach, the most recent paper of Hall et al. (2013) provided evidence of the existence of strong support for homogeneity condition, defined as the proportionality in the presence of omitted variables, in the PPP specification.[8] In a related vein, Coakley et al. (2005b) find that it is the price index whether it is consumer price index (CPI) or producer price index (PPI) that matters for the support for the PPP.

The unit root tests, on the other hand, showed that the real exchange rates are non-stationary and that they follow a random walk process instead (see Sarno and Taylor, 2002; and Taylor, 2006). Though, using long span of data (e.g., Edison, 1987; Lothian and Taylor, 1996; and Taylor, 2002; among many others)[9] or panel data

[8] The authors use two experiments. The first experiment involves data on nine euro area countries and the euro area as an aggregate economy over the period 1999: M1-2011: M3, while the second experiment extends the same group of countries by Canada, Japan and Mexico and uses a longer period 1957: M4-2011: M3.

[9] Besides of the difficulty in obtaining long span of data, one more caveat with using long span of data, as stated earlier in this chapter, is that it involves different exchange rate regimes, and hence the interpretation of the results is not straightforward.

techniques (e.g., Lothian, 1997; Frankel and Rose, 1996; Wu, 1996; Papell, 1997; and Coakley et al., 2005a; among others) produced favourable outcomes.[10]

A notorious example of this type of nonstationary finding relates to yen real exchange rate (e.g., Evans and Lothian, 1993; Chortareas and Kapetanios, 2004). It is not surprising, therefore, that the monetary model has explicitly broken down for the dollar-yen exchange rate. Empirical studies have attributed this inadequacy with PPP to the influence of real economic shocks. In particular, the effect that was first formulated by Balassa (1964) and Samuelson (1964) emphasises productivity differences that arise between traded and non-traded goods' sectors. Shocks to government spending may also have implications for the nature of production and the size of the non-traded goods' sector. Furthermore, the real oil price may induce permanent deviations in real exchange rates.

Based on the conjecture that the economy consists of traded and non-traded goods' sectors, the Balassa-Samuelson effect shows that a rise in the productivity of the traded sector results in an increase in the wages of such a sector. Due to internal labour mobility between both sectors within the economy, the prices of non-traded sector's goods will increase as a result of wage increases in the non-traded sector. Incorporating productivity shocks, specifically, into the monetary model of the dollar-yen exchange rate has hypothetical appeal as the business cycles in the US and Japan have undergone differently over the past three decades. Compared to the US, Japan experienced faster economic growth in the 1980s seemingly due to rapid productivity growth in the traded -manufacturing- sector (Chinn, 2000); however, its economy sank into a deep recession

[10] Coakley et al. (2005a), using data from 19 OECD and 26 developing countries over the period 1970: 01-1998: 12, lend support to what they term it as general relative PPP, where the long-run elasticity of the nominal exchange rate with respect to national price ratio is unity without restricting the residuals to be stationary.

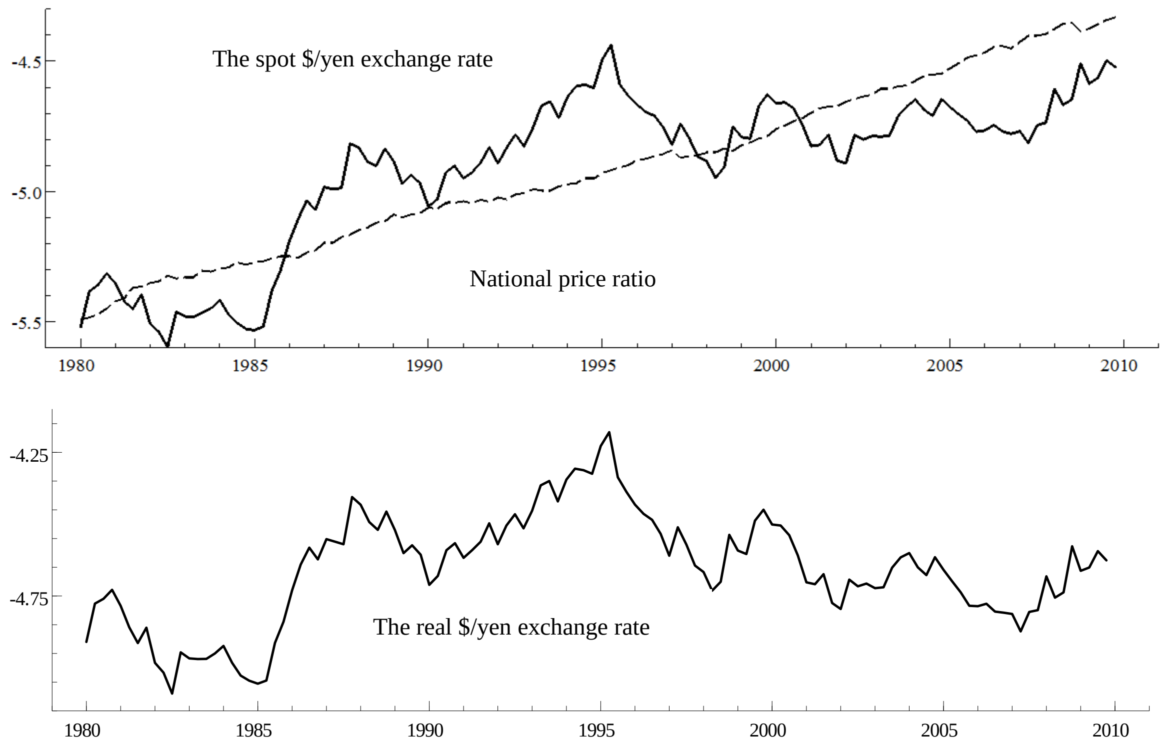

Figure 2.2. The behaviour of the spot dollar/yen exchange rate and national price ratio (upper panel) and the behaviour of the real dollar/yen exchange rate (lower panel) for the period 1980:Q1-2009:Q4.

from the early of 1990s to the middle of 2000s.[11] In the context of the yen-dollar exchange rate, Hsieh (1982), Marston (1987), Yoshikawa (1990), Chinn (1997; 2000), Wang and Dunne (2003), and Tsen (2011) altogether confirmed the impact of the real economic factors on the persistence of the real yen-dollar exchange rate during the post-Bretton Woods period.

The purpose of the graphical analysis below is to support our conjecture of the inadequacy of conventional PPP in explaining the dollar-yen exchange rate. Figure 2.2, upper panel, shows the evolution of the spot dollar-yen exchange rate and the national price ratio between the US and Japan and, lower panel, plots the behaviour of the real dollar-yen exchange rate for the period under investigation. Clearly, the graphical

[11] Relatively low interest rates in Japan over the period 1995-2005 led to the inception of what is known as yen carry trades. Investors borrowed the Japanese yen in order to invest in a high interest currency, mainly the US dollar.

analysis indicates the inadequacy of the conventional PPP for the dollar-yen exchange rate. The real dollar-yen exchange rate appears to be persistent and non-mean reverting. The spot exchange rate, on the other hand, seems comparatively poor in explaining the evolution of the price differential in that its performance appears to be much more erratic than that of the national price ratio.

Juselius and MacDonald (2004) suggested that the behaviour of the spot dollar-yen exchange rate resulted from speculative actions in the capital market rather than the price differential in the goods market. The empirical findings in this chapter show that different components of the shocks that impact on the dollar-yen exchange rate relate to both the real economy and the financial markets (i.e., productivity differential and relative real stock prices).

To this end, real economic factors, namely productivity differentials, relative government spending, and real oil prices may be incorporated into the monetary model to take into consideration the effect of such factors on the persistence of the real dollar-yen exchange rate. This is accomplished by paying an explicit attention to the fact that aggregate price levels in the domestic and foreign country can be decomposed into the prices of traded P_t^T and non-traded goods P_t^{NT}:

$$p_t = (1 - a)\, p_t^T + a p_t^{NT} = p_t^T + a\,(p_t^{NT} - p_t^T), \tag{2.14}$$

$$p_t^* = (1 - a)\, p_t^{*T} + a p_t^{*NT} = p_t^{*T} + a\,(p_t^{*NT} - p_t^{*T}). \tag{2.15}$$

where $a(1-a)$ denotes the proportion of non-traded (traded) goods in the economy. Meanwhile, the real exchange rate is the nominal exchange rate adjusted for the price levels in the domestic and foreign country:

$$q_t = e_t - p_t + p_t^*, \tag{2.16}$$

where q_t denotes the real exchange rate. Substituting the aggregate price levels from Eqs. (2.14) and (2.15) into Eq. (2.16), the real exchange rate can be expressed as follows:

$$q_t = e_t - [p_t^T + a\,(p_t^{NT} - p_t^T)] + [p_t^{*T} + a\,(p_t^{*NT} - p_t^{*T})],$$
$$= (e_t - p_t^T + p_t^{*T}) - a[(p_t^{NT} - p_t^T) - (p_t^{*NT} - p_t^{*T})]. \qquad (2.17)$$

If arbitrage applies primarily to traded goods, then PPP applies only to the traded goods and the component $(e_t - p_t^T + p_t^{*T})$ in Eq. (2.17) should be zero. An example that suggests why this should apply to the yen-dollar exchange rate is provided by Schnabl (2001). Thus, the real exchange rate can be expressed in terms of both traded and non-traded goods as follows:

$$q_t = -\,a\,[(p_t^{NT} - p_t^T) - (p_t^{*NT} - p_t^{*T})]. \qquad (2.18)$$

Eq. (2.18) shows that the relative prices of non-traded to traded goods between the domestic and foreign country determine the real exchange rate. Positive changes in the domestic prices of non-traded to traded goods relative to the foreign counterpart imply an appreciation of the real exchange rate. The deviations from PPP will be permanent if these changes are persistent. The idea that deviations from PPP are explained by relative prices between traded and non-traded goods' prices has gained support not only on a macro basis, but also on a micro level as well. Prompted by the seminal work of Engle and Rogers (1996), who examined the variation of relative prices using disaggregated city data from the US and Canada, both Parsley and Wei (1996) and Cecchetti et al. (2002) discerned the influence of traded and non-traded goods in examining the PPP using respectively panel data from 48 and 19 cities in the US. While Parsley and Wei (1996) observed that the speed of convergence towards PPP is much faster in the traded as compared to the non-traded goods, Cecchetti et al. (2002) concluded that the slow convergence- half-life of nine years – being observed towards PPP is due to the influence of the relatively persistent shocks of non-traded goods compared to the traded ones on overall prices.

Following Strauss (1999), in order to observe the relative price movements of non-traded goods, a market in a competitive world is assumed to hold where firms arrange their price setting in line with unit labour costs in each sector:

$$p_t^T = w_t - prod_t^T, p_t^{NT} = w_t - prod_t^{NT}, p_t^{*T} = w_t^* - prod_t^{*T}, p_t^{*NT} = w_t^* - prod_t^{*NT}, \qquad (2.19)$$

where w_t is the wage rate equated across both the traded and non-traded sectors due to internal labour mobility, while $prod_t^T$ and $prod_t^{NT}$ indicate the productivity in the traded and non-traded sectors, respectively. Hence, relative price movements of non-traded goods are explained by the relative productivity between the traded and non-traded goods' sectors:

$$p_t^{NT} - p_t^T = prod_t^T - prod_t^{NT}, p_t^{*NT} - p_t^{*T} = prod_t^{*T} - prod_t^{*NT}. \qquad (2.20)$$

Substituting Eq. (2.20) into Eq. (2.18), we get the following:

$$q_t = -a[(prod_t^T - prod_t^{*T}) - (prod_t^{NT} - prod_t^{*NT})]. \qquad (2.21)$$

Eq. (2.21) implies that an increase in the traded sector productivity relative to the non-traded sector counterpart results in a fall in the traded sector's goods prices and then an exchange rate appreciation. Chinn (2000), on the other hand, argued that in order to capture demand-side shocks as well, government spending should be taken into account as it is the primary variable of interest in this respect. Intuitively, as government consumption is anticipated to be spent largely on non-traded goods such as services, then government spending should increase the relative price of non-traded goods, giving rise to an exchange rate appreciation. That is, Eq. (2.21) is augmented as follows:

$$q_t = -a[(prod_t^{T} - prod_t^{NT}) - (prod_t^{*T} - prod_t^{*NT})] + \mu\,(gs_t - gs_t^{*}), \qquad (2.22)$$

where $gs_t(gs_t^*)$ denotes the domestic (foreign) government consumption as a percentage of GDP (based on the above discussion, it is expected that $\mu < 0$).

Finally, Amano and van Norden (1998a; 1998b), Chaudhuri and Daniel (1998), Chen and Chen (2007), and Benassy-Quere et al. (2007) showed the influence of the real oil price on the real exchange rate in a bivariate framework. This is also in line with more recent research by Lizardo and Mollick (2010), who concluded that the real oil price explains the movement of the nominal dollar exchange rates to a great extent. This suggests that Eq. (2.22) can be augmented as follows:

$$q_t = -a[(prod_t^{T} - prod_t^{NT}) - (prod_t^{*T} - prod_t^{*NT})] + \mu\,(gs_t - gs_t^{*}) + \gamma\, roil_t, \qquad (2.23)$$

where $roil_t$ is the oil price, represented by the West Texas Intermediate (WTI) spot price, deflated by the US Consumer Price Index (CPI) (as will be explained below, it is expected that $\gamma < 0$). As the availability of quarterly data on the non-traded sector is limited, this leads to the assumption that $prod_t^{NT} = prod_t^{*NT}$. This is in line with Chinn (1997), Wang and Dunne (2003), and Egert (2002a; 2002b), among others. As a result Eq. (2.23) is given by:

$$q_t = -a(prod_t^{T} - prod_t^{*T}) + \mu(gs_t - gs_t^{*}) + \gamma\, roil_t. \tag{2.24}$$

Using Eq. (2.24) along with modified money demand equations given by Eq. (2.1)′ and Eq. (2.2)′, the flexible-price monetary model can be amended, and denoted as MFPM, as follows:

$$e_t = \beta_1(m_t - m_t^*) + \beta_2(y_t - y_t^*) + \beta_3(i_t - i_t^*) + \beta_4(s_t - s_t^*) + \beta_5({prod^T}_t - {prod^{T^*}}_t) + \beta_6(gs_t - gs_t^*) + \beta_7\, roil_t + \varepsilon_t. \qquad (2.5)'$$

In addition to the coefficient restrictions discussed earlier $(\beta_1 = 1, \beta_2 < 0, \beta_3 > 0)$, the modified monetary model suggests that the sign of the coefficient of relative real stock prices, β_4, could be either negative (wealth effect) or positive (substitution effect), as stated earlier. Based on the derivation provided above, the sign of the coefficient on the productivity differential depends on the relative competitiveness of the traded goods sector. Specifically, an increase in the productivity of the traded sector relative to the non-traded sector in the domestic economy compared to the foreign one results in a fall in the domestic traded sector's goods prices relative to the foreign counterpart, and then an exchange rate appreciation $(\beta_5 < 0)$. The differential in government expenditure captures differences in demand side shocks (Chinn, 2000). As government expenditure is anticipated to be spent largely on non-tradable goods such as services, an increase in domestic government spending relative to the foreign counterpart should then increase the relative price of domestic non-tradable goods, leading to an exchange rate appreciation ($\beta_6 < 0$).

With regard to the sign of the coefficient on the real oil price, β_7, Lizardo and Mollick (2010) argue that a higher real oil price results in an appreciation of the US dollar especially against net-importing countries' currencies such as Japan. Japan is in fact considered the third largest oil consumer and importer country after the United States and China. Japanese oil consumption and imports in 2010 were respectively 4452 and 4394 thousands of barrels per day, constituting 23% and 42.7% of the respective consumption and imports of the United States.[12] Hence, funding such huge oil imports has a significant impact on the value of the yen against the US dollar as by international convention oil is priced in dollars and so it is expected that $\beta_7 < 0$; Amano and van Norden (1998b) also provide a theoretical justification of this.

The real interest rate differential model can also be modified, denoted as MRID, and is explained as follows:

$$e_t = \beta_1(m_t - m_t^*) + \beta_2\,(y_t - y_t^*) + \beta_3\,(i_t^s - i_t^{s*}) + \beta_4\,(i_t^l - i_t^{l*}) + \beta_5\,(s_t - s_t^*) + \beta_6\,({prod^T}_t - {prod^{T^*}}_t) + \beta_7\,(gs_t - gs_t^*) + \beta_8\,roil_t + \varepsilon_t. \qquad (2.13)'$$

Based on the aforementioned theoretical explanations, the coefficients are expected to be as follows: $\beta_1 = 1$, $\beta_2 < 0$, $\beta_3 < 0$, $\beta_4 > 0$, $\beta_5 > 0$ or $\beta_5 < 0$, $\beta_6 < 0$, $\beta_7 < 0$, $\beta_8 < 0$. Eq. (2.5)′ and Eq. (2.13)′ modify both the standard flexible-price model, Eq. (2.5), and the standard real interest rate differential model, Eq. (2.13), in the previous subsection to adjust for factors which have a substantial impact on the adequacy of the underlying pillars of these models.

[12] The figures are from CIA World Factbook (2011).

2.3. The econometric approach and data

2.3.1. The econometric approach

Johansen (1988; 1995) developed a maximum likelihood test procedure to investigate cointegration in a multivariate framework. MacDonald (2007, Ch6) argues that the Johansen cointegration technique has many advantages in the context of monetary models of the exchange rate. Such an analysis also bears comparison with recent developments in the study of the long-run relation between exchange rates and monetary fundamentals (e.g., MacDonald and Taylor, 1991; 1994; Tawadros, 2001; Kouretas, 1997; and Cushman 2000; among others).

Johansen (1995) formulates an unrestricted VAR model of order p with n endogenous variables, all integrated of order one ($I(1)$), forced by a vector of ($n \times 1$) independent Gaussian errors, with the following error-correction representation:

$$\Delta X_t = \prod X_{t-1} + \Gamma_1 \Delta X_{t-1} + ... + \Gamma_p \Delta X_{t-(p-1)} + \gamma D_t + \varepsilon_t , \qquad (2.25)$$

where X_t is an ($n \times 1$) vector of variables related to the hypothesised long-run relation that derives from the monetary models, D_t is a vector of constants, centred seasonal dummies[13], and impulse dummies; Γ_i (i =1,....., p-1) are ($n \times n$) parameter matrices capturing the short-run dynamics among the variables, and finally Π is an ($n \times n$) matrix which is partitioned as $\Pi = \alpha\beta'$, where α and β matrices represent the speed of adjustment and long-run parameters, respectively.

Johansen proposed two likelihood ratio tests which represent the key statistics for testing for cointegration, and hence determining the rank r of the long-run matrix Π. The tests are based on the trace and the maximum eigenvalue statistics. However, we

[13] We use centred seasonal dummies because these dummies are averaged to zero.

use the trace test to determine the rank r of Π in this chapter. Johansen (1995) explains that the test has an optimal sequence starting with the null hypothesis $r = 0$ (no cointegration) against the alternative $r \leq 1$ (at least one cointegrating vector) and subsequent further orders of cointegration $r = i$ against the alternative $r \leq i+1$; the sequence stops at $r = i$ when the null cannot be rejected. The trace test is also supported by Monte Carlo evidence (e.g., Lütkepohl et al., 2001) that implies that it has better power performance especially in relatively small samples. The trace test can be written in terms of eigenvalues (λ_i) and sample size (T) with:

$$\lambda_{trace} = -T\sum_{i=1}^{n}(1-\lambda_i). \tag{2.26}$$

As the basis of the Johansen test is an unrestricted VAR model, then the results associated with the Johansen test are well defined when the underlying VAR is well specified (Johansen, 1995). The most appropriate lag length of the VAR model is often based on model selection criteria such as the Akaike Information Criterion (AIC). However, Burke and Hunter (2007) suggest that there can be a substantial size distortion of the trace test relative to the null distribution when the selected lag order is sub–optimal;[14] therefore, we consider extending the VAR models by further lags in the presence of serial correlation.

As a result of sharp changes in monetary policy in the United States and Japan throughout the sample period, we include impulse dummies, which remove the impact of extreme observations relating to 1980:4, 1982:3, 2002:2, and 2008:4. The fourth

[14] As can be seen from their simulations, in the near cointegration case the true DGP is a first order Vector Moving Average model that exhibits considerable size distortion with samples as large as T = 400 observations. It does not go away as the sample evolves.

quarter of 1980 corresponds with the end point of the fiscally liberal 60s and 70s that led Ronald Reagan to enter the presidential office and the associated changes in the fiscal stance meshed with the Volker reforms at the Federal Reserve enacted earlier in the year. The corresponding known events for the other dummies relate to the large short-term interest rate fluctuations in the United States and Japan in the early 1980s, the monetary expansion (now termed Quantitative Easing [QE] policy) adopted by the Bank of Japan from March 2001 to March 2003 in which M1 increased sharply (see Miyao, 2005), and the QE in the United States as a result of the 2007–2008 banking crisis.[15]

Using the above scenario to examine the existence of stable long-run relationships among the variables in monetary models, we are particularly interested in adopting a specific-to-general approach in the econometric estimation of the information set. A similar approach has been used by Juselius and MacDonald (2004) in conducting joint modelling approach of the international parity relations between the US and Japan. We examine the FPM, RID, MFPM, MRID models[16] econometrically by estimating Eq. (2.25) using the following variable vectors in the respective levels:

$$X'_{(FPM)t} = [e_t, m_t - m_t^*, y_t - y_t^*, i_t^s - i_t^{s*}],$$

$$X'_{(RID)t} = [e_t, m_t - m_t^*, y_t - y_t^*, i_t^s - i_t^{s*}, i_t^l - i_t^{l*}],$$

[15] The money supply (M1) increased sharply in the US in late 2008 as liquidity dried up in the banking system after the collapse of Lehman Brothers.

[16] Extending the RID model by domestic and foreign trade balances yields the sticky price model of Hooper and Morton (1982). However, the net trade balance is likely to be a stationary variable as opposed to the rest of the variables. Examination of the sticky price model of Hooper and Morton (1982) and its modification is left for future research.

$$X'_{(MFPM)t}=[e_t, m_t - m_t^*, y_t - y_t^*, i_t^s - i_t^{s*}, s_t - s_t^*, prod_t^T - prod_t^{T*}, gs_t - gs_t^*, roil_t],$$

$$X'_{(MRID)t}=[e_t, m_t - m_t^*, y_t - y_t^*, i_t^s - i_t^{s*}, i_t^l - i_t^{l*}, s_t - s_t^*, prod_t^T - prod_t^{T*}, gs_t - gs_t^*, roil_t].$$

We suggest that by investigating these four variable sets, we might be able to determine the key factors that identify the long-run monetary model of the exchange rate. The monetary model depends on stable money demand relations and the assumption that PPP holds. Using series that are $I(1)$, we can observe an exchange rate equation by finding a cointegrating relation and showing via a likelihood ratio test that this variable is neither long-run excluded (Juselius, 1995), nor weakly exogenous (Johansen, 1992). According to Burke and Hunter (2005), such a finding can help in interpreting and identifying a long-run relation.[17]

2.3.2. Data:

The study employs quarterly seasonally unadjusted (where available) data for the United States *vis-à-vis* Japan over the period 1980:1 – 2009:4. The start of the sample is chosen in order to control the structural changes in the Japanese financial system in which the deregulation of interest rates in the interbank market, the removal of capital controls, and the inception of Certificates of Deposit (CDs) were all accomplished by the end of 1979 (McCornac, 1991).[18] Thus, our sample period is characterised by high international capital mobility between the US and Japan. We make use of quarterly data since data on GDP are not available on a monthly basis so as

[17] The empirical results are obtained using CATS 2.0 in RATS (see Dennis et al., 2005).

[18] These important changes in the Japanese financial system play a significant role in the interactions between financial markets such as foreign exchange, stock, and money markets.

not to limit the analysis by the use of an incomplete measure of national income such as industrial production.

The short-term interest rates are represented by the official discount rates[19], whereas the long-term interest rates are the 10-year government bond yields. Moreover, we use the Consumer Price Index (CPI) to deflate stock price indices in the US and Japan represented by the S&P 500 index and the Nikkei 225 index, respectively. Government spending is defined as government consumption as a percentage of GDP, whereas the productivity is defined as industrial production divided by the corresponding level of employment. The real oil price is the West Texas Intermediate (WTI) Cushing crude oil spot price, dollars per barrel, deflated by the US CPI. The exchange rate (denoted as $/yen), interest rates, income, industrial production for productivity measure and price levels (CPI) are extracted from the IMF's *International Financial Statistics (IFS),* whereas money supply (M1), oil price, and stock prices are retrieved from Thomson DataStream.[20] On the other hand, both government spending construction and employment figures are obtained from the OECD main economic indicators (MEI) database.

Stock prices, exchange rates, oil price, and short-term interest rates are end of period data. All of the variables have been expressed in their logarithmic form except interest rate variables. Finally, the graphs of the variables in levels and first differences are displayed in Appendix B2 (Figures B2.1-B2.2).

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[19] The official discount rate has long been a major policy instrument for the Bank of Japan and other short-term interest rates such as the call rate and bills discount rates have moved in line with the official discount rate (see Ueda, 1996).

[20] With regard to the oil price, it is available in DataStream from 1982 onwards. So, the preceding observations are obtained from the Federal Reserve Bank of St. Louis. These observations to be consistent with the DataStream ones which are end of period, the last month snapshot in each quarter are considered.

2.4. Empirical results

2.4.1. The results of the standard monetary models of exchange rates

A prerequisite step of conducting cointegration tests is checking the time series properties of the variables under investigation as to whether they have a single unit root or not and thus the order of integration. Computing an Augmented Dicky-Fuller (ADF) test (Dickey and Fuller, 1981), the results, as displayed in Table A2.1 (see Appendix A2), indicate that the variables are first difference stationary; thus, they are $I(1)$. As discussed later in the chapter, stationarity tests are also conducted in the multivariate setting by fixing the i^{th} element of a single cointegrating vector to unity and in turn the other elements to zero (to obtain further insight in this regard, see Johansen, 1995).

Having established that the variables are $I(1)$, the empirical analysis in this section involves an examination of standard monetary models of the exchange rate. Hence, the VAR analysis is based on the $X'_{(FPM)t}$ and $X'_{(RID)t}$ vectors. In determining the proper lag length, the AIC indicates a lag length of 1 in the specification of the VAR for both models. However, when $p = 1$ is selected, the misspecification tests show the presence of serial correlation. Thus, we sequentially add lags in order to remove serial correlation in the models. At lag 4, the specifications of the VAR models appear to be improved substantially. The implied misspecification tests of both sets of VAR models at lag 4 are reported in Appendix A2 (Table A2.2, where panel A and B correspond to the standard flexible-price and real interest rate differential monetary models, respectively). Evidently, Table A2.2 shows that, at the 5% level, both models are free from serial correlation, using LM tests of order (8), and ARCH effects, using the ARCH test of order (8). However, the normality tests indicate the existence of non-normality in the residuals only in the standard flexible-price monetary model as a result of excess kurtosis. Since Gonzalo (1994) demonstrated the insensitivity of cointegration to normality as a consequence of excess kurtosis, we conclude that the models are reasonably well-specified.

On the basis of this conclusion, the estimated eigenvalues and trace statistics of the standard monetary models of the exchange rate are reported in Table 2.1. Panel A and B present the results for the standard flexible-price and the standard real interest rate differential monetary models, respectively. As evident from Table 2.1, the trace test indicates that the null hypothesis of no cointegration may be accepted for the standard flexible-price monetary model at the 5% level. However, it is rejected for the standard real interest differential model of Frankel (1979), where there is evidence of one cointergating vector amongst the variables of this model.

The result of no cointergation for the standard flexible-price monetary model is consistent with that of Chinn and Moore (2011), who also failed to uncover cointegration between the dollar-yen exchange rate and its monetary fundamentals in examining the conventional flexible-price monetary model in its restricted form.

Were one to consider the case where $r = 1$, for the standard real interest rate differential model, then the estimated cointegrating vector is reported in Table 2.2 after normalising on the nominal exchange rate as the key variable of interest. An inspection of the results would show that the coefficient on the relative money supply has the sign expected by theory; therefore, based on one-sided inference, we consider it significant at the 5% level. All other variables (relative income, short-term and long-term interest rate differentials) have the wrong sign, though the relative income and the long-term interest rate differential are highly significant. If the real interest rate differential model of Frankel (1979) is accepted as a long-run equilibrium framework for the exchange rate determination, it would yield results that are not theoretically consistent with monetary theory as the estimated coefficients do not have their hypothesised signs.

Table 2.1. Johansen cointegration test results for the standard monetary models.

Panel A. The standard flexible-price monetary model (FPM)

System comprises of $[e, m - m^{*}, y - y^{*}, i^{s} - i^{s*}]$

$(p - r)$	r	Eigenvalue	Trace Test	95% Critical Value	p-value
4	$r = 0$	0.220	47.509	47.707	0.052
3	$r \leq 1$	0.118	18.684	29.804	0.526
2	$r \leq 2$	0.033	4.060	15.408	0.892
1	$r \leq 3$	0.001	0.167	3.841	0.682

Panel B. The standard real interest rate differential monetary model (RID)

System comprises of $[e, m - m^{*}, y - y^{*}, i^{s} - i^{s*}, i^{l} - i^{l*}]$

$(p - r)$	r	Eigenvalue	Trace Test	95% Critical Value	p-value
5	$r = 0$	0.398	100.940	69.611	0.001***
4	$r \leq 1$	0.186	42.054	47.707	0.526
3	$r \leq 2$	0.119	18.219	29.804	0.832
2	$r \leq 3$	0.021	3.534	15.408	0.965
1	$r \leq 4$	0.009	1.053	3.841	0.344

Notes: r denotes the number of cointegrating vectors. *** indicates statistical significance at the 1% level.

Table 2.2. The estimated cointegrating relation for the RID model normalised on the exchange rate.

	$m - m^{*}$	$y - y^{*}$	$i^{s} - i^{s*}$	$i^{l} - i^{l*}$
Coef.	2.250	14.988	0.090	-1.065
t-stat	1.664**	3.301***	1.014	7.732***

Notes: *** and ** indicate statistical significance at the 1% and 5% levels, respectively. The coefficient on relative money supply $(m - m^{*})$ is significant at the 5% level using one sided inference.

It is often felt that normalisation is innocuous, but Boswijk (1996) has suggested that the validity of an identifying restriction requires testing via further rank conditions. However, as shown in Burke and Hunter (2005, ch5), a coherent strategy for identification is to preclude normalisation on variables that are either long-run excluded or weakly exogenous. Furthermore, cointegrating vectors define linear combinations of nonstationary series, thus invalidating normalisation on a stationary variable.

The tests of long-run exclusion (LE), weak exogeneity (WE), and stationarity are asymptotically distributed chi-squared (Johansen, 1992), and in Table 2.3 we report our results on a variable by variable basis. The LE tests indicate that except for the relative income and the long-term interest rate differential, all the other variables can be excluded from the cointegration space. Hence, a long-run model based on the exchange rate may be ill defined, as the related parameter is not different from zero. In the subsequent panel, the proposition that the exchange rate and the short-term interest rate differential are weakly exogenous cannot be rejected. The nominal exchange rate being found weakly exogenous implies that it is not adjusting to the long-run equilibrium and is forcing the system instead. Hence, at best, the long run ought to be conditioned on the exchange rate. Similar results are found in Hunter (1992)[21] and Engle and West (2005), among others. This evidence of the incoherence of the standard RID model casts doubt on the strength of much of the existing empirical evidence where the proposition that the exchange rate is weakly exogenous was not tested despite a large literature that suggest the exchange rate is well explained by a stochastic trend or that it follows a random walk.

On the other hand, the short-term interest rate differential being found as weakly exogenous compared with the long-term interest rate differential is consistent with the term structure of interest rates in that the transmission is running from the nominal short-term interest rate to the long-term interest rate. In the context of a single cointegrating vector this is a less powerful result as it is equally acceptable to normalise on any of the non-exogenous variables in this vector. This result contradicts that of Juselius and MacDonald (2004), who found that the long-run transmission is running

[21] Hunter (1992) finds that a number of variables are weakly exogenous, for the two cointegrating vectors, but in any system there are a maximum of *n-r* variables that satisfy WE. In the final model different restrictions are imposed that suggest a quasi-diagonal structure on α and these along with restrictions on β in terms of the exchange rate give rise to cointegrating exogeneity instead.

Table 2.3. Long-run exclusion, weak exogeneity and the stationarity tests for standard RID model.

	e	$m - m^*$	$y - y^*$	$i^s - i^{s*}$	$i^l - i^{l*}$
Panel A. Long-run exclusion tests					
$x^2(1)$	2.364	1.894	7.241	0.412	27.797
p-value	0.124	0.169	0.000***	0.521	0.000***
Panel B. Long-run weak exogeneity tests					
	e	$m - m^*$	$y - y^*$	$i^s - i^{s*}$	$i^l - i^{l*}$
$x^2(1)$	1.579	18.722	4.751	0.280	16.274
p-value	0.209	0.000***	0.029**	0.597	0.000***
Panel C. Stationarity tests					
	e	$m - m^*$	$y - y^*$	$i^s - i^{s*}$	$i^l - i^{l*}$
$x^2(1)$	47.410	56.959	55.866	46.315	27.492
p-value	0.000***	0.000***	0.000***	0.000***	0.000***

Note: *** and ** indicate statistical significance at the 1% and 5% levels, respectively.

from the long-term to the short-term interest rates when they analysed the international parity relations between the US and Japan.

WE is sensitive to changes in the information set (Juselius and MacDonald, 2004). Using this informative tool in the next subsection will enable us to analyse the long-run interrelations among the variables under consideration in a transparent manner especially the interaction of the nominal exchange rate with other variables in the system. Finally, the stationarity tests confirm that all the series in the system are non-stationary prior to differencing.

In sum, the standard monetary models of the exchange rate appear not to be coherent with the theory and this is consistent with previous studies in the literature. In particular, the results of the standard flexible-price model showed evidence of no cointegration among its variables. With regard to the real interest differential model, it would not seem possible to identify an exchange rate equation solved from the model as a valid explanation of the long-run equilibrium detected by the Johansen methodology.

2.4.2. The results of the modified monetary models of exchange rates

The preceding analysis sheds doubt on the conventional monetary models of the exchange rate. Accordingly, it is of paramount interest to investigate what factors may cause this failure. A *priori*, the possible sources of the inadequacy of the models may be deduced from their underlying building blocks: stable money demand equations and conventional PPP. In this respect, real stock prices reflect a broader understanding of what defines transactions in a monetary model.[22] Nonetheless, the direct impact of stock prices on real money balances was initially considered by Friedman (1988). Real economic variables, on the other hand, may be possible sources of permanent deviations of the real exchange rate and these are used to modify the monetary models. Thus, our analysis involves an examination of modified monetary models of the exchange rate. The VAR for these modified models are based on the $X^{'}_{(MFPM)t}$ and $X^{'}_{(MRID)t}$ vectors.

Since the price of oil is a global factor and all other variables are relatives between the US and Japan in the systems, we treat the real oil price as an exogenous variable in both systems. However, we conduct a long-run weak exogeneity test to confirm this conjecture.[23] The conjecture is also consistent with the intuition of Hamilton (1985) and Amano and van Norden (1998b). These authors argued that the behaviour of the oil price over the past few decades has been governed by major shocks

[22] This follows from Keynes (1936) where financial market efficiency gives rise to a model in which the bond is the representative long-term asset. Hence, interest rates reflect the financial markets, while stocks obtain the same return for an asset with the same risk and term.

[23] Johansen and Juselius (1992) assume that the real oil price is strictly exogenous. Hunter (1992) shows that this corresponds in the long-run to the oil price being WE and LE, but these restrictions were found to be rejected.

which were on the supply-side as a consequence of political instability in the Middle East region, and thence exogenous to a macroeconomic structure. Indeed, in the long-run the real oil price is found to be weakly exogenous with respect to the cointegrating relation in both the modified flexible-price monetary model and the modified real interest differential model with strong non-rejection related to *p*-values of 71.5% and 74.1%, respectively. The real oil price being found as weakly exogenous with respect to the exchange rate is in line with the evidence of Amano and van Norden (1998a; 1998b), Chaudhuri and Daniel (1998), among others.

With regard to the specification of the VARs of both modified models, the AIC indicates a lag length of 3 and 1 for the modified flexible-price and the modified real interest differential models, respectively. Although the diagnostic tests of the former model indicate that such model is well-specified when a lag of 3 is applied, the corresponding tests of the latter with a lag of one do not seem to hold. We sequentially increase the lag of the latter model in order to remove the misspecifications i.e., serial correlation. At lag 3, the model also shows a significant improvement.[24]

It is evident from Table A2.3, displayed in Appendix A2, that both models are free from serial correlation, using LM tests of order (8), and also from ARCH effect, using ARCH tests of order (8). However, the multivariate normality test is rejected. By tracing the origin of such failure in both models, it appears that relative money supply and productivity differential are the sources of such failure and primarily as a result of excess kurtosis. Since Gonzalo (1994) demonstrated the insensitivity of inference on the cointegrating rank as a result of excess kurtosis, it is concluded that these findings are likely to be robust.

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[24] The implied misspecification tests for both models are listed in Appendix A2 (Table A2.3 where panel A and B correspond to the modified flexible-price and real interest differential monetary models, respectively).

The Johansen rank tests related to the modified flexible-price monetary model and the modified real interest rate differential model are reported in Tables 2.4 and 2.5, respectively. It is evident from the trace tests in both Tables that the null hypothesis of non-cointegration is rejected, where there is evidence of one cointergating vector at the 5% level for the two models. Hence, by contrast, the modified flexible-price monetary model is a valid equilibrium framework for the dollar-yen exchange rate. Although the cointegrating rank has not changed in the modified real interest rate differential model, the augmented factors are following a stochastic trend that is common to the nominal exchange rate and monetary fundamentals in the standard real interest differential model.

Tests of long-run exclusion are likely to be informative in this respect as the contribution of the augmented factors can be discerned more thoroughly or the nature of the variables that may be normalised on is reduced. Tables 2.6 and 2.7 report tests of the LE, WE, and stationarity of the variables included in the modified flexible-price and the modified real interest rate differential models, respectively. The stationarity tests in both Tables 2.6 and 2.7 imply that none of the variables in the cointegration spaces of both models is stationary. The LE tests, on the other hand, indicate that most of the variables cannot be excluded from the cointegration spaces of both models especially the variables that have been used to augment these models with the exception of the real oil price. It appears on the basis of this specification that the real oil price and relative income can be excluded from the modified flexible price model (as a block). Similarly, for the modified real interest differential model the results show that the real oil price on which the system has been conditioned, along with the relative money supply and relative real stock prices are not significant, with the latter at best significant at the 11% level for a two-sided inference.

The above analysis is not used as a reason to exclude these variables that seemingly do not belong to a long-run relationship. In the next subsection we follow a general-to-specific methodology in relation to the long-run relations to obtain both parsimonious and better formulated models.

Furthermore, with the exception of the nominal exchange rate, relative spending, and the productivity differential, the long-run WE tests indicate that all variables can be viewed as weakly exogenous in the modified flexible-price model.[25] With regard to the long-run WE tests of the modified real interest differential model, displayed in Table 2.7, it appears that the nominal exchange rate is not weakly exogenous after the extension of the information set. Change in the long-run weak exogeneity status of variables in a model is a de facto indication of changes in the long-run feedback, and hence it is of paramount interest (Juselius and Macdonald, 2004). Thus, in contrast to the standard real interest rate differential model, the findings on weak exogeneity indicate that the nominal exchange rate is adjusting to the long-run equilibrium and not forcing the system after augmenting the model by relative real stock prices, productivity differential, relative government spending, and the real oil price. It appears that all the other variables in the system are weakly exogenous at the 5% level, except the long-term interest rate differential.[26]

[25] It is important to note only *n-r* variables can be found to be weakly exogenous as otherwise the rank condition is violated and that tests on a single variable may not be further supported when joint tests are considered.

[26] With r =1, then there might be *n-1* variables on which the exchange rate is conditioned. This would identify the relationship as a long-run exchange rate equation. However, this finding relies on a joint test of weak exogeneity and the findings suggest the existence of at least another endogenous variable in the system in the long-run.

Table 2.4. Johansen cointegration test results for the MFPM model.

System comprises of [e, $m - m^*$, $y - y^*$, $i^s - i^{s*}$, $s - s^*$, $gs - gs^*$, $prod^T - prod^{T*}$, $roil$]

($p - r$)	r	Eigenvalue	Trace Test	95% Critical Value	p-value
7	$r = 0$	0.407	186.129***	166.049	0.002***
6	$r \leq 1$	0.332	124.915	131.097	0.111
5	$r \leq 2$	0.217	77.680	100.127	0.587
4	$r \leq 3$	0.163	48.984	73.128	0.783
3	$r \leq 4$	0.118	28.155	50.075	0.866
2	$r \leq 5$	0.100	13.507	30.912	0.895
1	$r \leq 6$	0.011	1.236	15.331	0.999

Notes: r denotes the number of cointegrating vectors. *** indicates statistical significance at the 1% level.

Table 2.5. Johansen cointegration test results for the MRID model.

System comprises of [e, $m - m^*$, $y - y^*$, $i^s - i^{s*}$, $i^l - i^{l*}$, $s - s^*$, $gs - gs^*$, $prod^T - prod^{T*}$, $roil$]

($p - r$)	r	Eigenvalue	Trace Test	95% Critical Value	p-value
8	$r = 0$	0.443	230.054***	204.989	0.002***
7	$r \leq 1$	0.389	161.594	166.049	0.085
6	$r \leq 2$	0.245	103.946	131.097	0.630
5	$r \leq 3$	0.191	71.056	100.127	0.799
4	$r \leq 4$	0.144	46.276	73.128	0.864
3	$r \leq 5$	0.121	28.068	50.075	0.869
2	$r \leq 6$	0.095	12.942	30.912	0.917
1	$r \leq 7$	0.011	1.259	15.331	0.998

Notes: r denotes the number of cointegrating vectors. *** indicates statistical significance at the 1% level.

Table 2.6. Long-run exclusion, weak exogeneity, and sationarity tests for the MFPM model.

Panel A. Long-run exclusion tests								
	e	m - m^*	y - y^*	$i^s - i^{s*}$	$s - s^*$	$prod^T - prod^{T*}$	$gs - gs^*$	$roil$
$x^2(1)$	3.398	4.775	1.055	7.585	6.077	13.464	11.471	.002
p-value	.065*	.029**	.304	.006***	.014**	.000***	.001***	.965
Panel B. Long-run weak exogeneity tests								
	e	m - m^*	y - y^*	$i^s - i^{s*}$	$s - s^*$	$prod^T - prod^{T*}$	$gs - gs^*$	
$x^2(1)$	5.967	0.038	0.670	2.662	2.128	4.473	8.982	
p-value	.015**	.845	.413	.103	.145	.034**	.003***	
Panel C. Stationarity tests								
	e	m - m^*	y - y^*	$i^s - i^{s*}$	$s - s^*$	$prod^T - prod^{T*}$	$gs - gs^*$	
$x^2(1)$	40.174	40.509	39.995	40.483	30.971	31.827	35.169	
p-value	.000***	.000***	.000**	.000***	.000***	.000***	.000***	

Note: ***, ** , and * indicate statistical significance at the 1% , 5% and 10% levels, respectively.

Table 2.7. Long-run exclusion, weak exogeneity, and stationarity tests for the MRID model.

Panel A. Long-run exclusion tests									
	e	m - m^*	y - y^*	$i^s - i^{s*}$	$i^l - i^{l*}$	$s - s^*$	$prod^T$ -$prod^{T*}$	$gs - gs^*$	$roil$
$x^2(1)$	5.641	2.058	4.556	9.756	7.942	2.561	6.057	9.163	0.845
p-value	.018**	.150	.033**	.002***	.005***	0.110	.014**	.002***	.358
Panel B. Long-run weak exogeneity tests									
	e	m -m^*	y - y^*	$i^s - i^{s*}$	$i^l - i^{l*}$	$s - s^*$	$prod^T$ -$prod^{T*}$	$gs - gs^*$	
$x^2(1)$	3.978	0.192	0.225	0.006	4.220	2.461	1.627	3.261	
p-value	.046**	.662	.636	.939	0.040**	.117	.202	.071*	
Panel C. Stationarity tests									
	e	m -m^*	y - y^*	$i^s - i^{s*}$	$i^l - i^{l*}$	$s - s^*$	$prod^T$ -$prod^{T*}$	$gs - gs^*$	
$x^2(1)$	38.03	38.25	37.13	39.45	23.56	30.83	29.99	33.191	
p-value	.000***	.000***	.000***	.000***	.000***	.000***	.000***	.000***	

Note: $^{***, **}$ and * indicate statistical significance at the 1%, 5% and 10% levels, respectively.

Overall, the results imply that the long-run status of the nominal exchange rate, relative money supply and relative income have changed after extending the standard RID model by factors likely to affect money demand stability and deviations from PPP. These results are also consistent with the monetary mainstream. The long-run feedback reverse for the nominal exchange rate and the acceptance of weak exogeneity individually of relative real stock prices and real economic variables (productivity differential, relative government spending, and real oil price) indicate an important policy implication. The finding suggests that the cumulated shocks to the nominal exchange rate originate from shocks that relate to both the real economy and financial markets i.e., productivity differential, relative government spending, real oil price, and relative real equity prices. Lastrapes (1992), Chen and Wu (1997), and Enders and Lee (1997) altogether confirmed the dominance of real shocks on the nominal and real exchange rates.

The long-run weak exogeneity states of the short-term as well as the long-term interest rate differentials have not changed, consistent with the term structure of interest rates. This is another piece of important information in the conduct of monetary policy, especially in the light of the findings in the literature on the term structure of interest rates.

On the basis of these findings, the cointegrating vectors in the modified monetary models are normalised on the nominal exchange rate, listed in Table 2.8 in Panel A and B. The results of the modified models indicate a substantial improvement over the standard monetary models. With regard to the modified flexible price monetary model, the results show that all monetary fundamentals of the model are statistically significant. The coefficient of relative money supply has the expected positive sign, significant and reasonably close to one. The coefficient of relative income also has a sign as is expected and statistically significant. The interest rate differential is also significant, but it has a negative coefficient. The interest rate being found negative is consistent with the sticky-price monetary model instead of the flexible-price monetary model. It indicates that a higher domestic interest rate relative to the foreign counterpart induces capital inflows, and hence nominal exchange rate appreciation.

The results of the modified real interest differential model, on the other hand, are also consistent with monetary theory. Strictly speaking, the coefficient of relative money supply is numerically close to one and significant at the 5% level, based on a one-sided test. All other monetary variables (relative income, short-term and long-term interest rate differentials) have their *prior* hypothesised signs and significant at the 1% level. Furthermore, as hypothesised by Frankel (1979), the parameter on the long-term interest rate differential is greater than that on the short-term interest rate differential in absolute value.

With regard to the factors that have been augmented to the models, all of them are significant, except the real oil price. The relative real stock price coefficient in both models is negative. This implies that the wealth effect dominates the money demand functions in the US and Japan, consistent with the evidence of Friedman (1988), McCornac (1991) and Caruso (2001). The coefficients of the productivity differential in industry and relative government spending in both models are also negative and significant. Thus, a higher domestic productivity or government spending relative to the foreign counterpart results in an exchange rate appreciation.

The effect of the real oil price in both models is weak which is surprising, though this is consistent with the suggestion in Johansen and Juselius (1992), where the real oil price to be treated as a strictly exogenous. That is, it only impacts the long-run indirectly via the short-run dynamics. It is negative in the modified flexible-price monetary model, consistent with theory, and positive in the modified real interest rate

Table 2.8. The estimated cointegrating vectors of the modified monetary models of exchange rates.

Panel A. The modified flexible-price monetary model

	$m-m^*$	$y-y^*$	i^s-i^{s*}	i^l-i^{l*}	$s-s^*$	$prod^T-prod^{T*}$	$gs-gs^*$	$roil$
Coef.	1.40	-2.956	- 0.129	_	-0.768	-9.786	-15.601	-0.008
t-stat	2.55***	1.85*	3.87***	_	3.96***	6.49***	9.93***	0.059

Panel B. The modified real interest rate differential monetary model

	$m-m^*$	$y-y^*$	i^s-i^{s*}	i^l-i^{l*}	$s-s^*$	$prod^T-prod^{T*}$	$gs-gs^*$	$roil$
Coef.	0.935	- 5.524	-0.214	0.262	- 0.477	-7.822	-12.420	0.205
t-stat	1.77**	3.61***	5.46***	4.83***	2.52***	5.52***	8.34***	1.25

Notes: ***, ** and * indicate statistical significance at the 1%, 5% and 10% levels, respectively. The coefficient on relative money supply ($m-m^*$) is significant at the 5% level using one sided inference.

differential model. The weak effect of oil price may be due to the importance of the other factors in the models in governing the movement of the dollar-yen exchange rate.

2.4.3. Validation of the modified monetary models of exchange rates

The above results strongly infer the inadequacy of the conventional monetary models of Bilson (1978) and Frankel (1979), applied to the dollar-yen exchange rate, as a result of the breakdown of their underlying building blocks: stable money demand relations and conventional PPP. Once we have accounted for factors that affect these two building blocks, the monetary models are considerably improved. However, in order to obtain a parsimonious and robust formulation of the above long-run relationships, we follow the general-to-specific approach (Hendry and Mizon, 1993) subject to the results on long-run exclusion and weak exogeneity tests. Having detected that $r=1$ in both modified models, the following structure related to the α and β vectors is observed for the modified flexible-price model. First:

$$\begin{array}{lcccccccc} & e & dm & dy & di^s & ds & dprod^T & dgs & roil \\ \alpha' = & [\alpha_0 & \alpha_1 & \alpha_2 & \alpha_3 & \alpha_4 & \alpha_5 & \alpha_6 & 0], \end{array}$$

with α_7= 0 imposed for the weak exogeneity of the real oil price and the notation *d* represents the differential of the variables for the US and Japan. Second:

$$\begin{array}{lcccccccc} & e & dm & dy & di^s & ds & dprod^T & dgs & roil \\ \beta' = & [-1 & \beta_1 & \beta_2 & \beta_3 & \beta_4 & \beta_5 & \beta_6 & 0], \end{array}$$

with the only restriction imposed on β associated with the normalisation on the exchange rate (β_0= -1). Likewise, the α and β vectors for the modified real interest rate differential model are as follows:

$$\begin{array}{lccccccccc} & e & dm & dy & di^s & di^l & ds & dprod^T & dgs & roil \\ \alpha' = & [\alpha_0 & \alpha_1 & \alpha_2 & \alpha_3 & \alpha_4 & \alpha_5 & \alpha_6 & \alpha_7 & 0], \end{array}$$

with α_8= 0 imposed for the weak exogeneity of the real oil price and (β_0= -1) for the normalisation on the exchange rate:

$$\begin{array}{lccccccccc} & e & dm & dy & di^s & di^l & ds & dprod^T & dgs & roil \\ \beta' = & [-1 & \beta_1 & \beta_2 & \beta_3 & \beta_4 & \beta_5 & \beta_6 & \beta_7 & 0]. \end{array}$$

On the basis of this structure, the long-run exclusion tests show strongly the exclusion of the augmented real oil price (see Tables 2.6 and 2.7), hence β_7= 0 for the MFPM and β_8= 0 for the MRID. Conditioning on such an exclusion for each model, we also sequentially impose zero restrictions on the loading factors, α, of the standard monetary fundamentals (relative money supply, relative income, and short-term interest rate differential) in both models as the implicit weak exogeneity restrictions of such variables are consistent with the monetary mainstream and are plausible given the small size of the adjustment coefficients.[27]

We do not impose the weak exogeneity of the long-term interest rate differential in the modified real interest rate differential model as it appears to be endogenous and this is consistent with the term-structure of interest rates, as stated earlier. The particular tests, displayed in Table 2.9, indicate that the imposed restrictions are strongly accepted and the constrained final long-run relationships normalised on the exchange rate indicate the significance of all variables with their hypothesised *prior* signs. This indicates the robustness of our results in terms of the long-run formulation and the importance of the augmenting factors in explaining the conventional monetary models of the exchange rate. More specifically, the coefficient on the relative money supply in the two modified models is correctly signed as opposed to the corresponding recent reported results of Lizardo and Mollick (2010) and Chinn and Moore (2011) for the conventional monetary model of the dollar-yen exchange rate. However, the coefficient on the relative income is larger in absolute value; it is -3.4 (-4) for the modified flexible-price (real interest differential) model compared to -1.8 in Chinn and Moore (2011). Lizardo and Mollick (2010), by contrast, found such a coefficient to be wrongly signed. The coefficient on the short-term interest rate differential is smaller in absolute value; that is, it is -0.12 (-0.16) for the modified flexible-price (real interest differential) model compared to -0.51 in Chinn and Moore (2011) which was being found insignificant.

[27] The long-run weak exogeneity tests of these variables confirmed their status being weakly exogenous with respect to the exchange rate in the long-run relationships of the two modified models, see Table 2.6 and 2.7.

Table 2.9. Joint tests of weak exogeneity and long-run exclusion conditional on *r* =1.

Panel A. The modified flexible-price monetary model

Tests under the null:	Statistics [*p*-value]
(i) $\beta_7 = 0$	$x^2(1) = 0.002$ [0.965]
(ii) $\beta_7 = 0, \alpha_1 = 0$	$x^2(2) = 0.058$ [0.971]
(iii) $\beta_7 = 0, \alpha_1 = 0, \alpha_2 = 0$	$x^2(3) = 0.797$ [0.850]
(iv) $\beta_7 = 0, \alpha_1 = 0, \alpha_2 = 0, \alpha_3 = 0$	$x^2(4) = 3.016$ [0.555]

The implied long-run relationship by test (iv):

	$m - m^*$	$y - y^*$	$i^s - i^{s*}$	$s - s^*$	$prod^T - prod^{T*}$	$gs - gs^*$
Coef.	1.472	- 3.416	- 0.129	- 0.603	- 8.726	- 14.620
t-stat	-3.17***	2.47***	4.37***	3.51***	6.46***	10.60***

Panel B. The modified real interest rate differential model

Tests under the null:	Statistics [*p*-value]
(i) $\beta_8 = 0$	$x^2(1) = 0.845$ [0.358]
(ii) $\beta_8 = 0, \alpha_3 = 0$	x^2 (2) = 0.856 [0.652]
(iii) $\beta_8 = 0, \alpha_3 = 0, \alpha_1 = 0$	x^2 (3) = 1.690 [0.639]
(iv) $\beta_8 = 0, \alpha_3 = 0, \alpha_1 = 0, \alpha_2 = 0$	x^2 (4) = 2.010 [0.734]

The implied long-run relationship by test (iv):

	$m - m^*$	$y - y^*$	$i^s - i^{s*}$	$i^l - i^{l*}$	$s - s^*$	$prod^T - prod^{T*}$	$gs - gs^*$
Coef.	0.740	- 4.028	-0.169	0.172	- 0.557	- 6.748	-11.23
t-stat	-1.74**	3.36***	5.75***	4.52***	3.66***	5.74***	9.09***

Notes: *** and ** indicate statistical significance at the 1% and 5% levels, respectively. The coefficient on relative money supply ($m - m^*$) is significant at the 5% level using one sided inference.

Then, we subject our proposed modified monetary models of the exchange rate to the array of forward and backward recursive stability tests proposed by Hansen and Johansen (1999) to gain further insights into the adequacy of these models. The tests are on the eigenvalues λ_i, for the constancy of the log-likelihood, and the max test of β, displayed respectively in Figures C2.1-C2.3 (see Appendix C2) and D2.1-D2.3 (see Appendix D2) for the modified flexible-price and modified real interest differential models (left and right panels display forward and backward tests, respectively).[28] The corresponding 5% critical value is represented by the solid line.

The base samples of the modified flexible–price model for the forward and backward tests are 1980:Q4-1995:Q3 and 2009:Q4-1995:Q1, respectively**.** The base samples of the modified real interest rate differential model, on the other hand, are respectively 1980:Q4-1998:Q2 and 2009:Q4-1992:Q2 for the forward and backward tests. The forward test hinges on the estimation of the base sample and then it is recursively extended by one observation until the whole sample is restored. The backward test starts with the base sample. Then, the base sample is recursively extended backward until the whole sample is covered. In discussing these tests, the short-run effects are concentrated out (X(t)).

Broadly speaking, both models show a reasonable degree of stability of the parameters in their corresponding cointegrating vectors. The graphs of R1(t) in all cases lie below the line that indicates the 5% critical value. This is aside from the forward recursive test of the constancy of the cointegrating relation (β) for the modified flexible-price model, where the corresponding graph violates the line for the period of the Asian financial crisis in 1997 and returns to a regime related to parameter constancy in 2001. Since the said graph returns to a parameter constancy regime in a short period and all other forward and backward tests of the modified flexible-price model exhibit parameter constancy, the stability of the overall model is secured. Overall, both models appear to be adequate and do not exhibit any structural breaks over the period under observation.

2.4.4. The error correcting models and the out-of-sample forecasting

Having found the existence of a long-run equilibrium relationship among the variables of the two modified monetary models of the exchange rate in subsection 2.4.2,

[28] See also Dimitraki and Menla Ali (2013) for the use of these recursive stability tests.

in this subsection we formulate the inherent dynamic error correction models in order to examine the short-run dynamics among the variables in the two models. Yet, in order to diagnose the performance of the proposed models in terms of out-of-sample forecasting, the last 8 observations are reserved. In particular, the models are estimated for the period 1980:Q1-2007:Q4, then out-of-sample forecasts at different horizons, namely 3, 4, 6, and 8 are obtained. Following the common practice in the literature, rolling window technique is employed to conduct the estimation which starts 2008:Q1.

As far as the estimation of the short-run dynamics is concerned, the long-run weak exogeneity findings are instructive in the formulation. It follows that the error correction model for the modified flexible-price monetary model would include 3 simultaneous equations for the variables Δe_t, $\Delta prod_t^{Td}$ and Δgs_t^{Td} (the superscript d denotes the differential of the variables for the US and Japan) and conditioned on the remaining weakly exogenous variables at the 5% level. The error correction model for the modified real interest rate differential model, on the other hand, is based on two simultaneous equations for the variables Δe_t and Δi_t^{ld} and conditioned on the rest of weakly exogenous variables at the 5% level. However, since the central variable of interest in this study is the exchange rate, we design only the implied single error correction models of the nominal exchange rate for both models. That is, the error correction models are constructed for both modified models as follows:

The modified flexible-price monetary model:

$$\Delta e_t = \varphi_0 + \gamma_1 ect_{1t-1} + \sum_{i=1}^{k} \varphi_{1i} \Delta e_{t-i} + \sum_{i=1}^{k} \varphi_{2i} \Delta m_{t-i}^{d} + \sum_{i=1}^{k} \varphi_{3i} \Delta y_{t-i}^{d} + \sum_{i=1}^{k} \varphi_{4i} \Delta i_{t-i}^{sd} + \sum_{i=1}^{k} \varphi_{5i} \Delta s_{t-i}^{d} + \sum_{i=1}^{k} \varphi_{6i} \Delta prod_{t-i}^{Td} + \sum_{i=1}^{k} \varphi_{7i} \Delta gs_{t-i}^{d} + \sum_{i=1}^{k} \varphi_{8i} \Delta roil_{t-i} + \xi_1 S_1 + \xi_2 S_2 + \xi_3 S_3 + \delta_1 D_{1982:Q3} + \delta_2 D_{2002:Q2} + u_t, \quad (2.27)$$

The modified real interest differential monetary model:

$$\Delta e_t = \theta_0 + \gamma_2 ect_{1t-1} + \sum_{i=1}^{k} \theta_{1i} \Delta e_{t-i} + \sum_{i=1}^{k} \theta_{2i} \Delta m_{t-i}^{d} + \sum_{i=1}^{k} \theta_{3i} \Delta y_{t-i}^{d} + \sum_{i=1}^{k} \theta_{4i} \Delta i_{t-i}^{sd} + \sum_{i=1}^{k} \theta_{5i} \Delta i_{t-i}^{ld} + \sum_{i=1}^{k} \theta_{6i} \Delta s_{t-i}^{d} + \sum_{i=1}^{k} \theta_{7i} \Delta prod_{t-i}^{Td} + \sum_{i=1}^{k} \theta_{8i} \Delta gs_{t-i}^{d} + \sum_{i=1}^{k} \theta_{9i} \Delta roil_{\,t-i} + \omega_1 S_1 + \omega_2 S_2 + \omega_3 S_3 + \lambda_1 D_{1982:Q3} + \lambda_2 D_{2002:Q2} + v_t, \quad (2.28)$$

where u_t and v_t are error terms assumed to be white noise; S_1, S_2, and S_3 indicate quarterly centred seasonal dummies; $D_{1982:Q3}$ and $D_{2002:Q2}$ are two monetary impulse dummies; ect_{1t-1} and ect_{2t-1} denote the lagged error correction terms considered by normalising the cointegrating vector on the nominal exchange rate e_t in both the modified flexible-price and the modified real interest differential models, respectively, and finally γ_1 and γ_2 represent the corresponding speed of adjustments towards the long-run equilibrium in both models.

To estimate the above equations, Eqs. (2.27) and (2.28), we utilise the 'general-to-specific' approach of Hendry and Mizon (1993) in order to obtain the parsimonious models. That is, we allow for six lags in each model, as a preliminary step, which are sufficient to capture the dynamics in the data in addition to the dummy variables specified above, then we exclude the least insignificant variables sequentially.[29] The obtained final short-run dynamics for both models are presented in Table A2.4 in Appendix A2, which include the significant coefficients at least at the 10% level. As shown from Table A2.4, the two models are well-specified with no evidence of ARCH effects, normality and serial correlation.

Moreover, the error correction terms in both models are negative and significant, implying that the exchange rate corrects to the long-run equilibrium in the models. In

[29] In conducting the 'general-to-specific' approach, we not only sequentially remove the insignificant coefficients based on *t*-statistics and *F*-statistics, but also we make use of the PcGive (version 13)' PROGRESS feature which observes the progress of the model as the reduction sequentially proceeds.

terms of the coefficients signs, conventional exchange rate theories do not indicate any hints about the short-run dynamics of the monetary models. The only exception is the Dornbusch (1976) overshooting model. In a broad sense, the results indicate that most of the variables are consistent with monetary mainstream and the short-run coefficients are much smaller than the long-run counterpart. This is in line with the existing common practice in the literature (see, for example, Civcir, 2003).

Turning the focus onto the out-of-sample forecasting performance, the Root Mean Square Errors (RMSEs) for both models at the considered horizons are calculated and compared with the random walk model with drift and without drift in order to assess the predictive power of the models. The results of the out-of-sample forecasting performance are displayed in Table 2.10. Several points are noteworthy. The results show that the implied dynamic error correction models outperform the random walk benchmark at all horizons, except horizon 3. A similar finding was obtained by Frömmel et al. (2005) using the Markov switching approach. It is evident from Table 2.10 that as the forecasting horizon is extended, the predictive power of the models rises substantially in relative to the random walk benchmark.

The fact that the modified monetary models of the exchange rate provide great improvement in the medium and long-term in relative to the short-term signifies the important role of real economic variables at loner horizons. A tentative explanation of the inadequacy of the models to beat the random walk in the short-term is the omission of factors capturing capital flows between the US and Japan. It is proved that capital flows play a significant role in the movements of the exchange rates in the short-run as real macroeconomic variables do in the long-run. Another explanation for the overthrow of the models in the short-term and at horizon 3, in particular, is that such a forecasting horizon corresponds to the financial turmoil ensued the collapse of the Lehman Brothers in September 15, 2008.

Table 2.10. The out-of-sample forecasts: Random walk versus modified monetary models.

Horizon	Statistics	Random walk without drift	Random walk with drift	Modified flexible-price model	Modified real interest differential model
3	RMSE	0.086506	0.099894	0.10619	0.097733
4	RMSE	0.12441	0.14366	0.10453	0.10958
6	RMSE	0.12330	0.15054	0.086668	0.11936
8	RMSE	0.13712	0.17351	0.084201	0.11897

Notes: The estimation period is from 1980:Q1 to 2007:Q4, while the forecasting period starts from 2008:Q1 to 2009:Q4.

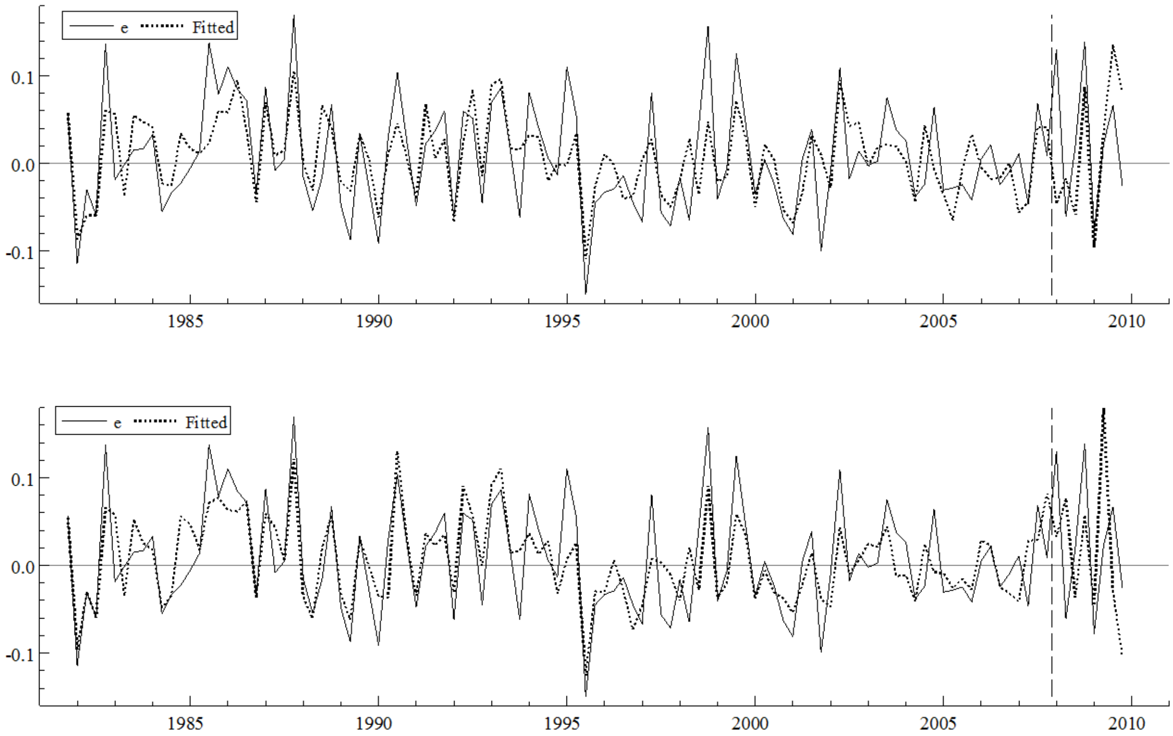

Figure 2.3. The actual and the fitted values from the modified flexible-price model (upper panel) and the modified real interest differential model (lower panel).

Further inference can be obtained by taking a close look at the plot of the actual and fitted values from both models in Figure 2.3 for the in-sample, as well as the out-of-sample forecasting period. Upper (lower) panel displays the fitted and the actual values for the modified flexible-price (real interest differential) monetary model. Visual inspection indicates that both models perform reasonably well in tracking the turning points of the actual exchange rate for the in-sample, as well as the out-of-sample, especially at longer and medium horizons.

2.5. Conclusions

In this chapter, we conduct a thorough empirical scrutiny of the conventional flexible-price monetary model of Bilson (1978), as well as the conventional real interest rate differential monetary model of Frankel (1979), applied to quarterly dollar-yen exchange rate over the period 1980-2009. This particular period is characterised by high international capital mobility between the US and Japan, as well as high volatility of the US dollar against the Japanese yen. By employing the Johansen cointegration technique and long-run exclusion and weak exogeneity tests, we demonstrate that the breakdown of the two conventional monetary models of the dollar-yen exchange rate is due to the breakdown of their underlying building blocks: stable money demand and PPP. Accounting the monetary models for factors affecting these building blocks provides supportive results. In particular, adjusting the models for real stock prices to capture the stability of money demand on one hand and also for real economic variables such as productivity differential, relative government spending, and real oil price to explain the persistence in the real exchange rate on the other provide long-run relationships that appear consistent with the monetary models.

The enhanced performance of the modified models derives from the following considerations to the conventional monetary models. First, the stability of money demand relations is taken into account by the inclusion of key variables that impact on transactions (Friedman, 1988). A key feature of globalised financial markets is a highly active market in cross-border investments, mergers and acquisitions, and cross-listed stocks. In particular, the futures contract on the Nikkei is listed as an asset in the US stock market. Second, the persistence of the real exchange rate, which reflects primarily the impact of the non-traded goods, is taken into consideration by accounting for productivity and government expenditure differences. In essence, these differences may be due to the relatively insular nature of Japanese society limiting the effectiveness of arbitrage. The literature also suggests that the real oil price affects such persistence, but the empirical findings herein show an indirect impact of such a price via the dynamic specification of the VAR model.

Contrary to the conventional monetary models, the results also suggest that the dollar–yen exchange rate in the modified models is driven by money, income, and short-term interest rate differentials, but not the reverse. The results of the out-of-sample forecasting of the proposed modified models also show their superiority over the random walk benchmark in the medium- and long-term, but not the short-term. This implies a substantial role for real economic and financial market variables in a well-formulated monetary model for the determination of the exchange rate.

Appendix A2

Table A2.1. Augmented Dickey–Fuller unit root test results.

Variable	Levels		First differences	
	Without trend	With trend	Without trend	With trend
e	-1.420 (3)	-2.375 (4)	-5.217 (2)***	-5.206 (2)***
$m - m^*$	-1.450(4)	-2.142 (4)	-3.403 (4)**	-3.428 (4)
$y - y^*$	-0.516 (3)	-2.250 (3)	-3.753 (2)***	-3.905(2)***
$i^s - i^{s*}$	-2.752 (1)	-2.725 (1)	-4.877 (4)***	-4.837 (4)***
$i^l - i^{l*}$	-2.408 (4)	-2.392 (1)	-5.379 (4)***	-5.324 (4)***
$s - s^*$	-.4927 (1)	-1.990 (1)	-6.634 (1)***	-6.621 (1)***
$gs - gs^*$	-.2737 (3)	-2.296 (3)	-4.301 (3)***	-4.321 (3)***
$prod^T - prod^{T*}$	-.8268 (4)	-3.443 (4)	-6.621 (4)***	-5.666 (4)***
$roil$	-1.964 (2)	-1.994 (2)	-9.764 (1)***	-6.876 (4)***

Note: the 1% and 5% critical values for the ADF tests are respectively -3.486 and -2.885 (without trend) and -4.03, -3.448 (with the trend); the proper lag length is selected by the AIC, representing in parentheses. *** and ** indicate statistical significance at the 1% and 5% levels, respectively.

Table A2.2. Misspecification tests of the standard monetary models.

Panel A. The standard flexible-price monetary model (FPM)

Equations	ARCH(8)	Normality	Skewness	Ex. Kurtosis
		Single tests (k=4)		
e	0.318 [0.957]	0.027 [0.049]	0.425	4.212
$m - m^*$	1.105 [0.365]	7.978 [0.018]	-0.007	4.138
$y - y^*$	1.028 [0.419]	4.100 [0.128]	0.279	3.608
$i^s - i^{s*}$	0.567 [0.802]	6.660 [0.035]	-0.477	4.632
		System tests		
LM(8)	1.200 [0.116]			
Normality	24.47 [0.001]			

Panel B. The standard real interest rate differential model (RID)

Equations	ARCH(8)	Normality	Skewness	Ex. Kurtosis
		Single tests (k=4)		
e	0.188 [0.992]	4.414 [0.110]	0.442	3.063
$m - m^*$	0.718 [0.674]	2.676 [0.262]	0.054	3.480
$y - y^*$	0.079 [0.999]	2.628 [0.268]	0.194	3.430
$i^s - i^{s*}$	0.609 [0.768]	5.308 [0.070]	-0.566	3.286
$i^l - i^{l*}$	1.139 [0.343]	0.916 [0.632]	-0.027	3.246
		System tests		
LM(8)	1.204 [0.088]			
Normality	17.25 [0.068]			

Notes: k denotes number of lags, whereas p-values are in square brackets [.].

Table A2.3. Misspecification tests of the modified monetary models.

Panel A. The modified flexible-price monetary model (MFPM)

	Single tests (k=3)			
Equations	ARCH(8)	Normality	Skewness	Ex. Kurtosis
e	0.790 [0.612]	4.443 [0.108]	0.344	3.261
$m - m^*$	0.745 [0.651]	13.52 [0.001]	-0.299	4.760
$y - y^*$	1.259 [0.273]	3.616 [0.163]	0.385	2.919
$i^s - i^{s*}$	0.648 [0.735]	5.368 [0.068]	-0.287	3.759
$s - s^*$	0.852 [0.559]	2.562 [0.277]	0.300	3.270
$gs - gs^*$	0.845 [0.564]	2.375 [0.305]	-0.216	3.617
$prod^T - prod^{T*}$	0.577 [0.849]	41.42 [0.000]	1.019	7.614
	System tests			
LM(8)	1.083 [0.292]			
Normality	69.43 [0.000]			

Panel B. The modified real interest rate differential model (MRID)

	Single tests (k=3)			
Equations	ARCH(8)	Normality	Skewness	Ex. Kurtosis
e	0.641 [0.741]	4.105 [0.128]	0.348	3.239
$m - m^*$	1.265 [0.270]	14.88[0.000]	-0.243	4.745
$y - y^*$	0.866 [0.547]	1.320 [0.516]	0.298	2.973
$i^s - i^{s*}$	0.840 [0.569]	2.305 [0.315]	-0.343	3.349
$i^l - i^{l*}$	1.465 [0.179]	5.706 [0.057]	-0.307	2.748
$s - s^*$	0.915 [0.507]	2.672 [0.262]	0.300	3.431
$gs - gs^*$	0.901 [0.518]	1.990 [0.369]	-0.055	3.544
$prod^T - prod^{T*}$	0.455 [0.884]	41.80 [0.000]	0.857	7.317
	System tests			
LM(8)	1.272 [0.052]			
Normality	63.80 [0.000]			

Notes: k denotes number of lags, whereas p-values are in square brackets [.].

Table A2.4. The estimates of the error correction models.

Panel A. The modified flexible-price model			*Panel B. The modified real interest differential model*		
Variables	Coef	*t*-stat	Variables	Coef	*t*-stat
Constant	0.151	2.01	Constant	0.252	6.19
Δe_{t-2}	-0.284	-3.28	Δe_{t-2}	-0.235	-2.25
Δe_{t-3}	0.200	2.34	Δe_{t-4}	-0.170	-1.69
Δm_{t-3}^{d}	0.395	3.21	Δe_{t-6}	0.184	2.02
Δm_{t-4}^{d}	0.340	2.49	Δm_{t-3}^{d}	0.685	4.12
Δm_{t-6}^{d}	-0.430	-2.32	Δy_{t-1}^{d}	1.530	2.92
Δy_{t-2}^{d}	-0.814	-1.80	Δy_{t-5}^{d}	2.175	3.33
Δi_{t-5}^{sd}	-0.019	-2.50	Δi_{t-1}^{sd}	-0.021	-1.82
Δs_{t-2}^{d}	-0.098	-1.82	Δi_{t-3}^{sd}	-0.0259	-2.03
Δs_{t-4}^{d}	-0.099	-1.78	Δi_{t-5}^{sd}	-0.0361	-4.20
Δs_{t-5}^{d}	0.137	2.47	Δi_{t-1}^{ld}	0.0307	2.62
$\Delta prod_{t-1}^{Td}$	-0.416	-2.40	Δi_{t-2}^{ld}	0.028	2.72
$\Delta prod_{t-4}^{Td}$	0.361	2.38	Δi_{t-3}^{ld}	0.026	2.29
Δgs_{t-3}^{d}	1.007	2.60	Δi_{t-6}^{ld}	0.032	3.29
Δgs_{t-6}^{d}	-0.707	-1.92	Δs_{t-5}^{d}	0.166	2.73
$\Delta roil_{t-6}$	-0.069	-2.27	$\Delta prod_{t-1}^{Td}$	0.862	-3.56
ect_{t-1}	-0.051	-4.10	$\Delta prod_{t-2}^{Td}$	-0.860	-3.78
S_1	-0.055	-3.81	$\Delta prod_{t-3}^{Td}$	-0.722	-3.56
S_2	-0.060	-2.93	$\Delta prod_{t-5}^{Td}$	-0.496	-2.33
$D_{1982:Q3}$	-0.139	-2.60	Δgs_{t-3}^{d}	0.899	2.02
			Δgs_{t-4}^{d}	0.801	1.90
			Δgs_{t-5}^{d}	1.220	2.30
			Δgs_{t-6}^{d}	-1.112	-2.70
			$\Delta roil_{t-1}$	0.103	3.15
			$\Delta roil_{t-4}$	0.105	3.15
			S_1	-0.049	-2.94
			S_2	-0.029	-1.94
			ect_{t-1}	-0.128	-5.70
			$D_{1982:Q3}$	-0.153	-2.55
σ	0.047		σ	0.046	
LM (8)	1.107 [0.374]		LM(8)	0.989 [0.457]	
ARCH (8)	0.416 [0.908]		ARCH (8)	0.598 [0.776]	
Normality	2.565 [0.277]		Normality	0.098 [0.952]	

Notes: The superscript *d* denotes the differential of the variables for the US and Japan, while *p*-values are reported in square brackets [.].

Appendix B2

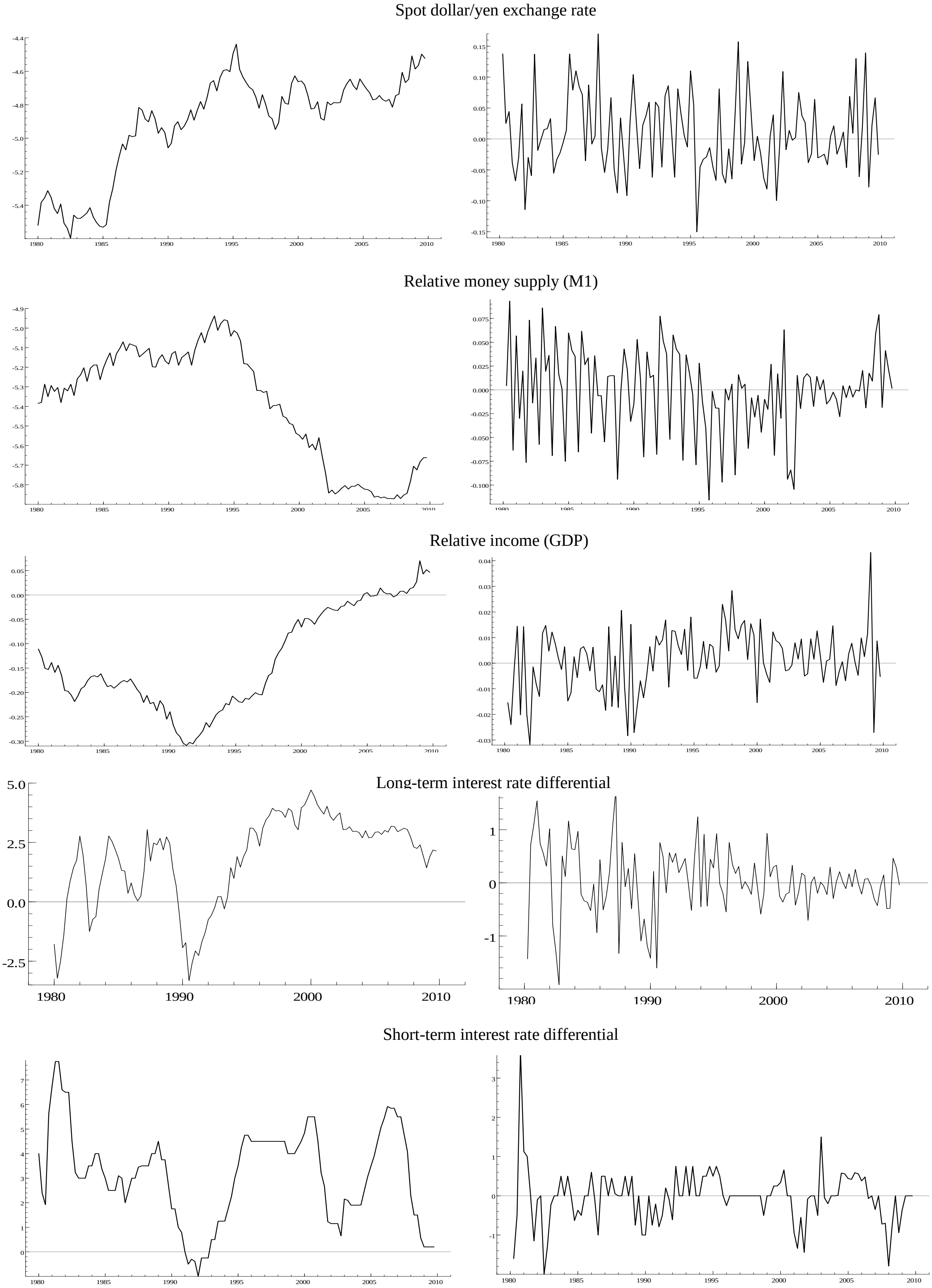

Figure B2.1. The graphs of the standard monetary models variables in levels (left panel) and first differences (right panel).

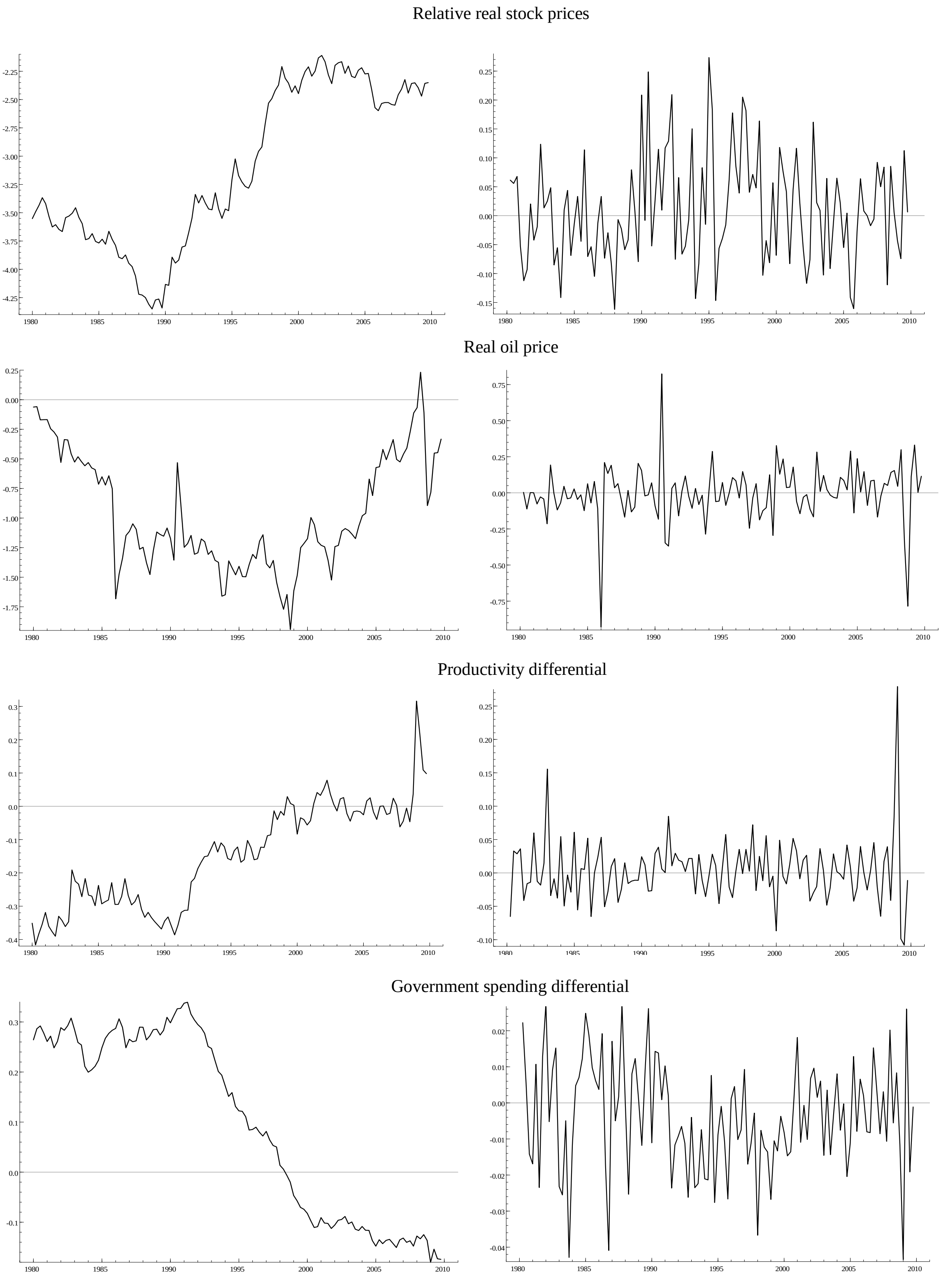

Figure B2.2. The graphs of variables affecting the monetary models' building blocks in levels (left panel) and first differences (right panel).

Appendix C2

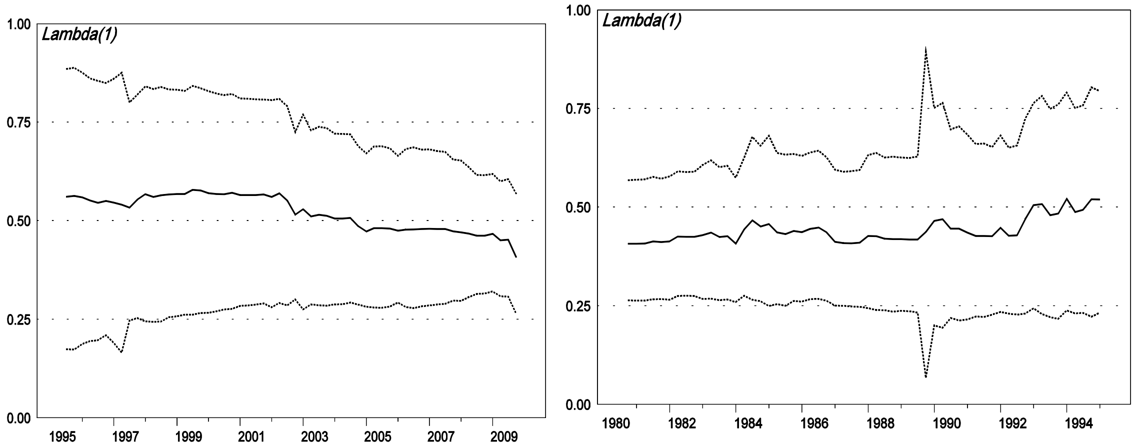

Figure C2.1. Recursively calculated test for the eigenvalues in the modified flexible-price monetary model (1.0 corresponds to 5% critical value).

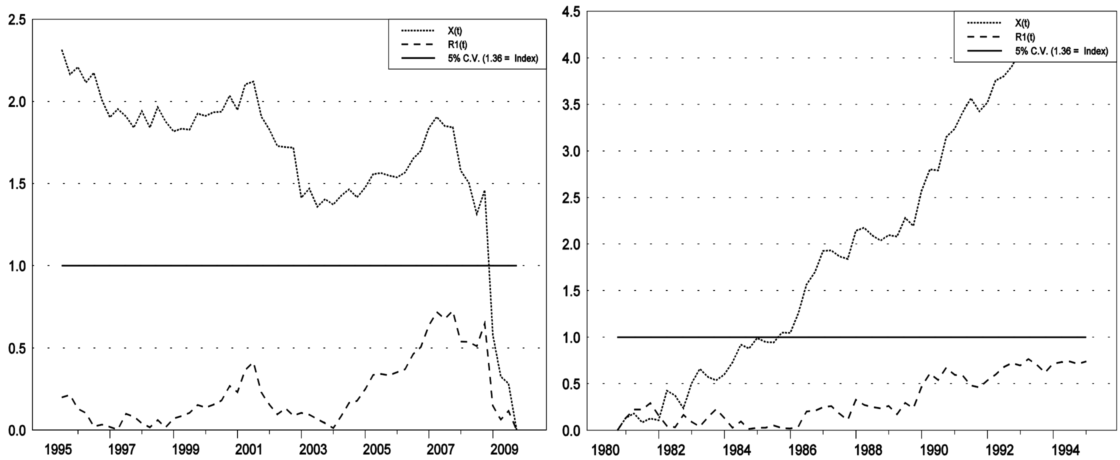

Figure C2.2. Recursively calculated test for the constancy of the log-likelihood in the modified flexible-price monetary model (1.0 corresponds to 5% critical value).

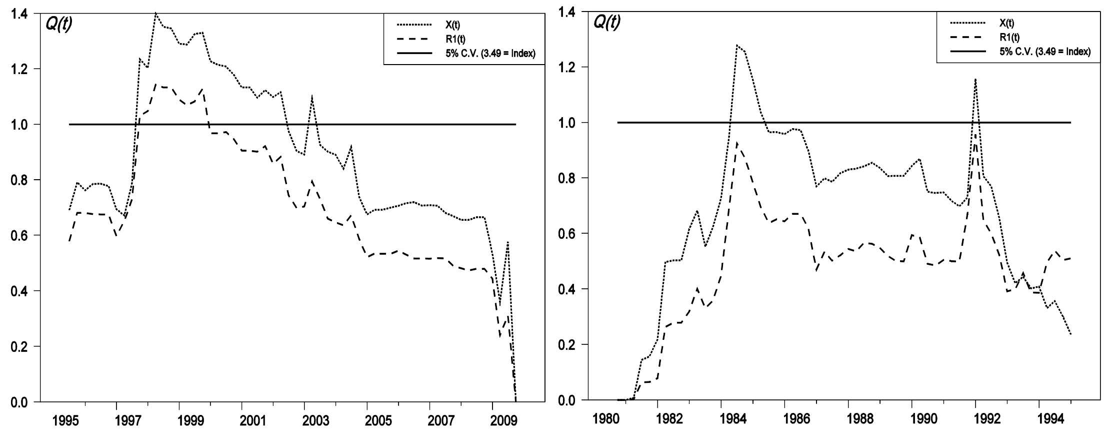

Figure C2.3. Recursively calculated test for the constancy of beta in the modified flexible-price monetary model (1.0 corresponds to 5% critical value).

Appendix D2

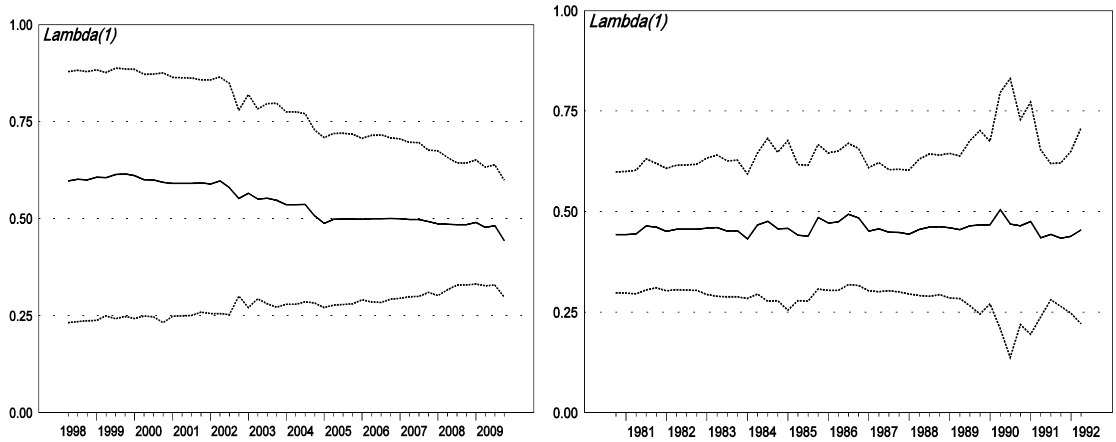

Figure D2.1. Recursively calculated test for the eigenvalues in the modified real interest rate differential monetary model (1.0 corresponds to 5% critical value).

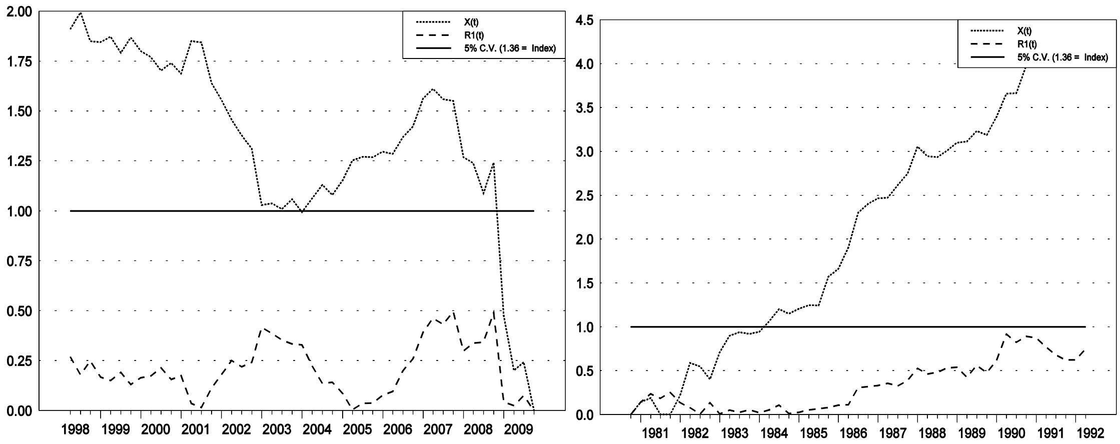

Figure D2.2. Recursively calculated test for the constancy of the log-likelihood in the modified real interest rate differential monetary model (1.0 corresponds to 5% critical value).

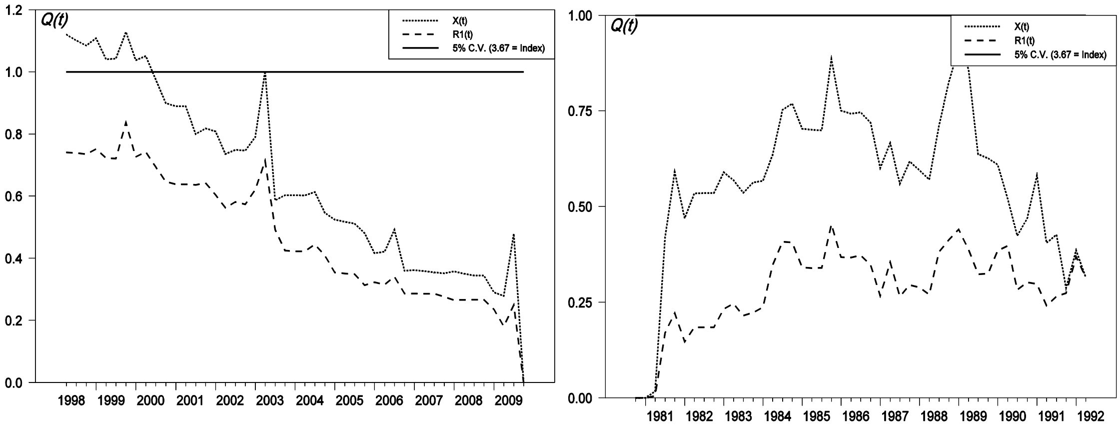

Figure D2.3. Recursively calculated test for the constancy of beta in the modified real interest rate differential monetary model (1.0 corresponds to 5% critical value).

CHAPTER THREE

ON THE LINKAGES BETWEEN STOCK PRICES AND EXCHANGE RATES: EVIDENCE FROM THE BANKING CRISIS OF 2007–2010

3.1. Introduction

The collapse on September 15th 2008 of Lehman Brothers (LB, until that point the fourth largest investment bank in the US) sent a wave of global panic across financial markets. Following global bank failures and the resulting collapse in liquidity and inter-bank lending, stock market indices in most developed economies experienced significant declines. Higher uncertainty also generated turbulence in the foreign exchange markets, with the major currencies being hit by a reduction in international transactions and a flight to value. An interesting issue is whether financial markets have become more dependent as a result of the uncertainty created by the crisis. Aloui et al. (2011), Kenourgios et al. (2011), Samarakoon (2011), Dufrénot et al. (2011), Dimitriou et al. (2013), and Kotkatvuori-Ornberg et al. (2013) among others find indeed an increase in dependence between international stock markets, and similar findings are reported by Coudert et al. (2011) and Bubák et al. (2011) among others for foreign exchange markets.

Surprisingly, the linkages between stock market prices and exchange rates during the recent financial crisis have drawn less attention. To the best of knowledge, the studies by Wong and Li (2010), Tsai (2012), Tsagkanos and Siriopoulos (2013), and Chkili and Nguyen (2014) are the only one to date to have examined the interactions between stock prices and exchange rates during the recent crisis; however, they have some limitations. In particular, Wong and Li (2010) and Tsai (2012) use monthly data which cannot capture the timing of events such as the bailouts of AIG in the US and RBS and HBOS in the UK. Also, their analysis ends in 2008 and 2009, respectively, thereby ignoring the turbulent period following the collapse of LB and the European sovereign debt crisis. While Tsagkanos and Siriopoulos (2013) and Chkili and Nguyen (2014) use higher frequency data and longer sample periods to cover the recent financial crisis, their short-run dynamics results are characterised by significant deviations from normality and conditional heteroscedasticity (see Engle, 1982) that are not captured by their setup. Studies of Tsai (2012) and Chkili and Nguyen (2014) also do not pay a particular attention to the recent financial crisis and focus on the Asian countries and the BRICS economies.

The present chapter contributes to the existing literature by addressing the interactions between stock returns and exchange rate changes as well as their volatilities in a comprehensive manner by analysing weekly data for six advanced economies, namely the US, the UK, Canada, Japan, the euro area and Switzerland. That is, the linkages between the two financial returns and their volatilities are modelled simultaneously. This approach will enable us to capture the time-varying volatility associated with financial data. Also, Ross (1989) pointed out that volatility is a measure of information flow, hence analysing returns and volatilities of stock returns and exchange rate changes simultaneously will enable us to examine not only which type of financial returns predict the mean of the others, but also the transmission of information between the two financial markets.

More specifically, the chapter examines two sub-periods: the pre-crisis (August 6, 2003-August 8, 2007) and the crisis period (August 15, 2007-December 28, 2011). These are selected to enable us to analyse linkages in both normal and turbulent times, which can provide important insights to investors in terms of portfolio management strategies by focusing their attention on the right segments of the markets during such times with the aim of minimising risk and maximising returns in highly integrated financial markets.

The chosen econometric framework is a bivariate VAR-GARCH model in the BEKK representation of Engle and Kroner (1995). Unlike the DCC model which estimates the time-varying conditional correlations directly, the BEKK specification allows for interactions in the variances and covariances in a lead-lag framework. The 'curse of dimensionality' highlighted by Caporin and McAleer (2012) associated with it is not a serious issue in our application with only two variables. Furthermore, to circumvent potential missing variable errors in the conditional mean, the model is extended to incorporate the underlying short-run deviations between stock prices and exchange rates in the conditional mean in case both variables are cointegrated. Therefore, a thorough econometric analysis is conducted of the dependence between stock prices and exchange rates during the period under examination.

The chapter is organised as follows. Section 3.2 provides a brief background of the dynamics of exchange rates and stock prices during the recent financial crisis. Section 3.3 provides a review of the theoretical and empirical literature on the relationship between stock prices and exchange rates. Section 3.4 describes the data and conducts the preliminary analysis. Section 3.5 outlines the econometric methodology used in the chapter. Section 3.6 discusses the empirical results and Section 3.7 concludes.

3.2. The recent financial crisis and the dynamics of stock prices and exchange rates

The recent financial crisis initiated by the crisis with mortgage-backed securities and the failure of Fannie Mae and Freddie Mac in late 2007 as well as the collapse of

LB in September 2008 triggered unprecedented turbulence in the global financial markets, possibly since the great depression. As a consequence of the collapse in liquidity and inter-bank and cross-border lending, stock market indices in most developed economies experienced severe downturns. From early October 2007 until the second week of March 2009, the S&P 500 (US), FTSE 350 (UK), and Stoxx 50 Euro (euro area) indices declined by approximately 56%, 48%, and 59%, respectively. Similar stock market falls occurred in Switzerland and Japan, which also ended with low points in the second week of March 2009 following peaks on June 1 and July 10, 2007, respectively.[30] With such an evaporated confidence among financial institutions, global capital flows also declined sharply during the crisis; they turned negative after the collapse of LB following a steady increase over the last three decades (see Milesi-Ferretti and Tille, 2011, for a detailed analysis).

Other repercussions of the crisis were the real fall in economic activity. Foreign exchange markets also became turbulent, with the major currencies being hit by significant changes and driven by the flight to value. The British pound and the Canadian dollar depreciated against the currencies of their trading partners by approximately 30% (from September 3, 2007 to January 22, 2009) and 28% (from November 7, 2007 to March 9, 2009), respectively. The US dollar experienced a slight appreciation on the onset of the LB collapse, but then depreciated by approximately 20% from March 7, 2009 to July 26, 2011. By contrast, the Japanese yen and the Swiss franc appreciated steadily (by approximately 38% and 61% until late 2011) against the currencies of their trading partners following the failure of the credit market in early August 2007. These two currencies were seen as safe havens during the crisis, hence what was observed was a flight to security.

[30] The Swiss market index, specifically, declined by approximately 54%, whereas the Japanese Nikkei 225 index dropped by roughly 61% during the period.

It follows that the heterogeneous pattern of the foreign exchange movements was evident during the crisis. Fratzscher (2009) found evidence that countries' financial liabilities, FX reserves, countries' current account positions have been the major factors in the global foreign exchange movements during the crisis period.

Although there is a substantial literature examines the dynamics and the linkages between international financial markets during the recent financial crisis such as across international equity markets (e.g., Aloui et al., 2011, Kenourgios et al., 2011; Dimitriou et al., 2013; and Kotkatvuori-Ornberg et al., 2013; among others) and across foreign exchange markets (e.g., Melvin and Taylor, 2009; Fratzscher, 2009; Bubák et al., 2011; among others), there are very few studies investigating the linkages between stock prices and exchange rates during the period. This study provides a good opportunity to explore how the uncertainty generated by the recent crisis affected the dynamic linkages between the two financial markets, with a particular focus on developed economies. As a result of the heterogeneous strength of the major currencies against each other, as discussed earlier, and the heterogeneous pattern of the global capital flows across countries as pointed out by Milesi-Ferretti and Tille (2011), as well as the role of pull and push factors in driving these flows during the period (see Fratzscher, 2012), it is anticipated that nature of the linkages between stock prices and exchange rates may differ across countries and also during the turbulent period compared to the pre-crisis period.

All in all, at times of financial turmoil, the high volatility of stock markets generates speculative actions by investors and capital flight to value and this may lead to considerable instability in other markets such as foreign exchange markets. This has been shown in the case of the Asian financial crisis (see, e.g., Granger et al., 2000; Caporale et al., 2002) and also for the recent financial crisis (see, e.g., Tsagkanos and Siriopoulos, 2013) when stock markets led the foreign exchange markets. However, in turbulent times, decoupling may also occur: when stock markets experience severe downturns, investors may only focus on markets where their assets can be seen as safe havens irrespective of foreign exchange movements; consequently, there might not be interactions between different markets. In fact, Hatemi-J and Roca (2005) concluded that there were no interactions between the stock markets and exchange rates during the Asian crisis once the empirical distributions of the tests for causality were corrected using bootstrapping, as opposed to the findings of Granger et al. (2000).

Knowledge of the interactions between stock market prices and exchange rates during the recent crisis period and what were the nature and the direction of causation during the period are of paramount interest. The present study seeks to answer these questions in a comprehensive manner by analysing weekly data for six advanced economies, namely the US, the UK, Canada, the euro area, Japan, and Switzerland, over the banking crisis of 2007–2010.

3.3. A review of the literature

There are two main types of theoretical models analysing the linkages between exchange rates and stock prices. The traditional approach based on 'flow-oriented' models (Dornbusch and Fischer, 1980) posits that causality runs from the former to the latter, whereas the portfolio approach based on 'stock-oriented' models (Branson, 1983; Frankel, 1983) suggests the opposite. In the first case a more competitive exchange rate will improve the trade position of an economy and stimulate the real economy through firm profitability and stock market prices.[31] However, domestic firms utilising imported inputs will experience an increase in production costs, leading to a reduction in the

[31] This approach has been given some empirical support in the literature on asset pricing models based on consumption and income (Gregoriou et al., 2009), as well as output (Sousa, 2010).

firms' sales and their earnings, which in turn will lead to a decline in their stock prices. Hence, the impact of exchange rates on stock prices can be either positive or negative.

In the second case, the exchange rate is thought to respond to increases in the demand for financial assets such as bonds and stocks. Hence, a bullish domestic stock market will signal favourable domestic economic prospects, thereby inducing capital inflows and an appreciation of the exchange rate (Kollias et al., 2012). Another channel for this type of causality stems from the demand for money (Gavin, 1989). The increased money demand leads to a higher domestic interest rate which in turn attracts investment in the domestic country. This stimulates foreign investors to reallocate their internationally held portfolios by flying out of the foreign assets and buying the domestic ones simultaneously. Consequently, the domestic country will experience capital inflows and an appreciation of its currency.

If, however, both traditional and portfolio approaches are empirically relevant, a bidirectional relation between the two variables will be found with an arbitrary correlation (Granger et al., 2000).

Furthermore, Hau and Rey (2006) recently developed a theoretical framework for the relationship between equity return differential and exchange rate changes on the basis of portfolio rebalancing motive as follows: higher domestic equity returns relative to the foreign counterpart are associated with domestic currency depreciation under incomplete foreign exchange risk. Their rationalisation of this hypothesis is that if unexpected shock gives rise to domestic equity returns relative to the foreign equity returns, the share of domestic equity increases in an internationally held portfolio. Foreign investors find it favourable to fly out a portion of domestic equity to reduce the exposure of the portfolio to foreign exchange risk. Outflows from the domestic equity market as a result of investors' portfolio rebalancing results in a depreciation of the domestic currency.[32]

The empirical literature on the relationship between stock prices and exchange rates is extensive and also provides mixed results. Early studies used the two-step cointegration procedure of Engle and Granger (1987) and the maximum likelihood technique of Johansen (1995) to examine the time series properties of both stock market prices and exchange rates in the long run. Using monthly data on the US economy for the period 1973-1988, Bahmani-Oskooee and Sohrabian (1992) found that these two variables are not cointegrated, yet there is a bidirectional feedback in the short run. Similar findings were reported by Nieh and Lee (2001), who investigated stock prices and exchange rates for the G-7 countries and found one-day significant linkages in some countries.

By considering nine Asian economies and using the Gregory and Hansen (1996) cointegration technique, Granger et al. (2000) also found no evidence of cointegration between stock prices and exchange rates in all cases. However, the results based on Granger causality tests and impulse responses concluded the importance of the stock market as the leader or the existence of bidirectional causality between the two variables during the Asian flu period. By contrast, Ajayi and Mougoue (1996) provided evidence of cointegration between stock prices and exchange rates in all eight advanced countries under their investigation, namely Canada, Germany, France, Italy, Japan, the UK, the US, and the Netherlands, using daily data from 1985 to 1991. Also, significant feedback interactions between both variables were found in the short-run.

[32] While Hau and Rey (2006) found that the portfolio rebalancing motive is strongly supported for 17 OECD countries, Chaban (2009) and Ferreira Filipe (2012) found that such hypothesis is weak for commodity-exporting countries. Note that the analysis of the portfolio rebalancing motive of Hau and Rey (2006) during the recent financial crisis is left for future research.

Using data from January 1992 to December 2005, Alagidede et al. (2011) also failed to uncover cointegration between stock prices and exchange rates using the Johansen (1995), as well as Saikkonen and Lütkepohl (2000a; 2000b; 2000c) cointegration tests. Using different variants of Granger causality tests, there existed linear Granger causality from exchange rates to stock prices in Canada, Switzerland and the UK, whereas the results of nonlinear causality of Hiemstra and Jones (1994) showed that stock prices lead exchange rates in Japan.

Cointegration may not be detected as a result of model misspecification, and in particular the omission of variables. Phylaktis and Ravazzolo (2005) found that US stock prices were a key channel linking the exchange rates of five Pacific Basin countries to their stock indices. Chortareas et al. (2011) also found that both stock prices and exchange rates are interlinked via oil price in three out of four Middle Eastern countries, namely Egypt, Saudi Arabia and Oman. On the other hand, Ülkü and Demirci (2012) showed that global developed and emerging stock market returns explain a large portion of the permanent comovement between stock and foreign exchange markets for eight European emerging economies. Considering six Asian countries (Singapore, Thailand, Malaysia, the Philippines, South Korea, and Taiwan) from January 1992 to December 2009 and using the quantile regression model, Tsai (2012) found that the negative impact of stock prices on exchange rates, prevailed by the portfolio balance approach, is more evident when exchange rates are extremely high or low.

The seminal article of Engle (1982) showed that the ARCH family of models can capture volatility clustering and ARCH effects in financial returns such as those of stock markets and exchange rates. Kanas (2000) found positive volatility spillover effects from stock returns to exchange rate changes for all the G-7 countries except Germany. The failure to find volatility spillover effects between both variables in the case of Germany was attributed to the intervention of the Bundesbank in the currency markets during the era of the Exchange Rate Mechanism (ERM). Caporale et al. (2002) also found that causality-in-variance runs from stock returns to exchange rate changes over the whole sample (1/1/1987 – 20/1/2000) and for all four East Asian countries (Indonesia, Japan, South Korea and Thailand). In line with Granger et al. (2000), their evidence for the post-Asian crisis period indicated the dominance of the stock market in the flow of information or the existence of a feedback relation in terms of the second moment between the two financial markets.

Ning (2010), instead, used copulas to show that there is significant symmetric upper and lower tail dependence between the stock and foreign exchange markets of the G-5 countries (US, UK, Germany, France and Japan). Katechos (2011) found that the sign of the link between global stock market returns and exchange rate changes depends on whether the currency in question is a high yielding (positive) or a lower yielding one (negative). Chkili et al. (2012), who estimated a bivariate CCC-FIAPARCH specification to capture asymmetry and long memory in daily data from January 1999 to December 2010 for three major European countries (namely France, Germany, and the UK), reported a strong correlation between the two variables and more accurate in-sample estimates, as well as better out-of-sample performance than in the case of GARCH specifications.

More recently, Tsagkanos and Siriopoulos (2013) employed the structural nonparametric cointegrating regression and found the existence of a long run (short run) causal relationship from stock prices to the exchange rates in the EU (US) during the recent financial crisis (2008-2012). Using regime switching VAR models for the BRICS countries (Brazil, Russia, India, China, and South Africa), Chkili and Nguyen (2014) also found that stock markets have more impact on exchange rates during tranquil, as well as turbulent periods using weekly data over the period 1997-2013. Moore and Wang (2014), on the other hand, showed that the dynamic relationship between real exchange rates and stock return differential of the emerging Asian markets in relation to the US is driven by trade balance. However, in the case of developed economies, the interest rate differential was found to be the driving force of the dynamic relationship between the two variables.

The above empirical literature on the linkages between stock market prices and exchange rates has essentially been prompted by the early studies on the sensitivity of the value of firms to foreign exchange exposure (e.g., Aggarwal, 1981; Soenen and Hennigar, 1988; Jorion, 1990).[33] The latter strand of literature is also extensive, with mixed results. For example, Aggarwal (1981) found a significant positive correlation between US stock prices and the strength of the US dollar using monthly data between 1974 and 1978, although Soenen and Hennigar (1988) reported that the sign depends on the sample used.

By examining the exposure of US multinationals to foreign currency risk over the period 1971-1987, Jorion (1990) found that the statistically significant contemporaneous effect of a change in the trade-weighted exchange rate on the value of the firm is existed in only 15 firms out of 287 US multinational firms. However, using a sample of 208 firms for the period between 1978 and 1990, Bartov and Bodnar (1994) found that no contemporaneous but a lagged change in the value of the dollar has a significant influence on the abnormal returns of these firms.

Griffin and Stulz (2001) also observed, by employing weekly data on industry indices from the US, the UK, Canada, France, Germany and Japan over the period 1975-1997, that the influence of exchange rate shocks on the value of industries is negligible. Nonetheless, using a sample of 171 Japanese multinationals for the period 1979-1993, He and Ng (1998) found that about 25% of these firms exhibited

[33] See Muller and Verschoor (2006b) for a detailed theoretical and empirical review.

economically significant positive contemporaneous exposure to exchange rate changes. Considering a sample of 817 European multinational firms, Muller and Vershoor (2006a), by contrast, reported that 14% of the firms exhibited economically significant exposure effects to the US dollar, 13% to the Japanese yen and 22% to the UK pound. Using a large sample of non-financial firms from 37 Countries, Bartram and Bodnar (2012) found the existence of noticeable differences in the impact of exchange rates on the returns of firms across the considered countries, with 30–40% of firms in open and emerging market countries being found significantly exposed to foreign exchange rate risk.

Furthermore, by using a sample of automotive firms from the US and Japan, Williamson (2001) confirmed the existence of time-varying foreign exchange rate exposure across countries for multinational firms and global competitors. The paper argued that the time-varying exposure is due to changes in the structure of the industry and its competition through time. Using a dynamic framework based on vector GARCH, Koutmos and Martin (2007) also found that the exchange rate exposure of US stocks is time varying. The average time-varying exposure was found to be statistically significant for the size-based, as well as sector-based portfolios.

Aggarwal and Harper (2010) examined US domestic firms and found that these firms exhibit significant foreign exchange exposure, which is, on average, not significantly different from the exposures of the corresponding multinational firms. Agyei-Ampomah et al. (2013) found that the sensitivity of foreign exchange exposure is model-dependent. Using a sample of 269 UK non-financial firms from January 1991 to December 2010, the paper found that Jorion's (1990) model implies that 14.93% (30.50%) of the firms are exposed, directly or indirectly, to the UK pound (US dollar, euro, or Japanese yen). However, the exposure increases substantially to 85.13% (96.65%) when the orthogonalised GARCH-based two-factor asset pricing model with time varying coefficients is adopted. The more recent study by Tsai et al. (2014) report, by examining various Taiwanese industries over the period 2001-2010, that stock returns show less sensitivity to foreign exchange exposure when the effect of hot money on stock returns is taken into account.

3.4. Data description and preliminary analysis

Weekly data (Wednesday to Wednesday) are employed in order to analyse the linkages between stock market prices and exchange rates because daily or intra-daily data are affected by the synchronicity of trading between the various markets. Also, daily or intra-daily data are affected by noise and anomalies such as day-of-the-week effects, while monthly data may be inadequate to trace the short-run evolution of capital across international financial markets. We consider six advanced economies: Canada, the euro area, Japan, Switzerland, the UK, and the US from August 6, 2003 to December 28, 2011, a sample of 441 observations. The exchange rates used are trade-weighted (as calculated by the Bank of England), thus providing a better measure of the competitiveness of these economies (Kanas, 2000), while the stock prices are the main local stock exchange indices. The currencies of these economies are the most actively traded in the foreign exchange markets, while their stock markets are the largest among the developed economies in terms of market capitalisation. The data have been obtained from Thomson DataStream.

We consider two sub-periods: a tranquil or pre-crisis period from August 6, 2003 to August 8, 2007, and a crisis period from August 15, 2007 to December 28, 2011. It is well known that the former corresponds to the so-called 'Great Moderation' (see Stock and Watson, 2002), which was characterised by stable and low inflation and a decline in the volatility of other macroeconomic fundamentals. The subsequent global financial crisis (and the associated 'Great Recession') clearly represents a new regime.

The start date of the pre-crisis sample is chosen to avoid the impact of major global events such as the 9/11 terrorist attacks and their anniversary in 2002 (see Gregoriou et al., 2009), and the ensuing conflicts in Afghanistan and Iraq as well as the dotcom bubble that burst in late 2002. On the other hand, the crisis period is defined as starting with the first signs of the subprime mortgage crisis in the US in the summer of 2007, ahead of the failure of Fannie Mae and Freddie Mac and the collapse of LB and AIG. This is also consistent with the study of Melvin and Taylor (2009), who consider August 16th 2007 as the beginning of the crisis in the foreign exchange markets.

The variables in levels are denoted by s_t and e_t, respectively the log stock prices and log exchange rates, while their first differences ($R_{S,t}$ and $R_{E,t}$) are continuously compounded returns; the data are in percentages and are multiplied by 100. To evaluate the stochastic features of stock market returns and exchange rate changes, a wide range of descriptive statistics is displayed in Table A3.1 (Appendix A3) where Panel A and B indicate respectively the two sub-periods under observation, pre-crisis and crisis periods.

The mean weekly changes in exchange rates are positive (appreciation) for the UK, the euro area, and Canada, whereas they are negative (depreciation) for the US, Switzerland and Japan during the pre-crisis period. During the crisis period, the mean changes for the US dollar, the euro, and the British pound are negative (depreciation), while they are positive (appreciation) for the rest of the other currencies. The averages of weekly returns for stock indices, on the other hand, are positive during the pre-crisis period; however, the crisis period indicates the reverse for all cases. Thus, this implies the severe downturns being experienced in stock market indices ever since the financial crisis compared with the preceding period.

With regard to volatility measures, two remarks are in order. First, stock market returns exhibit higher volatility than exchange rate changes in both sub-samples.

Second, both stock returns and exchange rate changes, as expected, exhibit higher volatility during the crisis period compared with the pre-crisis one in all countries. In terms of the third and fourth moments, stock returns exhibit excess kurtosis and are negatively skewed during both sub-periods. Exchange rate changes also exhibit excess kurtosis and skewness. In the pre-crisis period, such changes are negatively skewed for Canada, and the UK, and positively skewed for the rest of the other countries. In the crisis period, they are positively skewed for the euro, the British pound and the Japanese yen, whereas for the rest of the other currencies they are negatively skewed.

Overall, the Jarque-Bera (JB) test statistics indicate a rejection of the null hypothesis that stock returns are normally distributed in all countries in both sub-periods. Although exchange rate changes indicate the existence of normality for the pre-crisis period for all countries except the euro area and Switzerland; however, the crisis period indicates the failure of such normality for all countries. With regard to the Ljung and Box (1978) *Q*-statistics of the return series and their squares calculated up to 10 lags, the corresponding statistics, broadly speaking, indicate the existence of significant linear and non-linear dependencies in the data, especially in the crisis sample.

Figure 3.1a and 3.1b show respectively the weekly evolution of the trade-weighted exchange rates and stock prices with their corresponding changes for the period under investigation. While the trade-weighted exchange rates of the UK and Switzerland did not experience substantial fluctuations in the pre-crisis period, the corresponding rates of Canada and the euro area appreciated steadily. The US dollar and the Japanese yen, by contrast, experienced depreciation in the pre-crisis period, with the depreciation of the latter being stronger. In the crisis period, the Swiss franc and the Japanese yen appreciated steadily, whereas the US dollar, the euro, the British pound, and the Canadian dollar depreciated over the period, with the latter turned to appreciate after 2009. While the Japanese yen continues its appreciation till the end of the crisis-period, the Swiss National Bank in the last quarter of 2011 intervened in the market for Swiss francs in order to ease the impact of its overvaluation on exports and the inflationary effect of this on the economy.[34] With regard to stock markets, although all indices rose in the pre-crisis period, they declined sharply following the onset of the crisis, as can be seen from the figures.

Stock returns and exchange rate changes exhibit volatility clustering, especially in the crisis period, which indicates an ARCH model might be appropriate for analysing the linkages between the two variables. The Figures also suggest that the log of exchange rates and stock prices might be non-stationary and follow a stochastic trend, while their first difference is co-variance stationary or has a finite variance. This is confirmed by a battery of unit root tests, including the augmented Dickey–Fuller (1981) test, the Phillips and Perron (1988) test, and the minimum LM test of Lee and Strazicich (2004) with one structural break in the intercept and the trend, displayed in Tables A3.2, A3.3 and A3.4, respectively (see Appendix A3). The latter test is advantageous to other alternatives with a single endogenous structural break such as Zivot and Andrews (1992) test.[35] This is because it is characterised by no size distortion and spurious rejections in the presence of a break under the null, and hence such a test circumvents the potential erroneous conclusions associated with the Zivot and Andrews (1992) test.

[34] The intervention was specifically conducted by reducing interest rates and setting a floor for the franc against the euro at a rate of 1.20 franc per euro in August 2011 and September 2011, respectively.

[35] The endogenous breakpoint in Zivot and Andrews (1992) test is chosen where a one-sided test statistic on the coefficient in the ADF test is minimised (i.e., the most negative). Hence, such a test favours to reject the null of unit root for a trend-stationary process with a break.

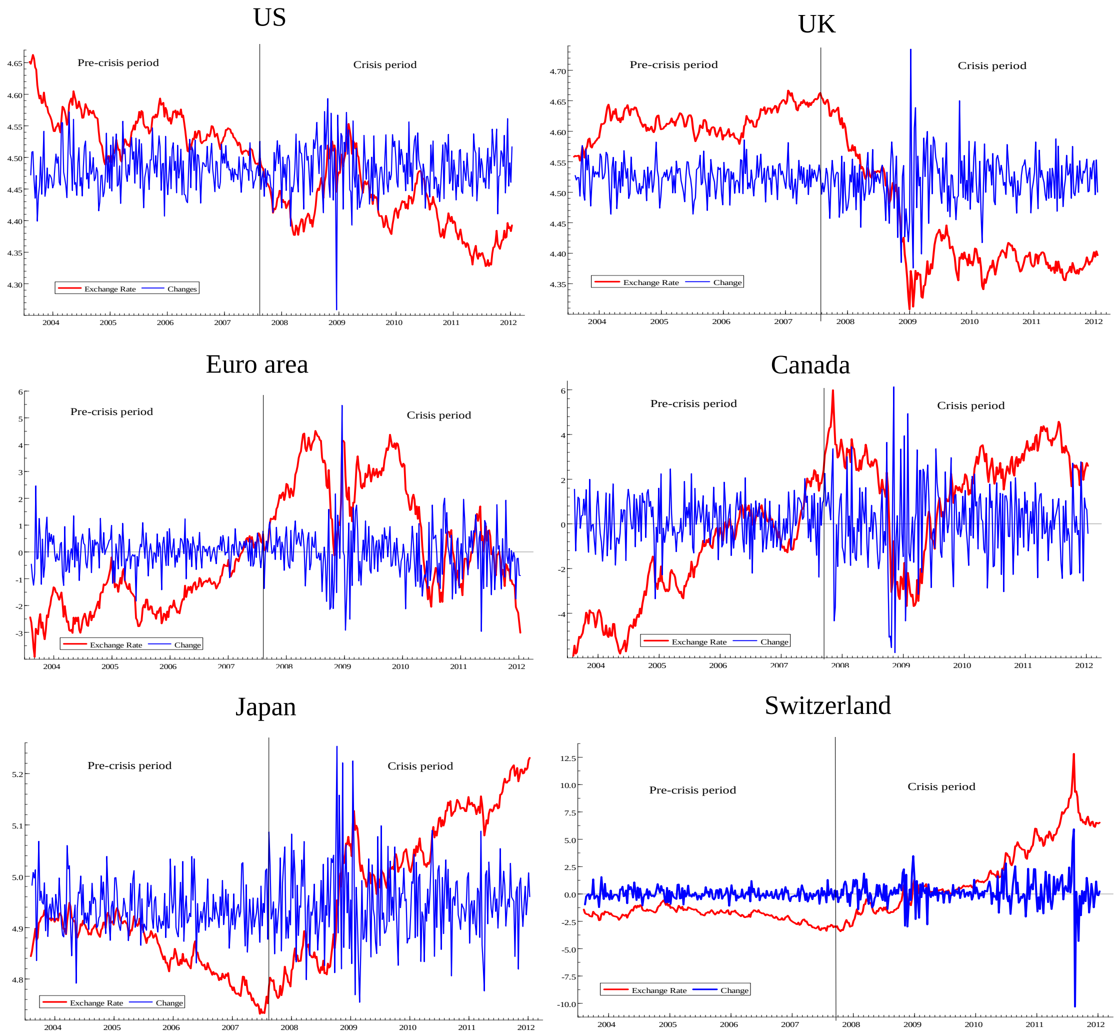

Figure 3.1a. Weekly trade-weighted exchange rates with their corresponding changes over the period August 6, 2003-December 28, 2011.

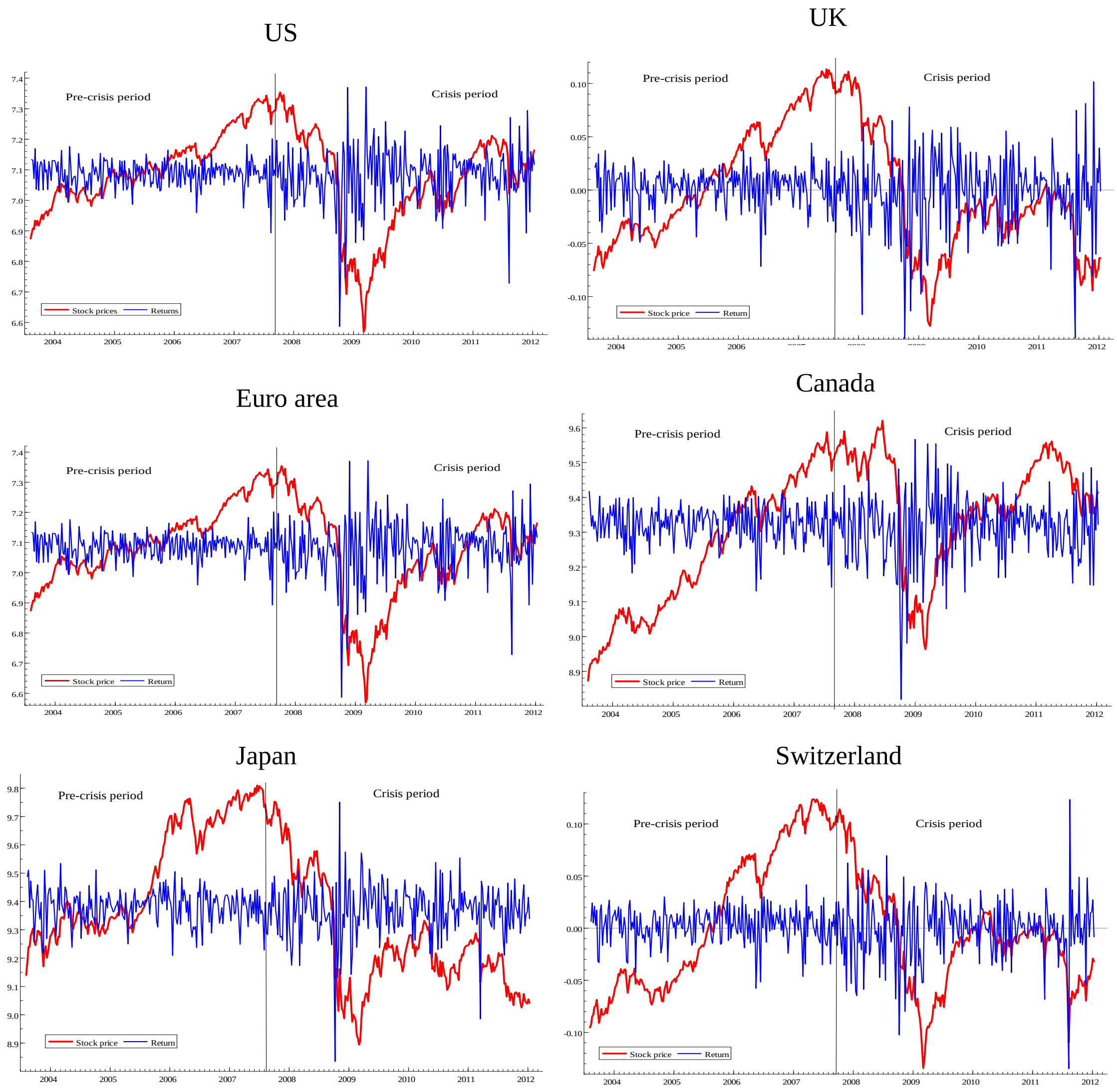

Figure 3.1b. Weekly stock market prices with their corresponding returns over the period August 6, 2003-December 28, 2011.

Employing a test which allows for a structural break is likely to be informative. The crisis period is characterised by many incidents (i.e., Lehman Brothers collapse, European debt crisis, downgrade of US debt status, etc.) which might influence the time path of both stock market prices and exchange rates. Earlier, Perron (1989)[36] provided evidence that not considering an actual structural change in a time series results in inefficient conclusions by not rejecting a false unit root null. Thus, the test in question can potentially establish findings associated with more accurately specified models (Strazicich et al., 2004). Furthermore, the identified breakpoints by LM tests in Table A3.4 (see Appendix A3) are found to be significant in most cases.

3.5. The econometric model

3.5.1. The VAR-GARCH model

We employ the BEKK representation of Engle and Kroner (1995) for our bivariate VAR-GARCH (1, 1) model to examine the joint processes governing weekly changes in stock market prices and exchange rates for the two sub-periods. This enables us to examine the dependence between both the first and the second moments of stock returns and exchange rate changes in a dynamic framework. Furthermore, the model also includes some exogenous variables to capture the effects of domestic monetary policy shocks as well as global shocks such as those of world stock market and oil prices. That is, the conditional mean equation is specified as follows:

$$\begin{gathered} R_t = \mu + \sum_{i=1}^{p} \psi_i R_{t-i} + \lambda R_{w,t} + \gamma R_{rf,t} + \delta \Delta P_{oil,t} + \varepsilon_t \\ \mu = \begin{bmatrix} \mu_S \\ \mu_E \end{bmatrix}, \psi_i = \begin{bmatrix} \psi_{SS}^{(i)} & \psi_{SE}^{(i)} \\ \psi_{ES}^{(i)} & \psi_{EE}^{(i)} \end{bmatrix}, \varepsilon_t = \begin{bmatrix} \varepsilon_{S,t} \\ \varepsilon_{E,t} \end{bmatrix} \end{gathered} \tag{3.1}$$

[36] The procedure proposed by Perron (1989) allows for a known or exogenous structural break.

where $R_t = [R_{S,t}, R_{E,t}]$, the innovation vector $\varepsilon_t \mid \Omega_{t-1} \sim \mathrm{N}(0, \mathrm{H}_t)$ is normally distributed with H_t being the corresponding variance-covariance matrix, and Ω_{t-1} is the information set available at time t-1. $\psi_{SS}^{(i)}$ and $\psi_{EE}^{(i)}$ indicate respectively the response of stock market returns and exchange rate changes to their own lags, whereas $\psi_{ES}^{(i)}$ and $\psi_{SE}^{(i)}$ measure respectively causality from stock market returns to exchange rate changes, and vice versa (*i* denotes the lagged time-period).

The model is augmented with some exogenous variables, namely $R_{w,t}$ (returns of the world stock index), $R_{rf,t}$ (the three-month domestic interest rate), and $p_{oil,t}$ (the logarithm of the world oil price).[37] Returns on the world stock market capture shocks from other financial markets around the globe; for example Caporale and Spagnolo (2011) used US stock market returns to proxy for market globalisation when they examined stock market integration between Central and Eastern European countries and both the UK and Russia. Interest rates reflect domestic monetary policy (i.e., quantitative easing policies in the crisis period, etc.) and the availability of credit, given that monetary authorities of the economies under consideration are using interest rate

[37] The interest rates are 3-month treasury bills (for the US, the UK, and Canada), 3-month certificate of deposit (for Japan), 3-month Swiss interbank rate (for Switzerland), and 3-month euribor rate (for the euro area). Returns on the world stock index for all countries in the sample except the US are represented by returns on the S&P 500 index. In the case of the US, the world stock index is represented by the MSCI world (excluding the US) index. Due to the nonsynchronous trading time of the US stock market with other markets under observation; that is, the US stock market closes after the stock markets of other countries under observation, we include (Tuesday to Tuesday) returns of US' S&P 500 index for all countries, except Canada. In the case of Canada, we use (Wednesday to Wednesday) returns due to the contemporaneous trading time between the US and Canada. The world oil price is represented by the West Texas Intermediate Cushing crude oil spot index, US dollars per barrel. The data have been obtained from DataStream.

rules responding to inflation and an output gap.[38] The oil price can be seen as representing supply shocks; for example Amano and van Norden (1998) showed that world oil price movements can capture the underlying shocks to the terms of trade. See also Lizardo and Mollick (2010) for the impact of the real oil price on the US dollar exchange rate and Filis et al. (2011) for the effect of the real oil price on stock market returns.

Given the nature of the data, in most cases a lag length *p=1* is sufficient to capture the dynamics associated with financial returns. If necessary, further lags are added to eliminate any serial correlation on the basis of the multivariate *Q*-statistic of Hosking (1981) applied to the standardised residuals, $z_{i,t} = \varepsilon_{i,t} / \sqrt{h_{i,t}}$ for $i = S, E$.

Note that in the event of detecting cointegration between stock market prices and exchange rates, Eq. (3.1) is also augmented by a lagged error correction term (ect_{t-1})[39], as in Li et al. (2001). The exclusion of an error correction term in the differenced VAR gives rise to a vector moving average term that is generally non-invertible (see Burke and Hunter, 2005). Cointegration is tested using the Engle and Granger (1987) two-step procedure, the Johansen (1995) trace test, and the Gregory and Hansen (1996) method that allows for a single unknown endogenous structural break; see subsection 3.5.2 for a detailed summary of these tests. Furthermore, if a structural shift in the long-run relationship between stock prices and exchange rates is detected and in order to examine the impact of this shift on the dynamic linkages between the two variables and the short-run adjustment towards the long-run equilibrium, we use a dummy variable and allow

[38] Recently, Laopodis (2013) showed that the dynamic relationship between the monetary policy and the stock market is monetary regime-dependent. Hnatkovska et al. (2013), on the other hand, found that the relationship between interest rates and the nominal exchange rate is non-monotonic; larger increases in the nominal interest rate depreciate the currency, whereas small increases appreciate it.

[39] The error correction term is measured from a cointegrating relation in a similar manner to Engle and Granger (1987), as in Eq. (3.7).

the parameters related to Granger causality between both variables, denoted by $\psi_{ES}^{(i)*}$ and $\psi_{SE}^{(i)*}$, as well as the error correction term, denoted by η^{*}, to shift as follows:

$$R_t = \mu + \sum_{i=1}^{p} \psi_i R_{t-i} + \lambda R_{w,t} + \gamma R_{rf,t} + \delta \Delta P_{oil,t} + \eta\, ect_{t-1} + \eta^{*} ect_{t-1} + \varepsilon_t,$$
$$\psi_i = \begin{bmatrix} \psi_{SS}^{(i)} & \psi_{SE}^{(i)} + \psi_{SE}^{(i)*} \\ \psi_{ES}^{(i)} + \psi_{ES}^{(i)*} & \psi_{EE}^{(i)} \end{bmatrix}. \tag{3.2}$$

Finally, as Granger causality tests do not provide information on the signs and timing of the relation between stock returns and exchange rate changes, we compute the generalised impulse response functions (GIRFs) of Pesaran and Shin (1998) based on Eq. 3.1 and Eq. 3.2 in order to explore the direction of spillovers between the two variables and their evolution over time.

Having specified the conditional mean equation, a differenced VAR (Sims, 1980) is estimated in the case of no cointegration between the two financial assets, whereas a vector error correction form is adopted (Johansen, 1995) when the variables are cointegrated. The model is then estimated conditional on the multivariate GARCH model in the BEKK specification.[40] Because of its quadratic forms, the estimated conditional variance-covariance matrices in the BEKK model are ensured to be positive definite. The conditional variance-covariance equation can be expressed as follows:

$$H_t = C'C + A'\varepsilon_{t-1}\varepsilon'_{t-1}A + B'H_{t-1}B. \tag{3.3}$$

[40] The published paper drawn from this chapter uses the UEDCC-GARCH model instead of the BEKK specification to estimate the conditional variance-covariance equation (see Caporale et al., 2014). However, the conclusion of the paper is relatively the same as that of this chapter.

More explicitly:

$$\begin{bmatrix} H_{SS,t} & H_{SE,t} \\ H_{ES,t} & H_{EE,t} \end{bmatrix} = C'C + A' \begin{bmatrix} \varepsilon_{S,t-1}^2 & \varepsilon_{S,t-1}\varepsilon_{E,t-1} \\ \varepsilon_{E,t-1}\varepsilon_{S,t-1} & \varepsilon_{E,t-1}^2 \end{bmatrix} A + B' \begin{bmatrix} H_{SS,t-1} & H_{SE,t-1} \\ H_{ES,t-1} & H_{EE,t-1} \end{bmatrix} B. \tag{3.4}$$

Where C is a lower triangular matrix, and A and B are ARCH and GARCH parameter matrices:

$$C = \begin{bmatrix} c_{SS} & 0 \\ c_{ES} & c_{EE} \end{bmatrix}, A = \begin{bmatrix} a_{SS} & a_{SE} \\ a_{ES} & a_{EE} \end{bmatrix}, B = \begin{bmatrix} b_{SS} & b_{SE} \\ b_{ES} & b_{EE} \end{bmatrix}.$$

It follows from Eq. (3.4) that in the BEKK specification each conditional variance and covariance in H_t is modelled as a linear function of lagged conditional variances and covariances, and lagged squared innovations and the cross-product of the innovations.

Note that the variance-covariance matrix is not extended to take into account asymmetric responses as sign and size bias tests (as in Engle and Ng, 1993), displayed in Table 3.1, show no evidence of asymmetry for the two variables. This applies to all cases except the following. Stock returns in Japan in the pre-crisis period show that the joint impact of the three effects (sign, negative and positive size bias) is significant, US stock returns in the pre-crisis period exhibit weak sign bias (significant at the 10% level), exchange rate changes in Switzerland in the crisis period exhibit positive size bias and also the joint impact of the three effects is significant, and finally exchange rate changes in Japan show significant sign bias in the crisis period.

Volatility is transmitted between stock returns and exchange rate changes through two channels represented by the off-diagonal parameters in the ARCH and

GARCH matrices: a symmetric shock $\varepsilon_{ii,t-1}$ and the conditional variance $H_{ii,t-1}$. Specifically, volatility transmission from stock returns to exchange rate changes is tested by setting $a_{SE} = b_{SE} = 0$, and in the reverse direction by $a_{ES} = b_{ES} = 0$. Using Monte Carlo simulation technique, Hafner and Hewartz (2008) showed that these causality-in-variance tests in the context of multivariate GARCH-BEKK models have superior power to the cross-correlation function (CCF) two-step approach proposed by Cheung and Ng (1996). Causality-in-variance is tested using a likelihood ratio test statistic:

$$LR = -2(L_r - L_{ur}) \sim \chi^2_{df} \tag{3.5}$$

where L_r and L_{ur} indicate the restricted and unrestricted log-likelihood function values, respectively; LR follows the chi-squared distribution with the degrees of freedom equal to the number of restricted parameters (df).

When the innovations are assumed to be normally distributed, the log-likelihood function is given by:

$$L(\theta) = \frac{-Tn}{2}\ln(2\pi) - \frac{1}{2}\sum_{t=1}^{t}(\ln|H_t| + \varepsilon'_t H^{-1}{}_t \varepsilon_t) \tag{3.6}$$

where $n = 2$, T is 209 and 228 respectively for the pre-crisis and crisis periods, and θ is a vector of unknown parameters. Specifically, the quasi-maximum likelihood estimator of Bollerslev and Wooldridge (1992) is applied as the corresponding computed standard

Table 3.1. Sign and size bias tests of Engle and Ng (1993).

Panel A. Pre-crisis period, 6 August 2003- 8 August 2007

		Sign	Neg. Size	Pos. Size	Joint
Canada	$R_{E,t}$	0.197[0.843]	0.152[0.878]	0.123[0.901]	0.043[0.144]
	$R_{S,t}$	0.197[0.843]	0.132[0.894]	0.559[0.575]	0.359[0.948]
Euro area	$R_{E,t}$	0.987[0.323]	0.163[0.870]	0.088[0.929]	2.702[0.439]
	$R_{S,t}$	0.830[0.406]	0.561[0.574]	0.873[0.382]	3.797[0.284]
Japan	$R_{E,t}$	0.147[0.882]	1.190[0.233]	0.063[0.949]	2.116[0.548]
	$R_{S,t}$	1.475[0.140]	0.124[0.900]	0.591[0.554]	8.514[0.036]
Switzerland	$R_{E,t}$	1.527[0.126]	0.233[0.815]	1.465[0.142]	2.927[0.402]
	$R_{S,t}$	0.634[0.525]	0.118[0.905]	0.018[0.985]	1.261[0.738]
UK	$R_{E,t}$	0.668[0.504]	0.646[0.517]	1.208[0.226]	2.263[0.519]
	$R_{S,t}$	0.988[0.322]	0.573[0.566]	1.104[0.269]	1.583[0.663]
US	$R_{E,t}$	0.398[0.690]	0.299[0.764]	0.433[0.664]	0.879[0.830]
	$R_{S,t}$	1.703[0.088]	0.170[0.864]	0.809[0.418]	5.668[0.128]

Panel B. Crisis period, 15 August 2007- 28 December 2011

Canada	$R_{E,t}$	0.803[0.421]	0.092[0.926]	0.371[0.710]	0.905[0.824]
	$R_{S,t}$	0.533[0.593]	1.167[0.242]	0.292[0.770]	5.376[0.146]
Euro area	$R_{E,t}$	0.790[0.429]	0.203[0.838]	0.359[0.719]	2.847[0.415]
	$R_{S,t}$	1.206[0.227]	0.211[0.832]	0.321[0.747]	5.296[0.151]
Japan	$R_{E,t}$	2.041[0.041]	1.351[0.176]	1.062[0.287]	4.324[0.228]
	$R_{S,t}$	0.733[0.463]	0.595[0.551]	0.937[0.348]	1.270[0.736]
Switzerland	$R_{E,t}$	0.073[0.941]	1.010[0.312]	2.620[0.008]	12.79[0.005]
	$R_{S,t}$	0.967[0.333]	0.231[0.816]	0.242[0.808]	3.258[0.353]
UK	$R_{E,t}$	0.790[0.429]	0.203[0.838]	0.359[0.719]	2.847[0.415]
	$R_{S,t}$	0.241[0.809]	0.706[0.480]	1.525[0.127]	5.158[0.160]
US	$R_{E,t}$	0.204[0.837]	0.205[0.837]	1.147[0.251]	1.846[0.604]
	$R_{S,t}$	0.106[0.914]	0.007[0.993]	1.394[0.163]	3.629[0.304]

Notes: The tests are conducted on the residuals from a univariate GARCH model for each series subject to sufficient lags in the mean and the variance to remove serial correlation in the residuals as well as the squared residuals. *P*-values are reported in square brackets [.].

errors are robust even when the error process is non-normal.[41] We also employ the multivariate Q-statistic (Hosking, 1981) for the squared standardised residuals to determine the adequacy of the estimated model of the conditional variance-covariance matrix in capturing the ARCH and GARCH dynamics.

3.5.2. Cointegrations tests

Ever since the seminal work of Engle and Granger (1987) and the subsequent development of Johansen (1988; 1995), cointegration has been the cornerstone technique to examine the features of non-stationary time series in the literature of economics and finance. Cointegration, loosely speaking, implies that a linear combination of nonstationary time series integrated of the same order is stationary where the economic sense of such a concept refers to the presence of a long-run relationship or long-run predictability between the series in question. In this chapter, we employ different cointegration techniques to investigate the long-run relationships between stock market prices and exchange rates within the national economies in the two sub-periods.

The first test is the pairwise Engle and Granger (1987) two-step procedure which is based on the following cointegrating regression:

$$s_t = \varkappa + \pi e_t + u_t, \tag{3.7}$$

where s_t and e_t denote, as stated earlier, respectively the log stock market price and the log exchange rate, which were found to be $I(1)$ series. If the estimated residual of the

[41] The procedure was implemented with a convergence criterion of 0.00001 in RATS 8.1, using the quasi-Newton method of Broyden, Fletcher, Goldfarb, and Shanno, which does not require exact estimates of the matrix of second derivatives in contrast to the approach of Berndt, Hall, Hall, and Hausman (see Sargan, 1988).

regression u_t is found to be stationary (i.e., $u_t \sim I(0)$) using the ADF unit root test, as shown below in Eq. (3.8), then cointegration is existed between s_t and e_t, otherwise not:

$$\Delta u_t = \partial + (\delta - 1)\, u_{t-1} + \sum_{i=1}^{p} \emptyset_i \Delta u_{t-i} + \vartheta_t, \tag{3.8}$$

where Δ is the first difference operator; ∂ is a constant; ϑ_t is a white noise term; and p is the number of lags. The test is conducted under the null hypothesis H_0: $\delta = 1$ (no cointegration) where the estimated test statistic is compared with the MacKinnon (1991) critical values. The optimal lag length in the ADF test is chosen by considering the Akaike information criterion (AIC).

The second test we use is the Johansen (1995) maximum likelihood cointegration technique. Johansen (1995) specifies an unrestricted Vector Autoregressive (VAR) model of order k with $(n \times 1)$ endogenous variables integrated of the same order (i.e., $I(1)$) forced by a vector of $(n \times 1)$ independent Gaussian errors. That is, it is expressed in an error correction form as:

$$\Delta Y_t = \Pi Y_{t-1} + \Gamma_1 \Delta Y_{t-1} + \ldots\ldots + \Gamma_k \Delta Y_{t-(k-1)} + \varepsilon_t, \tag{3.9}$$

where Y is an $(n \times 1)$ vector of $I(1)$ variables in question (log stock price and log exchange rate); Γ_i (i =1,..., k) are $(n \times n)$ parameter matrices capturing the short-run dynamics among the variables, and finally Π is an $(n \times n)$ matrix which is partitioned as $\Pi = \alpha\beta'$ where α and β matrices encompass the speed of adjustment and long-run parameters, respectively.

Johansen proposed two likelihood ratio tests which represent the key statistics for testing for cointegration, and hence determining the rank r of the long-run matrix Π. The tests are the trace test and the maximum eigenvalue statistics. However, as shown by the Monte Carlo evidence of Lütkepohl et al. (2001), the trace test has a better power performance especially in relatively small samples. The trace test can be expressed in terms of eigenvalues (λ_i) and sample size (T) with:

$$\lambda_{trace} = -\mathrm{T}\sum_{i=1}^{n}(1-\lambda_i). \tag{3.10}$$

As the basis of the Johansen test is an unrestricted VAR model, then the results associated with the Johansen test are well defined when the underlying VAR is well specified (i.e., the model is free from serial correlation). The most appropriate lag length of the VAR model is often based on model selection criteria such as the AIC. However, in the event serial correlation is existed, sufficient lags are added to remove such correlation.

Finally, we use the Gregory and Hansen (1996) residual-based cointegration test which allows for a possible structural change in the cointegrating relationship at unknown timing. Via Monte Carlo simulation technique, Gregory et al. (1996), Campos et al. (1996) and Gregory and Hansen (1996) altogether showed that the power of the Engle and Granger (1987) ADF based test which assumes constant parameter cointegration is deteriorated in the presence of a structural break. By using constant parameter cointegration tests, researchers therefore may end up with erroneous conclusions in that cointegration does not exist when it is present but governed by a structural break.

Allowing for a structural change is likely to be informative for the two sub-periods, pre-crisis and crisis. In the pre-crisis period, economies such as Japan and the euro area have been subjected to significant change. For example, Japan after a decade of deep recession started to recover in the middle of 2005 before being hit by the financial crisis. The euro also underwent significant changes rivalling the US dollar during the pre-crisis period. During the crisis period financial markets were hit by the collapse of LB, the European debt crisis, the downgrade of US government debt, etc. - a regime change might have occurred, with investors reacting differently to the new situations in the markets.[42]

Gregory and Hansen (1996) extend the Engle and Granger (1987) test to allow the intercept and the slope of the cointegrating vector to change endogenously at unknown date. They specifically propose three alternative models which accommodate structural changes in the parameters of the cointegrating vector under the alternative hypothesis.

The models are, specifically, model C (a level shift in the cointegrating relationship represented by only a change in the intercept), model C/T (a level shift in the intercept and allowing for the trend), and model C/S (a regime shift where the structural change allows both the intercept and the slope vector to shift). These models are displayed below respectively by Eqs. (3.11) – (3.13):

Model C: $$s_t = \omega_1 + \omega_2\varphi_{t\tau} + \xi e_t + v_t, \qquad t = 1,\dots, T \tag{3.11}$$

Model C/ T: $$s_t = \omega_1 + \omega_2\varphi_{t\tau} + \mu t + \xi e_t + v_t, \qquad t = 1,\dots, T \tag{3.12}$$

Model C/S: $$s_t = \omega_1 + \omega_2\varphi_{t\tau} + \xi_1 e_t + \xi_2 e_t\,\varphi_{t\tau} + v_t, \qquad t = 1,\dots, T \tag{3.13}$$

where ω_1 indicates the intercept before the shift, ω_2 indicates the change in the intercept at the time of the shift, t is the time trend, ξ_1 represents the slope of the cointegrating vector before the shift, and finally ξ_2 represents the change in the slope vector at the

[42] The Gregory and Hansen (1996) test is particularly suitable as it allows for a regime change at an unknown date and will likely capture any regime change not detected by the sample split. Using pre-specified break points instead will require prior observation of the data for each country and could introduce pre-testing problems as highlighted by Zivot and Andrews (1992). Furthermore, as pointed out by Cashin et al. (2004) there is not necessarily a one-to-one correspondence between possible causes of a structural shift and its occurrence in the data.

time of the shift. The structural change is represented by the dummy variable which is indicated as follows:

$$\varphi_{t\tau} = \begin{cases} 1 & \text{if} \quad t > [T_\tau] \\ 0 & \text{if} \quad t \leq [T_\tau] \end{cases}$$

where $\tau \in (0, 1)$ is the date of the provisional break point in the data. In particular, the endogenous breakpoint is chosen where a one-sided test statistic on the coefficient in the ADF unit root test, performed on the estimated residual e_t in each of Eqs. (3.11) - (3.13), is minimised (i.e., the most negative). Hence, such a test favours to reject the null hypothesis of no cointegration in favour of the alternative hypothesis of cointegration with a single structural break at unknown date.

3.6. Empirical results

3.6.1. Cointegration tests results

A prerequisite step of specifying the conditional mean equation, Eq. (3.1), is examining the long-run time series properties of stock market prices and exchange rates in which whether they are cointegrated or not, as stated earlier, since the series under observation appeared to be $I(1)$. In this subsection, we report the estimated cointegration results of the pairwise Engle and Granger (1987) two-step procedure, the Johansen (1995) multivariate technique, and the Gregory and Hansen (1996) test with a single structural break at unknown date, respectively.

3.6.1.1. Engle and Granger (1987) two-step cointegration tests results

The results of the Engle and Granger (1987) ADF based tests are reported in Table A3.5 (see Appendix A3) where the statistic when s_t (stock market price) is regressed on e_t (exchange rate) is listed in the third column, whereas the statistic on the reverse-order regression is reported in the last column.[43]

The results suggest that the null hypothesis of no cointegration between stock market prices and exchange rates is rejected in only one case which is in Japan, in the pre-crisis period. All other cases indicate patterns of no long-run relationship between both financial variables as the corresponding ADF test statistics are not significant.

3.6.1.2. Johansen (1995) cointegration tests results

The computed eigenvalues, trace test statistics, small sample Bartlett-corrected trace test statistics, and 95% asymptotic critical values from Johansen (1995) are reported in Table 3.2. The small sample Bartlett-corrected trace tests developed by Johansen (2002) are reported for comparison, though it seems likely for the sample size employed here that the corrected and uncorrected tests will not be dissimilar. The lag length of the VAR model is selected using the AIC, with some further lag augmentation to correct for serial correlation where appropriate.

The Johansen trace test and the Bartlett-corrected form of the test provide weak evidence of cointegration between stock prices and exchange rates. With the exception of Japan in the pre-crisis period, the null hypothesis of no cointegration (r = 0) cannot be rejected at the 5% level in all cases. To gain further insights into the long-run interrelationship between both variables in the case of Japan (in the pre-crisis period), we conduct long-run exclusion and long-run weak exogeneity tests by imposing zero restrictions respectively on each row of β and α of the long-run matrix $\Pi = \alpha\beta'$ (see Johansen, 1995). The results of these tests are listed in Table 3.3.

[43] Due to the normalisation issue of the Engle and Granger (1987) and the Gregory and Hansen (1996) tests, we report both statistics.

Table 3.2. The results of the Johansen (1995) cointegration tests between stock market prices and exchange rates.

	Sample	Lags	r	Eigenvalue	Trace test	Trace Test*	95% C.V	p-value	p-value*
US	Pre-crisis	1	$r = 0$	0.044	10.931	10.88	15.40	0.219	0.222
			$r \leq 1$	0.007	1.519	1.517	3.841	0.218	0.218
	Crisis	1	$r = 0$	0.043	11.86	11.82	15.40	0.165	0.167
			$r \leq 1$	0.008	1.843	1.840	3.841	0.175	0.175
UK	Pre-crisis	1	$r = 0$	0.043	10.13	10.09	15.40	0.275	0.278
			$r \leq 1$	0.005	1.031	1.029	3.841	0.310	0.310
	Crisis	1	$r = 0$	0.035	11.40	11.36	15.40	0.191	0.191
			$r \leq 1$	0.014	3.289	3.285	3.841	0.070	0.070
Japan	Pre-crisis	2	$r = 0$	0.102	22.58	22.23	15.40	0.003***	0.003***
			$r \leq 1$	0.001	0.200	0.194	3.841	0.655	0.660
	Crisis	2	$r = 0$	0.023	6.407	6.328	15.408	0.652	0.661
			$r \leq 1$	0.005	1.110	0.963	3.841	0.292	0.327
Euro area	Pre-crisis	1	$r = 0$	0.038	8.389	8.356	15.408	0.432	0.435
			$r \leq 1$	0.002	0.381	0.380	3.841	0.537	0.538
	Crisis	2	$r = 0$	0.020	6.472	6.385	15.408	0.645	0.655
			$r \leq 1$	0.009	1.989	1.540	3.841	0.158	0.215
Canada	Pre-crisis	1	$r = 0$	0.033	8.126	8.094	15.408	0.459	0.463
			$r \leq 1$	0.006	1.209	1.207	3.841	0.272	0.272
	Crisis	5	$r = 0$	0.040	11.27	11.27	15.40	0.198	0.198
			$r \leq 1$	0.009	2.040	2.040	3.841	0.153	0.153
Switzerland	Pre-crisis	1	$r = 0$	0.034	8.138	8.106	15.40	0.458	0.461
			$r \leq 1$	0.004	0.841	0.840	3.841	0.359	0.359
	Crisis	1	$r = 0$	0.022	5.580	5.560	15.40	0.746	0.748
			$r \leq 1$	0.002	0.498	0.497	3.841	0.480	0.481

Notes: The Table reports the Johansen trace test statistics (Johansen, 1995) and the Bartlett corrected trace tests (see Johansen, 2002) denoted by Trace test*. r is the cointegrating rank. The lag length is selected using the Akaike Information Criterion (AIC), subject to correction for serial correlation by the inclusion of further lags. The last two columns report the respective p-values.
*** indicates significance at the 1% level.

Table 3.3. Long-run exclusion (LE) and long-run weak exogeneity (WE) tests of the cointegrating relation (r =1) in Japan in the pre-crisis period.

LE tests		WE tests		The estimated long-run relationships	
Tests	$x^2(1)$	Tests	$x^2(1)$	e_t	s_t
$\beta_S = 0$	22.053***	$\alpha_S = 0$	19.869***	-1	-3.393(11.32)***
$\beta_E = 0$	19.869***	$\alpha_E = 0$	22.053***	- 0.295(11.96)***	-1

Notes: The critical values with one cointegrating vector are 6.64 and 3.84 at the 1% and 5% levels, respectively. e_t and s_t relate to the log of the exchange rate and the stock market price, respectively; -1 corresponds to the variable on which the cointegrating vector is normalised on; and t-statistics are reported in parentheses (.). *** indicates significance at the 1% level.

The long-run exclusion tests indicate that both variables are rejected to be excluded from the cointegration space; thus, the long-run relationship is well formulated. The long- run weak exogeneity tests employed indicate that neither variable is weakly exogenous. This implies that the relationship between stock prices and exchange rates for Japan (in the pre-crisis period) is bidirectional in the long-run. However, such relationship had broken down on the onset of the financial crisis because of the steady overvaluation of the yen since 2008. The yen hit a record high against the US dollar in late 2011, implying that the Japanese stock market and the yen exchange rate have not been linked since the crisis. In this case, the cointegrating relation can be normalised on each variable and the relation is found to be negative and significant in each case.

3.6.1.3. Gregory and Hansen (1996) cointegration tests results

The results of the Gregory and Hansen (1996) cointegration tests performed on models (C, C/T, C/S) are displayed in Table 3.4, where the statistics of a regression of s_t on e_t as well as of the reverse regression are reported for the two sub-periods. As shown in Table 3.4, the results indicate that the null hypothesis of no cointegration between stock prices and exchange rates is rejected for three of the cases in question. In particular, the null hypothesis of no cointegration is rejected for the euro area and Japan, in the pre-crisis period, and for the UK, in the crisis period.

In addition to Japan, the results suggest that the comovement between stock prices and exchange rates in the euro area had also broken down on the onset of the financial crisis. The depreciation of the euro and the uncertainty surrounding the single currency ever since the crisis might be the reason for the breakdown of the long-run relation. By contrast, it seems that the long-run relationship between the two financial markets in the UK was strengthened by the financial crisis, which led to both series being influenced by similar underlying factors and as a result sharing a single common stochastic trend.

No cointegration in the euro area, in the pre-crisis period, and the UK, in the crisis period, was detected by the Engle and Granger (1987), neither by the Johansen (1995) tests, as both tests are assumed to be time-invariant. This implies that the structure in these two cases has changed based on the degree of comovement between the two financial variables. However, there is no clear evidence that this was also the case in Japan (in the pre-crisis period) since the Engle and Granger (1987), the Johansen (1995), and the Gregory and Hansen (1996) tests all provide evidence of cointegration in this case. Furthermore, even though the Gregory and Hansen (1996) test detected cointegration with a structural break in Japan (in the pre-crisis); however, such a test can also detect cointegration with no structural shift and rejection of the null hypothesis of no cointegration may not be due to changes in the cointegrating relation.

It follows that to determine whether the cointegrating relationship exhibits a structural shift in the case of Japan (in the pre-crisis), following Gregory and Hansen (1996) we use the Hansen (1992a) instability test, which is applied to the residuals of a Fully Modified Ordinary Least Squares regression. The Hansen (1992a) instability test for this case implies a rejection of the null hypothesis of cointegration with constant parameters against the alternative of no cointegration due to parameter instability.[44] This suggests that the cointegrating relation in the case of Japan (in the pre-crisis period) has also been subjected to a structural shift.

[44] We use the Lc test to check the stability of the regression parameters. The *p*-value of the test statistic is 0.01.

Table 3.4. Results of the Gregory and Hansen (1996) cointergation tests allowing for a shift at unknown date.

	Model	US	UK	Euro area	Canada	Japan	Switzerland
Panel A. Pre-crisis period (August 6, 2003 – August 8, 2007)							
Regression of s_t on e_t	C	-3.191(0)	-3.081(4)	-3.828(5)	-3.670(7)	-4.702(6)**	-4.020(7)
		[2005:10:12]	[2004:03:17]	[2005:07:13]	[2006:09:20]	[2005:10:05]	[2004:04:21]
	C/T	-4.535(8)	-4.638(8)	-4.900(8)	-4.056(3)	-4.826(0)	-3.90222(1)
		[2007:01:10]	[2004:05:26]	[2004:05:26]	[2004:06:09]	[2005:10:19]	[2005:10:12]
	C/S	-4.535(0)	-3.178(4)	-3.873(3)	-3.869(0)	-4.555(6)	-4.161(1)
		[2005:10:12]	[2005:05:18]	[2005:04:20]	[2006:12:20]	[2005:10:05]	[2005:07:20]
Regression of e_t on s_t	C	-3.534(0)	-3.759(4)	-3.456(5)	-3.469(7)	-4.068(6)	-4.019(0)
		[2005:06:08]	[2006:09:13]	[2005:07:13]	[2006:09:20]	[2006:09:06]	[2004:04:21]
	C/T	-3.767(0)	-3.600(5)	-5.048(0)**	-3.536(7)	-4.454(6)	-4.677(0)
		[2005:06:08]	[2006:09:06]	[2005:05:25]	[2006:09:20]	[2004:11:10]	[2004:05:26]
	C/S	-4.537(0)	-3.836(4)	-3.498(5)	-3.874(0)	-4.338(6)	-4.069(0)
		[2005:06:08]	[2006:06:14]	[2005:07:13]	[2006:12:20]	[2006:07:12]	[2004:05:26]
Panel B. Crisis period (August 15, 2007- December 28, 2011)							
Regression of s_t on e_t	C	-3.840(0)	-4.408(0)	-2.823(7)	-3.113(0)	-3.563(0)	-2.966(5)
		[2008:10:01]	[2009:09:02]	[2008:08:06]	[2008:11:26]	[2009:12:16]	[2008:08:27]
	C/T	-4.24787 (0)	-4.433(0)	-3.106(8)	-3.137(8)	-4.041(0)	-3.133(0)
		[2008:10:01]	[2009:09:02]	[2008:07:30]	[2008:11:26]	[2008:06:04]	[2008:09:24]
	C/S	-4.28696(0)	-4.84225(0)	-2.659(7)	-3.099(8)	-4.336(0)	-3.746(5)
		[2008:09:10]	[2009:05:13]	[2008:08:06]	[2008:11:12]	[2009:07:15]	[2009:09:30]
Regression of e_t on s_t	C	-3.530(5)	-4.561(0)	-3.706(7)	-3.221(8)	-3.598(0)	-2.898(6)
		[2008:08:27]	[2009:09:02]	[2010:03:17]	[2009:08:05]	[2009:12:16]	[2010:06:30]
	C/T	-4.234(6)	-4.454(0)	-3.953(7)	-3.195(8)	-4.069(0)	-3.792(5)
		[2010:02:03]	[2008:11:19]	[2010:03:17]	[2009:07:29]	[2008:06:04]	[2010:10:20]
	C/S	-3.447(5)	-5.124(0)**	-3.963(7)	-3.288(8)	-3.919(0)	-3.207(6)
		[2008:04:23]	[2009:09:02]	[2010:04:14]	[2009:12:16]	[2009:12:16]	[2010:04:14]

Notes: The test due to Gregory and Hansen (1996) is conducted by regressing s_t on e_t and in the reverse regression. Model C allows for a shift in the intercept, Model C/T allows for a shift in the intercept and the trend, and Model C/S allows for a shift in both the intercept and the slope coefficient of the cointegrating relationship. The corresponding critical values for each model are from Table 1 in Gregory and Hansen (1996). The lag order is chosen on the basis of *t*-tests, reported in parenthesis (.), subject to a maximum of 8 lags. Breakpoints are reported in square brackets [.].
** indicates significance at the 5% level.

Next, we include the breaks identified by the Gregory and Hansen (1996) tests for the three cases in the VECM models, Eq. 3.2, in order to further capture the structural change in the relationship between stock prices and exchange rates. The identified break points are May 25, 2005, for the euro area; October 5, 2005, for Japan; and September 2, 2009, for the UK.

Note that the lack of cointegrating relations may due to model misspecification. In the long run, other fundamental economic variables may work as channels to link the two types of financial markets (stock and foreign exchange markets). However, our findings of limited cointegration between stock prices and exchange rates are in line with much of the existing empirical literature (see, e.g., Bahmani-Oskooee and Sohrabian, 1992; Granger et al., 2000; Nieh and Lee, 2001; Alagidede et al., 2011; among others).

3.6.2. VAR-BEKK results

The quasi-maximum likelihood estimates of the bivariate GARCH-BEKK parameters along with the associated multivariate *Q*-statistics (Hosking, 1981) are displayed in Tables 3.5–3.10 for Japan, the US, the euro area, Canada, Switzerland, and the UK in turn. Panels A and B concern the pre-crisis and crisis periods, respectively. On the basis of the results of the cointegration tests of subsection 3.6.1, the lagged error correction terms are included in the conditional mean equations for the cases for which cointegration was detected. Furthermore, since the detected cointegrating relations have been subject to structural change, causality parameters as well as parameters related to the lagged error correction terms in the VECM models are allowed to shift at the break points. The Hosking multivariate *Q*-statistics for the standardised residuals indicate no serial correlation at the 5% level. In all cases, a lag length of 1 captures the dynamics, except for the UK in the pre-crisis sample where $p = 3$ and the US and Switzerland in the crisis sample where p = 3 and p = 5 are required, respectively (note that the insignificant parameters in the mean equations are excluded). Overall, the estimated models appear to be well specified.

The dynamic interactions between the first moments of stock returns and exchange rate changes, captured by $\psi_{ES}^{(i)}$ and $\psi_{SE}^{(i)}$, suggest that there are limited dynamic linkages between the two variables in the pre-crisis compared with the crisis period. The results for the pre-crisis period imply the existence of unidirectional Granger causality from stock returns to exchange rate changes in the case of Japan, while there is causality in the opposite direction in the UK. However, since lagged error correction terms are included in the cases of cointegration, then the VECM model will allow to further differentiate between short-run and long-run Granger causality. In specific, long-run causality between the two variables will be through the error correction term if this is negative and significant as, for example, in Japan in the equations for both stock returns and exchange rate changes. This implies that both variables adjust to the steady-state equilibrium in Japan, and there is bidirectional feedback. However, the speed of adjustment of exchange rate changes towards equilibrium becomes slower after the break on October 5, 2005 as η_{E}^{*} is positive and significant.[45] By contrast, the lagged error correction term in the euro area is negative and significant only in the equation for exchange rate changes, suggesting that the adjustment towards equilibrium takes place through this variable.

In the crisis period, instead, the results indicate the existence of causality from stock returns to exchange rate changes in the US and the UK, in the opposite direction in Canada, and bidirectional causality in the euro area and Switzerland. With regard to

[45] Note that the error correction term is calculated from the estimation of a cointegrating relation of the stock price on the exchange rate.

the UK, the lagged error correction term in the equation for exchange rate changes is found to be negative and significant, implying an adjustment mechanism through the exchange rate and enforcing the evidence of causality from stock returns to exchange rate changes in the long-run, as well.

In contrast to the case of Japan (in the pre-crisis period), the impact of the breaks, identified by the Gregory and Hansen (1996) tests, on the linkages between stock returns and exchange rate changes in the euro area (in the pre-crisis period) and the UK (in the crisis period) seem to be limited, hence the causal structure between the two variables as well as the speed of adjustment towards the equilibrium are stable in these two cases. The significant change in the speed of adjustment of the exchange rate towards equilibrium in the case of Japan (in the pre-crisis period) may be due to the acceleration in the depreciation of the yen as a result of the decline in the Japanese long-term real interest rates compared to those for the US over the period April 2005 to June 2006 (Obstfeld, 2009).

To analyse further the dynamic linkages between stock returns and exchange rate changes, we estimate the GIRFs of Pesaran and Shin (1998) for the cases where Granger causality is not rejected. Overall, the results of the GIRFs (8 periods) from one standard error shock of the variable in question, displayed respectively in Figure 3.2a and Figure 3.2b for the pre-crisis and crisis periods, are in line with the findings for Granger causality.

In the pre-crisis period (see Figure 3.2a), a one standard error shock to stock returns in Japan leads to an appreciation of the yen in the first week. This is line with the portfolio approach on the linkage between stock prices and exchange rates, suggesting that stock prices lead exchange rates with a positive correlation. In the UK, the response of stock returns to a one standard error shock to exchange rate changes is significant and positive in the third week. This is consistent with the traditional

Table 3.5. The estimated bivariate GARCH-BEKK model for Japan.

Panel A. Pre-crisis period (6/8/ 2003-8/8/2007)			*Panel B. Crisis period (15/8/2007-28/12/2011)*		
	$R_{S,t}$ (i=S)	$R_{E,t}$ (i=E)		$R_{S,t}$ (i=S)	$R_{E,t}$ (i=E)
Conditional Mean Equation					
μ_i	0.065 (0.134)	$-$0.097 (0.064)	μ_i	-0.283^{**} (0.136)	0.202^{***} (0.072)
$\psi_{Si}^{(1)}$	0.136^{**} (0.057)		$\psi_{Ei}^{(1)}$		-0.151^{***} (0.049)
$\psi_{Ei}^{(1)}$	0.052^{**} (0.024)		λ_i	0.966^{***} (0.053)	-0.271^{***} (0.030)
η_i	-3.420^{**} (1.570)	-5.484^{***} (1.038)	δ_i		-0.044^{***} (0.015)
η_i^*		4.409^{**} (1.750)			
λ_i	0.859^{***} (0.089)				
Conditional Variance Equation					
c_{Si}	0.702^{***} (0.151)	0	c_{Si}	1.513^{***} (0.291)	0
c_{Ei}	-0.554^{***} (0.046)	0.000002 (0.353)	c_{Ei}	-0.591^{***} (0.080)	$-$0.00001 (0.044)
a_{Si}	0.159^{***} (0.057)	0.062 (0.065)	a_{Si}	0.053 (0.145)	0.115 (0.091)
a_{Ei}	-0.646^{***} (0.052)	0.391^{***} (0.121)	a_{Ei}	-0.466^{***} (0.136)	0.312^{***} (0.088)
b_{Si}	0.855^{***} (0.041)	0.124^{***} (0.027)	b_{Si}	0.719^{***} (0.037)	0.125^{***} (0.025)
b_{Ei}	0.147^{*} (0.091)	0.712^{***} (0.028)	b_{Ei}	0.053 (0.074)	0.931^{***} (0.044)
Loglik	$-$694.856		*Loglik*	$-$862.384	
$Q(6)$ 9.426[0.894]	$Q^2(6)$ 15.119[0.653]		$Q(6)$ 20.755[0.188]	$Q^2(6)$ 20.392[0.118]	
$Q(12)$ 23.78[0.851]	$Q^2(12)$ 43.472[0.249]		$Q(12)$ 32.592[0.437]	$Q^2(12)$ 26.587[0.644]	

Panel A		Panel B	
Tests of No Volatility Transmission:		Tests of No Volatility Transmission:	
(i) Bidirectional between $R_{S,t}$ and $R_{E,t}$		(i) Bidirectional between $R_{S,t}$ and $R_{E,t}$	
$H_0 \ : \ a_{SE} = a_{ES} = b_{SE} = b_{ES} = 0$	*LR*=14.103	$H_0 \ : \ a_{SE} = a_{ES} = b_{SE} = b_{ES} = 0$	*LR*=10.674
(ii) From $R_{S,t}$ to $R_{E,t}$		(ii) From $R_{S,t}$ to $R_{E,t}$	
$H_0 \ : \ a_{SE} = b_{SE} = 0$	*LR*= 5.637	$H_0 \ : \ a_{SE} = b_{SE} = 0$	*LR*= 8.177
(iii) From $R_{E,t}$ to $R_{S,t}$		(iii) From $R_{E,t}$ to $R_{S,t}$	
$H_0 \ : \ a_{ES} = b_{ES} = 0$	*LR*=12.343	$H_0 \ : \ a_{ES} = b_{ES} = 0$	*LR*= 5.637

Notes: $R_{S,t}$ and $R_{E,t}$ indicate stock market returns and exchange rate changes, respectively, while *LR* indicates the likelihood ratio test statistics. Heteroscedasticity-consistent standard errors are reported in parentheses (.), whereas *p*-values are reported in [.]. The superscripts of the ψ parameters denote the lagged time periods. $Q(p)$ and $Q^2(p)$ are the multivariate Hosking (1981) tests for the p^{th} order serial correlation on the standardised residuals, $Z_{i,t}$, and their squares, $Z_{i,t}^2$, respectively, where i = S (for stock market returns), E (for exchange rate changes). All the eigenvalues of $A_{11} \otimes A_{11} + B_{11} \otimes B_{11}$ being less than one in modulus, hence the covariance stationarity condition is fulfilled by all the estimated models.
***, **, and * indicate significance at the 1 %, 5%, and 10% levels, respectively.

Table 3.6. The estimated bivariate GARCH-BEKK model for the US.

	Panel A. Pre-crisis period (6/8/ 2003-8/8/2007)			*Panel B. Crisis period (15/8/2007-28/12/2011)*	
	$R_{S,t}\,(i{=}S)$	$R_{E,t}\,(i{=}E)$		$R_{S,t}\,(i{=}S)$	$R_{E,t}\,(i{=}E)$
Conditional Mean Equation					
μ_i	0.025 (0.053)	0.016 (0.046)	μ_i	0.165^{***} (0.082)	−0.016 (0.058)
$\psi_{Si}^{(1)}$	-0.217^{***} (0.043)		$\psi_{Si}^{(1)}$	-0.068^{**} (0.030)	
λ_i	0.522^{***} (0.033)	-0.244^{***} (0.026)	$\psi_{Ei}^{(1)}$	0.071^{**} (0.034)	
δ_i	-0.028^{***} (0.006)		$\psi_{Ei}^{(2)}$		0.128^{***} (0.043)
			$\psi_{Si}^{(3)}$	0.055^{*} (0.028)	
			λ_i	0.735^{***} (0.034)	-0.172^{***} (0.028)
			δ_i		-0.029^{**} (0.012)
Conditional Variance Equation					
c_{Si}	0.211 (0.260)	0	c_{Si}	0.093 (0.148)	0
c_{Ei}	0.260 (0.095)	0.0001 (0.147)	c_{Ei}	0.286^{***} (0.084)	−0.000006 (0.147)
a_{Si}	0.032 (0.092)	−0.028 (0.099)	a_{Si}	0.474^{***} (0.078)	0.120^{***} (0.054)
a_{Ei}	-0.207^{**} (0.082)	0.299^{***} (0.090)	a_{Ei}	−0.009 (0.115)	-0.213^{***} (0.055)
b_{Si}	0.150 (0.097)	-0.573^{***} (0.100)	b_{Si}	0.718^{***} (0.054)	-0.188^{***} (0.044)
b_{Ei}	0.610^{***} (0.126)	0.604^{***} (0.094)	b_{Ei}	0.543^{***} (0.083)	0.988^{***} (0.048)
Loglik	−504.510		*Loglik*	−670.683	
$Q(6)$	9.492[0.891]	$Q^2(6)$ 5.497[0.977]	$Q(6)$	25.932[0.054]	$Q^2(6)$ 21.370[0.092]
$Q(12)$	19.83[0.954]	$Q^2(12)$ 26.13[0.668]	$Q(12)$	43.881[0.078]	$Q^2(12)$ 41.429[0.080]

Panel A		Panel B	
Tests of No Volatility Transmission:		Tests of No Volatility Transmission:	
(i) Bidirectional between $R_{S,t}$ and $R_{E,t}$		(i) Bidirectional between $R_{S,t}$ and $R_{E,t}$	
$H_0\ :\ a_{SE}=a_{ES}=b_{SE}=b_{ES}=0$	*LR*=5.258	$H_0\ :\ a_{SE}=a_{ES}=b_{SE}=b_{ES}=0$	*LR*=28.155
(ii) From $R_{S,t}$ to $R_{E,t}$		(ii) From $R_{S,t}$ to $R_{E,t}$	
$H_0\ :\ a_{SE}=b_{SE}=0$	*LR*=2.894	$H_0\ :\ a_{SE}=b_{SE}=0$	*LR*=24.591
(iii) From $R_{E,t}$ to $R_{S,t}$		(iii) From $R_{E,t}$ to $R_{S,t}$	
$H_0\ :\ a_{ES}=b_{ES}=0$	*LR*=0.638	$H_0\ :\ a_{ES}=b_{ES}=0$	*LR*=21.034

Notes: See notes of Table 3.5.

Table 3.7. The estimated bivariate GARCH-BEKK model for the euro area.

Panel A. Pre-crisis period (6/8/ 2003-8/8/2007)			*Panel B. Crisis period (15/8/2007-28/12/2011)*		
	$R_{S,t}\,(i{=}S)$	$R_{E,t}\,(i{=}E)$		$R_{S,t}\,(i{=}S)$	$R_{E,t}\,(i{=}E)$
Conditional Mean Equation					
μ_i	0.118 (0.084)	-0.235^{*} (0.142)	μ_i	-0.176 (0.166)	-0.201^{***} (0.076)
η_i		-7.076^{***} (2.336)	$\psi_{Si}^{(1)}$	-0.175^{***} (0.043)	0.378^{**} (0.169)
λ_i	0.857^{***} (0.054)		$\psi_{Ei}^{(1)}$	-0.032^{**} (0.014)	
γ_i		0.096^{**} (0.045)	λ_i	1.000^{***} (0.051)	0.045^{*} (0.024)
			γ_i		0.050^{**} (0.023)
			δ_i		0.060^{***} (0.012)
Conditional Variance Equation					
c_{Si}	0.734^{**} (0.146)	0	c_{Si}	0.098 (0.221)	0
c_{Ei}	0.116 (0.074)	-0.003 (0.158)	c_{Ei}	-0.589^{***} (0.082)	-0.00001 (0.402)
a_{Si}	0.362^{***} (0.100)	0.038 (0.029)	a_{Si}	0.310^{***} (0.107)	-0.030 (0.036)
a_{Ei}	-0.033 (0.341)	0.163^{*} (0.102)	a_{Ei}	-0.461^{*} (0.249)	0.666^{***} (0.172)
b_{Si}	0.733^{***} (0.094)	-0.036 (0.042)	b_{Si}	0.870^{***} (0.040)	0.021 (0.031)
b_{Ei}	-0.223 (0.207)	0.951^{***} (0.056)	b_{Ei}	0.781^{***} (0.192)	0.435^{**} (0.205)
Loglik	-512.1292		*Loglik*	-794.673	
$Q(6)$	25.098[0.068]	$Q^2(6)$ 19.052[0.162]	$Q(6)$	16.313[0.431]	$Q^2(6)$ 10.362[0.664]
$Q(12)$	34.407[0.353]	$Q^2(12)$ 38.467[0.138]	$Q(12)$	35.831[0.293]	$Q^2(12)$ 25.235[0.665]

Panel A		Panel B	
Tests of No Volatility Transmission:		Tests of No Volatility Transmission:	
(i) Bidirectional between $R_{S,t}$ and $R_{E,t}$		(i) Bidirectional between $R_{S,t}$ and $R_{E,t}$	
$H_0\ :\ a_{SE}=a_{ES}=b_{SE}=b_{ES}=0$	LR=3.740	$H_0\ :\ a_{SE}=a_{ES}=b_{SE}=b_{ES}=0$	LR=7.254
(ii) From $R_{S,t}$ to $R_{E,t}$		(ii) From $R_{S,t}$ to $R_{E,t}$	
$H_0\ :\ a_{SE}=b_{SE}=0$	LR=1.567	$H_0\ :\ a_{SE}=b_{SE}=0$	LR=0.926
(iii) From $R_{E,t}$ to $R_{S,t}$		(iii) From $R_{E,t}$ to $R_{S,t}$	
$H_0\ :\ a_{ES}=b_{ES}=0$	LR=2.491	$H_0\ :\ a_{ES}=b_{ES}=0$	LR=9.466

Notes: See notes of Table 3.5.

Table 3.8. The estimated bivariate GARCH-BEKK model for Canada.

Panel A. Pre-crisis period (6/8/ 2003-8/8/2007)			*Panel B. Crisis period (15/8/2007-28/12/2011)*		
	$R_{S,t}\,(i=S)$	$R_{E,t}\,(i=E)$		$R_{S,t}\,(i=S)$	$R_{E,t}\,(i=E)$
Conditional Mean Equation					
μ_i	0.170^{**} (0.073)	0.105^{*} (0.061)	μ_i	-0.109 (0.091)	0.037 (0.071)
λ_i	0.750^{***} (0.056)	0.152^{***} 0.052	$\psi_{Si}^{(1)}$		-0.136^{**} (0.057)
			$\psi_{Ei}^{(1)}$		-0.107^{**} (0.051)
			λ_i	0.641^{***} (0.037)	0.279^{***} (0.025)
			δ_i	0.143^{***} (0.016)	0.061^{***} (0.014)
Conditional Variance Equation					
c_{Si}	0.371^{***} (0.133)	0	c_{Si}	0.256^{***} (0.153)	0
c_{Ei}	-0.210^{*} (0.126)	-0.000001 (0.075)	c_{Ei}	0.426^{***} (0.119)	0.000002 (0.091)
a_{Si}	0.258^{***} (0.068)	0.127 (0.083)	a_{Si}	-0.247^{*} (0.142)	0.254^{***} (0.090)
a_{Ei}	0.356^{***} (0.080)	-0.086 (0.112)	a_{Ei}	0.152^{***} (0.052)	0.454^{***} (0.118)
b_{Si}	0.846^{***} (0.042)	-0.097^{*} (0.052)	b_{Si}	0.915^{***} (0.069)	-0.128^{*} (0.077)
b_{Ei}	0.200^{***} (0.066)	0.952^{***} (0.024)	b_{Ei}	0.150^{*} (0.078)	0.784^{***} (0.106)
Loglik	-604.2614		*Loglik*	-749.813	
$Q(6)$ 20.303[0.206]		$Q^2(6)$ 19.401[0.150]	$Q(6)$ 20.990[0.178]		$Q^2(6)$ 10.310[0.739]
$Q(12)$ 37.567[0.229]		$Q^2(12)$ 31.000[0.415]	$Q(12)$ 31.581[0.487]		$Q^2(12)$ 29.445[0.494]

Panel A. Tests of No Volatility Transmission:

(i) Bidirectional between $R_{S,t}$ and $R_{E,t}$

$H_0\,:\,a_{SE}=a_{ES}=b_{SE}=b_{ES}=0$ LR=8.246

(ii) From $R_{S,t}$ to $R_{E,t}$

$H_0\,:\,a_{SE}=b_{SE}=0$ LR=3.171

(iii) From $R_{E,t}$ to $R_{S,t}$

$H_0\,:\,a_{ES}=b_{ES}=0$ LR=8.016

Panel B. Tests of No Volatility Transmission:

(i) Bidirectional between $R_{S,t}$ and $R_{E,t}$

$H_0\,:\,a_{SE}=a_{ES}=b_{SE}=b_{ES}=0$ LR=26.314

(ii) From $R_{S,t}$ to $R_{E,t}$

$H_0\,:\,a_{SE}=b_{SE}=0$ LR=20.633

(iii) From $R_{E,t}$ to $R_{S,t}$

$H_0\,:\,a_{ES}=b_{ES}=0$ LR=19.299

Notes: See notes of Table 3.5.

Table 3.9. The estimated bivariate GARCH-BEKK model for Switzerland.

Panel A. Pre-crisis period (6/8/ 2003-8/8/2007)			*Panel B. Crisis period (15/8/2007-28/12/2011)*		
	$R_{S,t}\,(i{=}S)$	$R_{E,t}\,(i{=}E)$		$R_{S,t}\,(i{=}S)$	$R_{E,t}\,(i{=}E)$
Conditional Mean Equation					
μ_i	0.192^{**} (0.092)	-0.016 (0.031)	μ_i	-0.136 (0.117)	0.172^{***} (0.048)
$\psi_{Si}^{(1)}$	-0.095^{**} (0.047)		$\psi_{Si}^{(1)}$	-0.107^{**} (0.048)	
λ_i	0.803^{***} (0.051)	-0.043^{**} (0.022)	$\psi_{Ei}^{(1)}$	0.048^{***} (0.017)	
δ_i	0.020^{*} (0.010)		$\psi_{Si}^{(3)}$		-0.166^{*} (0.092)
			$\psi_{Ei}^{(4)}$		-0.195^{***} (0.061)
			$\psi_{Ei}^{(5)}$		-0.185^{***} (0.054)
			λ_i	0.643^{***} (0.040)	-0.096^{***} (0.017)
Conditional Variance Equation					
c_{Si}	0.513^{***} (0.174)	0	c_{Si}	1.268^{***} (0.251)	0
c_{Ei}	-0.019 (0.060)	0.104^{*} (0.058)	c_{Ei}	-0.306^{***} (0.063)	-0.000001 (0.052)
a_{Si}	-0.047 (0.139)	0.120^{***} (0.030)	a_{Si}	0.203^{**} (0.081)	0.047^{**} (0.022)
a_{Ei}	0.819^{**} (0.412)	0.146 (0.188)	a_{Ei}	-0.553^{***} (0.174)	0.687^{***} (0.109)
b_{Si}	0.848^{***} (0.052)	-0.018 (0.049)	b_{Si}	0.665^{***} (0.130)	0.112^{***} (0.030)
b_{Ei}	-0.005 (0.217)	0.908^{***} (0.045)	b_{Ei}	0.252^{**} (0.117)	0.754^{***} (0.07)
Loglik	-452.951		*Loglik*	-746.106	
$Q(6)$ 20.090[0.216]		$Q^2(6)$ 14.296[0.427]	$Q(6)$ 30.444[0.062]		$Q^2(6)$ 8.630[0.853]
$Q(12)$ 32.829[0.426]		$Q^2(12)$ 29.145[0.509]	$Q(12)$ 39.509[0.169]		$Q^2(12)$ 20.28[0.908]

Panel A		Panel B	
Tests of No Volatility Transmission:		Tests of No Volatility Transmission:	
(i) Bidirectional between $R_{S,t}$ and $R_{E,t}$		(i) Bidirectional between $R_{S,t}$ and $R_{E,t}$	
$H_0\ :\ a_{SE}=a_{ES}=b_{SE}=b_{ES}=0$	LR=9.734	$H_0\ :\ a_{SE}=a_{ES}=b_{SE}=b_{ES}=0$	LR=18.952
(ii) From $R_{S,t}$ to $R_{E,t}$		(ii) From $R_{S,t}$ to $R_{E,t}$	
$H_0\ :\ a_{SE}=b_{SE}=0$	LR=4.945	$H_0\ :\ a_{SE}=b_{SE}=0$	LR= 5.262
(iii) From $R_{E,t}$ to $R_{S,t}$		(iii) From $R_{E,t}$ to $R_{S,t}$	
$H_0\ :\ a_{ES}=b_{ES}=0$	LR=1.316	$H_0\ :\ a_{ES}=b_{ES}=0$	LR=13.659

Notes: See notes of Table 3.5.

Table 3.10. The estimated bivariate GARCH-BEKK model for the UK.

Panel A. Pre-crisis period (6/8/ 2003-8/8/2007)			*Panel B. Crisis period (15/8/2007-28/12/2011)*		
	$R_{S,t}$ $(i{=}S)$	$R_{E,t}$ $(i{=}E)$		$R_{S,t}$ $(i{=}S)$	$R_{E,t}$ $(i{=}E)$
Conditional Mean Equation					
μ_i	0.117^{*} (0.071)	0.055 (0.044)	μ_i	−0.108 (0.136)	−0.042 (0.065)
$\psi_{Si}^{(1)}$	-0.102^{**} (0.046)		$\psi_{Si}^{(1)}$	-0.145^{***} (0.049)	
$\psi_{Si}^{(3)}$		0.235^{**} (0.097)	$\psi_{Ei}^{(1)}$	-0.062^{***} (0.020)	
$\psi_{Ei}^{(3)}$		-0.137^{**} (0.062)	η_i		-2.645^{***} (0.724)
λ_i	0.664^{***} (0.045)		λ_i	0.745^{***} (0.050)	
			δ_i	0.084^{***} (0.024)	0.045^{***} (0.015)
Conditional Variance Equation					
c_{Si}	0.092 (0.090)	0	c_{Si}	0.352^{**} (0.152)	0
c_{Ei}	-0.633^{***} (0.039)	−0.00001 (0.435)	c_{Ei}	−0.137 (0.472)	0.382^{***} (0.110)
a_{Si}	0.120 (0.096)	−0.063 (0.128)	a_{Si}	0.228^{***} (0.073)	−0.047 (0.043)
a_{Ei}	-0.301^{**} (0.144)	0.406^{***} (0.099)	a_{Ei}	−0.025 (0.102)	0.474^{***} (0.156)
b_{Si}	0.629^{***} (0.028)	−0.078 (0.139)	b_{Si}	0.959^{***} 0.025	0.002 (0.058)
b_{Ei}	-0.821^{***} (0.046)	0.058 (0.280)	b_{Ei}	0.041 (0.047)	0.801^{***} (0.113)
Loglik	−508.877		*Loglik*	−800.1193	
$Q(6)$	17.948[0.265]	$Q^2(6)$ 14.714[0.397]	$Q(6)$	18.119[0.316]	$Q^2(6)$ 7.921[0.893]
$Q(12)$	42.182[0.107]	$Q^2(12)$ 23.148[0.809]	$Q(12)$	29.472[0.595]	$Q^2(12)$ 15.29[0.987]

Panel A.

Tests of No Volatility Transmission:

(i) Bidirectional between $R_{S,t}$ and $R_{E,t}$

$H_0 : a_{SE} = a_{ES} = b_{SE} = b_{ES} = 0$ *LR*=12.821

(ii) From $R_{S,t}$ to $R_{E,t}$

$H_0 : a_{SE} = b_{SE} = 0$ *LR*=0.362

(iii) From $R_{E,t}$ to $R_{S,t}$

$H_0 : a_{ES} = b_{ES} = 0$ *LR*=11.817

Panel B.

Tests of No Volatility Transmission:

(i) Bidirectional between $R_{S,t}$ and $R_{E,t}$

$H_0 : a_{SE} = a_{ES} = b_{SE} = b_{ES} = 0$ *LR*= 2.313

(ii) From $R_{S,t}$ to $R_{E,t}$

$H_0 : a_{SE} = b_{SE} = 0$ *LR*= 1.317

(iii) From $R_{E,t}$ to $R_{S,t}$

$H_0 : a_{ES} = b_{ES} = 0$ *LR*= 0.867

Notes: See notes of Table 3.5.

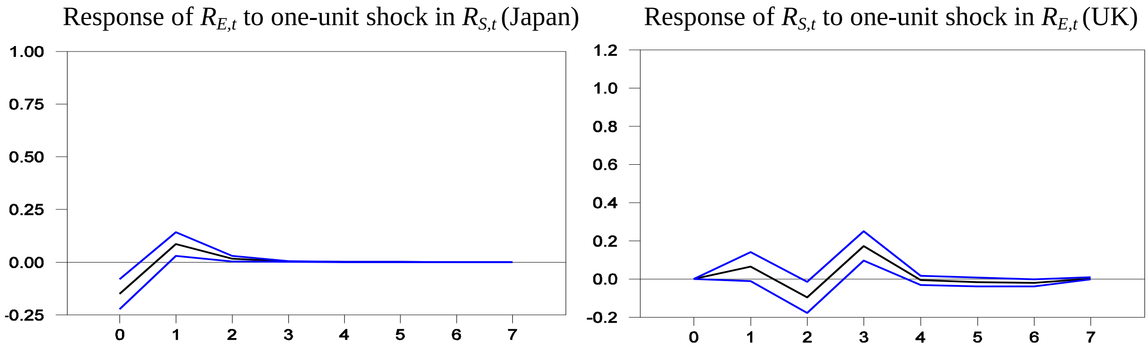

Figure 3.2a. Generalised impulse responses of significant short-run Granger causality between stock returns and exchange rate changes in the pre-crisis period (August 6, 2003-August 8, 2007).

approach on the linkage between the two variables, suggesting that exchange rates lead stock prices. However, the sign of the correlation can go either way depending on whether the domestic firm is an exporter or an importer. The net effect on the aggregate stock index cannot be determined a priori and hence the sign can be either positive or negative (Granger et al., 2000).

With regard to the crisis period (see Figure 3.2b), it is found that a one standard deviation shock to stock returns results in a depreciation of the exchange rate in the UK, which is not consistent with the portfolio approach, unlike in the US, where the positive sign supports empirically this approach. In Canada a shock to exchange rates decreases stock returns, in line with the traditional approach.

In Switzerland a shock to stock returns has a positive impact on corresponding exchange rates in the first week, whilst the response of stock returns to a shock to exchange rates is negative and significant in the third week. Finally, exchange rate changes (stock returns) in the euro area respond negatively (positively) to a shock to stock returns (exchange rate changes).

Granger et al. (2000) concluded that capital flows played a major role in the interactions between stock prices and exchange rates during the Asian flu period. Figure

3.3 shows the evolution of portfolio investment liabilities and current accounts as a percentage of GDP in all countries over the sample period considered here. The empirical support for the portfolio approach found for the US could be the result of capital flows. Given that the US was the centre of the crisis, the decline in its stock market at the onset of the crisis in late 2007, along with the collapse of LB and the downgrade of its debt status, induced capital outflows (see Figure 3.3) and a depreciation of its currency. With regard to the UK, the collapse of LB in the US and the shutdown of its offices in London sent a wave of panic right through the UK stock market followed by a severe downturn[46] and major changes in the British pound over the crisis period. Nonetheless, the causal effect as measured from the impulse responses seems to be more complex than implied by the portfolio approach as the sign does not validate such an approach, as stated earlier. By contrast, Canada experienced capital inflows as opposed to outflows during the crisis (see Figure 3.3) and its currency strengthened after 2009 (see Figure 3.1a) leading its stock market.

The lack of any interactions between stock returns and exchange rate changes in Japan, by contrast, can be attributed to country-specific factors. The fact that Japan has amassed huge foreign exchange reserves and had a strengthening real economy played a significant role in the appreciation of its currency and making it immune to the crisis (Wong and Li, 2010). Indeed, Obstfeld et al. (2009) showed that countries with large reserves exhibited less currency pressure. The finding also reflects the overall state of the Asian and Australasian countries whose banks and economies appeared not to be contaminated by the crisis that has been linked to the failure of the valuation of complex derivatives.

[46] According to Alistair Darling, the then Chancellor of the Exchequer, liquidity was compromised so that there was a fear that the cash machines would be empty.

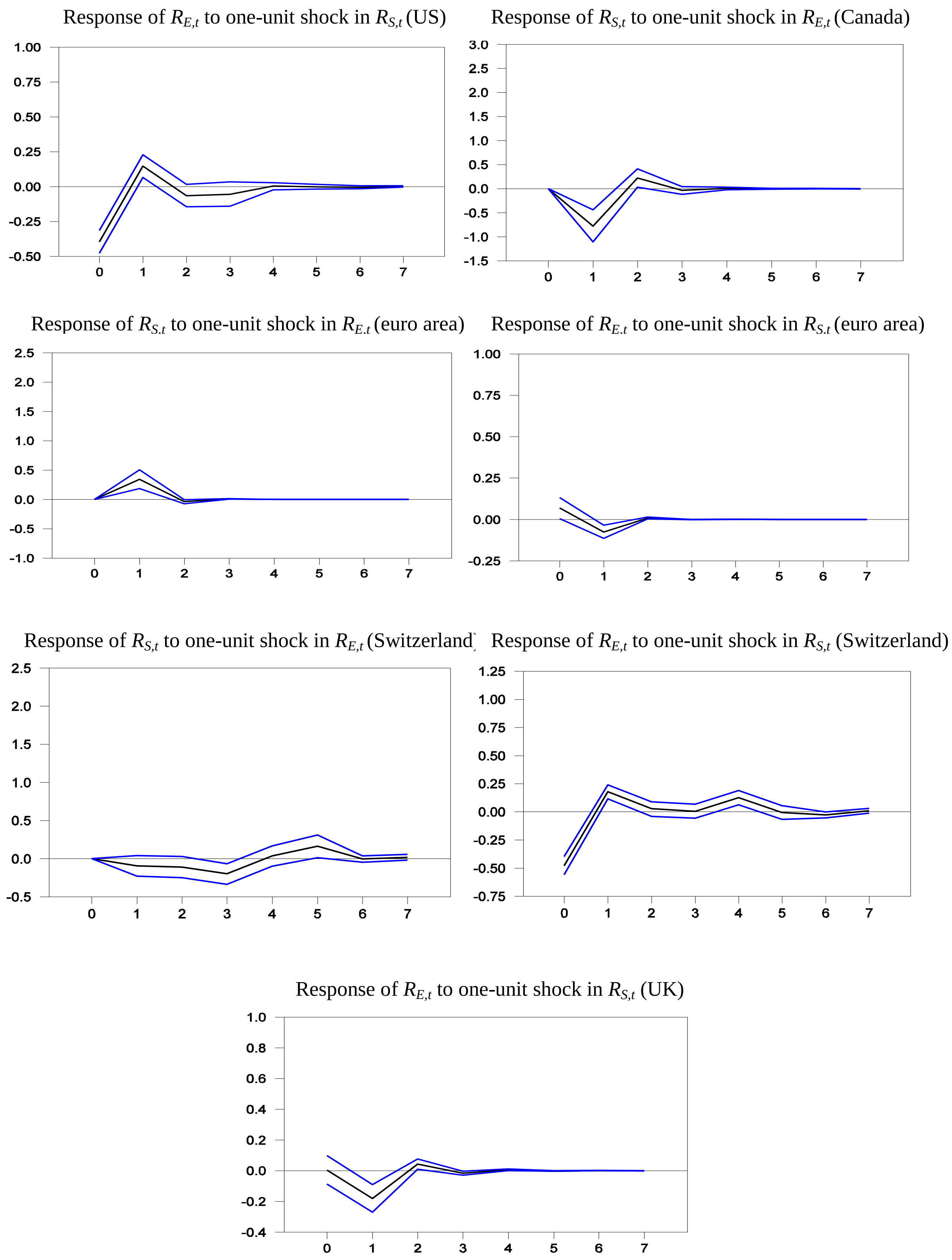

Figure 3.2b. Generalised impulse responses of significant short-run Granger causality between stock returns and exchange rate changes in the crisis period (August 15, 2007-December 28, 2011).

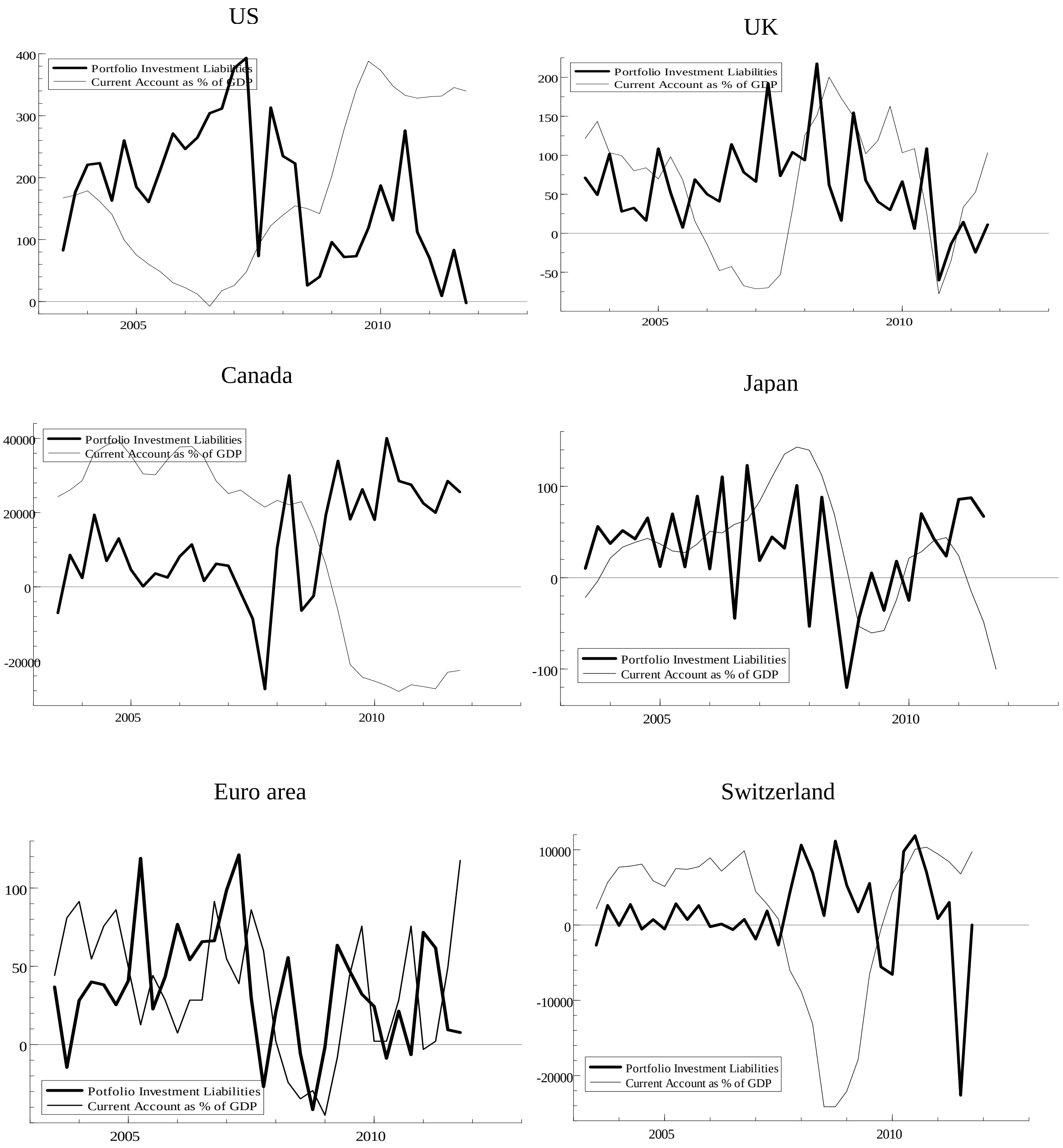

Figure 3.3. The evolution of portfolio investment liabilities and current accounts as percentage of GDP over the sample period (August 6, 2003-December 28, 2011) (Source: Bloomberg).

As far as the exogenous variables in the conditional mean equations are concerned, the return of the world stock index exerts strong influence on stock returns and exchange rate changes in most cases, especially in the crisis period, suggesting its dominance in the transmission of shocks and information to other markets around the globe. Specifically, the influence on stock returns is significant in all countries in the two sub-periods. With regard to its impact on exchange rate changes, the influence is significant in the US, Canada, and Switzerland, in the two sub-periods, and Japan and the euro area, in the crisis period.

The impact of the domestic interest rate, by contrast, appears to be limited. It only impacts on exchange rate changes in the euro area in the two sub-periods. The limited impact of the interest rate reinforces the notion that the quantitative easing policies adopted by the monetary authorities throughout the crisis period were ineffective (see also Lyonnet and Werner, 2012). One possible explanation is that the economic cycle did not respond because of the breakdown of both the financial system and the monetary transmission mechanism via the banks.

With regard to the influence of world oil price changes, this increased in the crisis period compared with the pre-crisis one in most countries, except Switzerland. The effects on stock returns in the case of Switzerland, in the pre-crisis period, and Canada and the UK, in the crisis period, are consistent with the findings of Filis et al. (2011), who argued that stock markets react positively to demand-side oil price shocks. The two periods in this chapter are characterised by such shocks. The pre-crisis period was accompanied by an increase in oil prices because of an increase in demand, primarily in China, whereas in the crisis period there was a decline in oil prices as a consequence of the global recession induced by the financial crisis.

The increased impact of oil price changes on exchange rate changes in the crisis period compared with the pre-crisis one, on the other hand, is in line with the findings of Roboredo (2011). During the crisis, the impact on the US dollar and the Japanese yen is negative, whereas the impact on the British pound, the Canadian dollar, and the euro is positive. The negative influence on the US dollar and the Japanese yen may due to the fact that the US and Japan are the first and third largest oil consumer and importer economies.[47] Hence, funding such huge oil imports in the US and Japan has a significant impact on the value of the dollar and yen. The positive influence of the oil price changes on the Canadian dollar is explained by the fact that Canada is a net-exporting economy.

Finally, with regard to the UK and the euro area, Lizardo and Mollick (2010) argued that the net imports and/or net exports of oil for these economies represent a small proportion of their total trade compared with the US and Japan. Thus, an increase in oil prices results in an appreciation in the exchange rates of these economies.

Next, the estimates of the conditional variance equations suggest that the stock return–exchange rate changes process in the two sub-periods displays strong conditional heteroscedasticity: the diagonal elements of the ARCH matrices are significant in 58% of the cases in the pre-crisis period and 91.6% of the cases during the crisis period. The conditional variances, on the other hand, exhibit persistence in all cases with only two exceptions, i.e., US stock returns and UK exchange rate changes in the pre-crisis period. More specifically, the estimated conditional variances of stock returns range from 0.62 (the lowest) for the UK to 0.85 (the highest) for Japan in the pre-crisis period, whilst they range from 0.66 (the lowest) for Switzerland to 0.95 (the highest) for the UK in the crisis period. The corresponding estimates of exchange rate changes range from 0.60

[47] US imports of crude oil in 2009 were 18.690 million barrels per day, which accounted for approximately 55% of its domestic consumption. Meanwhile, Japan imports in 2009 were 4.394 million barrels per day, which accounted for approximately 100% of its domestic consumption.

(US) to 0.95 (Canada and euro area) in the pre-crisis period and from 0.43 (euro area) to 0.98 (US) in the crisis period.

Furthermore, the off-diagonal elements of the ARCH and GARCH matrices indicate that shocks to exchange rate changes (stock returns) affect the conditional variance of stock returns (exchange rate changes) in Japan and Switzerland across the two sub-periods, the UK in the pre-crisis period, and Canada and the US in the crisis period; the 5% critical value of the chi-squared distribution with 4 degrees of freedom is 9.49.

More specifically, the results of the likelihood ratio tests suggest the existence of causality-in-variance, that operates as an information flow, from exchange rate changes to stock returns in the UK, Japan, and Canada in the pre-crisis period. In the crisis period, there is evidence of causality-in-variance from stock returns to exchange rate changes in Japan, in the opposite direction in the euro area and Switzerland, and of bidirectional feedback in the US and Canada. Therefore these two types of financial markets appear to have become integrated in all countries, except the UK, during the recent financial crisis.

The Hosking multivariate *Q*-statistics of order (6) and (12) for the squared standardised residuals suggest, at the 5% significance level, that the multivariate GARCH (1, 1) structure adequately captures volatility, and hence no further variance dynamics are required.

3.7. Conclusions

In this chapter, we have analysed the nature of the linkages between stock market returns and exchange rate changes in six advanced economies, namely the US, the UK, Canada, Japan, the euro area, and Switzerland. Specifically, we have examined the extent to which they have been affected by the banking crisis of 2007–2010 employing weekly data from August 2003 to December 2011. The estimation of bivariate GARCH-BEKK models provides evidence of unidirectional causality from stock returns to exchange rate changes in the US and the UK, in the opposite direction in Canada, and of bidirectional causality in the euro area and Switzerland during the recent financial crisis. Furthermore, causality-in-variance tests for the crisis period lend support to the existence of causality-in-variance from stock returns to exchange rate changes in Japan, in the opposite direction in the euro area and Switzerland, and of bidirectional causality-in-variance in the US and Canada. In a broad sense, our findings are consistent with those of Tsagkanos and Siriopoulos (2013), who found that stock returns lead exchange rate changes in the short run during the recent financial crisis in the EU and the US, while the reverse short run causality holds during the preceding tranquil period. Our findings are also consistent with those of Granger et al. (2000) and Caporale et al. (2002), who examined the 1997 Asian financial crisis.

The results reflect cross-country differences in terms of policies, cycle phases, expectations, the degree of liberalisation, and capital controls (Nieh and Lee, 2001). Furthermore, given the fact that the currencies under investigation are the most actively traded (the corresponding economies being the top trading countries), their heterogeneous strength throughout the financial crisis may have played an important role in generating capital inflows and outflows. This might be one of the reasons for the different results when analysing the interactions between stock returns and exchange rate changes in these economies. Indeed, Milesi-Ferretti and Tille (2011) showed that capital flows were heterogeneous across countries and throughout the crisis period. Though, in addition to the strength of macroeconomic fundamentals and policies, Fratzscher (2012) found evidence that the institutional quality and country risk have also been the main forces of the heterogeneity of capital flows during the crisis.

Finally, our findings also imply the existence of limited diversification opportunities on a domestic basis during financial crises. Since stock prices and exchange rates have been shown to be interlinked strongly within national economies, it follows that investors cannot use them as effective instruments for portfolio hedging and diversification strategies. This applies to all countries examined except the UK.

Appendix A3

Table A3.1. Summary of descriptive statistics for stock returns and traded-weighted exchange rate changes.

Panel A. Pre-crisis period (August 6, 2003 – August 8, 2007)

Statistics	Variable	Canada	UK	Euro area	Japan	Switzerland	US
Mean	$R_{E,t}$	0.128	0.045	0.034	-0.044	-0.022	-0.078
	$R_{S,t}$	0.313	0.216	0.278	0.288	0.279	0.209
St. Dev	$R_{E,t}$	0.987	0.692	0.573	0.976	0.468	0.887
	$R_{S,t}$	1.567	1.456	1.805	2.307	1.638	1.327
Skewness	$R_{E,t}$	-0.152	-0.060	0.040	0.300	0.418	0.256
	$R_{S,t}$	-0.584	-0.861	-0.747	-0.441	-0.714	-0.412
Ex. kurtosis	$R_{E,t}$	2.875	2.977	4.335	3.370	3.513	3.009
	$R_{S,t}$	3.497	5.227	4.056	3.445	4.012	3.137
JB	$R_{E,t}$	0.942	0.133	15.59***	4.341	8.398***	2.284
	$R_{S,t}$	14.04***	69.06***	29.19***	8.527**	26.70***	6.098**
*LB(*10)	$R_{E,t}$	8.299	15.379	14.91	17.039*	14.04	3.454
	$R_{S,t}$	12.684	12.435	9.081	10.439	7.159	16.687*
LB^2(10)	$R_{E,t}$	14.113	6.917	29.125***	23.68***	23.87***	11.081
	$R_{S,t}$	24.28***	7.402	4.876	16.027*	8.148	5.054

Panel B. Crisis period (August 15, 2007 – December 28, 2011)

Statistics	Variable	Canada	UK	Euro area	Japan	Switzerland	US
Mean	$R_{E,t}$	0.0153	-0.110	-0.024	0.186	0.122	-0.047
	$R_{S,t}$	-0.046	-0.045	-0.271	-0.294	-0.168	-0.051
St. Dev	$R_{E,t}$	1.754	1.2755	0.995	1.655	1.344	1.160
	$R_{S,t}$	2.952	3.061	4.526	3.755	2.926	3.159
Skewness	$R_{E,t}$	-0.217	0.250	0.432	0.595	-1.634	-0.554
	$R_{S,t}$	-0.817	-0.753	-0.677	-0.834	-0.485	-0.936
Ex. kurtosis	$R_{E,t}$	4.184	6.072	6.587	4.808	19.833	6.274
	$R_{S,t}$	6.046	4.502	4.939	8.018	6.1306	6.997
JB	$R_{E,t}$	15.12***	92.04***	129.36***	44.53***	2793.4***	113.58***
	$R_{S,t}$	113.5***	43.01***	53.18***	265.7***	102.06***	185.10***
LB(10)	$R_{E,t}$	15.365	20.98**	17.594*	12.489	21.872**	16.777*
	$R_{S,t}$	34.12***	17.910*	27.242***	13.784	14.475	16.506*
LB^2(10)	$R_{E,t}$	115.3***	106.1***	38.49***	51.49***	36.63***	21.726**
	$R_{S,t}$	54.80***	35.51***	27.24***	31.01***	46.75***	32.21***

Notes: $R_{S,t}$ and $R_{E,t}$ indicate stock market returns and exchange rate changes, respectively. *LB*(*p*) and $LB^2(p)$ are the Ljung and Box (1978) tests for the p^{th}-order serial correlation of the returns, $R_{i,t}$, and squared returns, $R^2_{i,t}$, respectively, where $i = S$ (for stock returns), E (for exchange rate changes). *JB* is the Jarque-Bera test for normality.
***, **, and * indicate significance at the 1 %, 5%, and 10% levels, respectively.

Table A3.2. Augmented Dicky-Fuller unit root tests.

	Sample	Canada	Euro area	Japan	UK	Switzerland	US
e_t	Pre-crisis	-0.95(0)	-1.06(1)	0.38(6)	-2.16(4)	-1.688(0)	-2.683(0)
$R_{E,t}$		-15.4(0)***	-15.82(0)***	-7.89(5)***	-9.34(3)***	-15.97(0)***	-13.5(0)***
s_t		-0.97(0)	-0.73(1)	-0.83(2)	-0.96(2)	-0.757(1)	-1.156(1)
$R_{S,t}$		-14.9(0)***	-15.9(0)***	-0.83(1)***	-11.59(1)***	-15.95(0)***	-16.6(0)***
e_t	crisis	-1.59(2)	-1.25(0)	-1.03(1)	-2.65(3)	-0.724(6)	-1.860(6)
$R_{E,t}$		-12.9(1)***	-6.51(6)***	-16.66(0)***	-8.51(2)***	-8.104(5)***	-5.71(5)***
s_t		-2.26(8)	-1.97(7)	-2.576(0)	-0.43(0)	-2.661(1)	-1.596(0)
$R_{S,t}$		-3.92(7)***	-6.11(6)***	-7.76(2)***	-15.33(0)***	-16.45(0)***	-15.2(0)***

Notes: s_t ($R_{S,t}$) indicates the log of stock market price (stock market return), whereas e_t ($R_{E,t}$) denotes the log of exchange rates (exchange rate changes); the 1% and 5% critical values for ADF tests are -3.43 and -2.8748, respectively; the proper lag order is chosen by considering Akaike Information Criteria (AIC), representing in parenthesis.
*** indicates statistical significance at the 1% level.

Table A3.3. Phillips-Perron unit root tests.

	Sample	Canada	Euro area	Japan	UK	Switzerland	US
e_t	Pre-crisis	-0.909	-1.035	0.251	-2.283	-1.593	-2.720
$R_{E,t}$		-15.48***	-15.86***	-14.34***	-14.18***	-16.10***	-13.51***
s_t		-1.167	-0.864	-1.347	-1.079	-0.732	-1.307
$R_{S,t}$		-15.41***	-16.04***	-13.02***	-19.62***	-16.04***	-16.69***
e_t	crisis	-1.711	-1.483	-1.167	-2.465	-1.009	-2.030
$R_{E,t}$		-15.78***	-14.18***	-16.67***	-15.69***	-15.69***	-15.04***
s_t		-1.669	-1.667	-2.079	-1.789	-2.075	-1.561
$R_{S,t}$		-15.17***	-16.55***	-16.20***	-15.62***	-16.81***	-15.27***

Notes: s_t ($R_{S,t}$) indicates the log of stock market price (stock market return), whereas e_t ($R_{E,t}$) denotes the log of exchange rates (exchange rate changes); the 1% and 5% critical values for PP tests are -3.4587 and -2.8739, respectively.
*** indicates statistical significance at the 1% level.

Table A3.4. Lee and Strazicich' minimum LM unit root tests with one break.

Panel A. Pre-crisis period (6 August 2003- 8 August 2007)

Series		US	Canada	Euro area	Japan	UK	Switzerland
s_t	Stat	-3.551(0)	-3.575(3)	-3.860(0)	-3.426(2)	-4.332(0)	-3.138(0)
	T_B	[2006:06:14]	[2006:06:07]	[2004:07:21]	[2005:10:19]	[2005:10:26]	[2005:08:31]
	λ	λ=0.7	λ=0.7	λ=0.24	λ=0.55	λ=0.55	λ=0.52
$R_{S,t}$	Stat	-16.22(2)***	-15.10(0)***	-15.72(0)***	-12.90 (0)***	-15.37(0)***	-16.08(0)***
	T_B	[2004:05:19]	[2004:05:26]	[2004:10:27]	[2004:12:22]	[2004:08:25]^{n}	[2004:08:18]
	λ	λ=0.2	λ=0.2	λ=0.3	λ=0.35	λ=0.27	λ=0.26
e_t	Stat	-3.598(1)	-2.916(0)	-3.565(3)	-3.171(6)	-2.956(4)	-4.092(0)
	T_B	[2005:06:01]	[2006:10:25]^{n}	[2005:08:10]^{n}	[2005:07:20]	[2005:07:20]	[2004:11:03]
	λ	λ=0.45	λ=0.8	λ=0.5	λ=0.5	λ=0.5	λ=0.3
$R_{E,t}$	Stat	-13.74(0)***	-15.22(0)***	-15.82(0)***	-10.00(5)***	-13.04(0)***	-15.38(0)***
	T_B	[2005:10:05]	[2007:02:14]	[2005:12:07]	[2004:03:03]	[2006:07:12]	[2004:07:21]
	λ	λ=0.545	λ=0.9	λ=0.6	λ=0.15	λ=0.73	λ=0.24

Panel B. Crisis period (15 August 2007- 28 December 2011)

Series		US	Canada	Euro area	Japan	UK	Switzerland
s_t	Stat	-2.698(0)	-2.666(6)	-2.868(6)	-3.333(3)	-2.880(5)	-2.856(0)
	T_B	[2008:11:05]	[2009:07:08]	[2009:07:08]	[2009:04:29]	[2009:07:08]	[2009:05:13]
	λ	λ=0.3	λ=0.4	λ=0.4	λ=0.4	λ=0.4	λ=0.4
$R_{S,t}$	Stat	-14.76(0)***	-6.147(6)***	-6.933(6)***	-16.19(0)***	-15.40(0)***	-16.81(0)***
	T_B	[2008:08:27]	[2009:01:21]	[2008:12:24]	[2009:02:18]	[2008:11:19]	[2011:05:11]^{n}
	λ	λ=0.2	λ=0.3	λ=0.3	λ=0.3	λ=0.3	λ=0.9
e_t	Stat	-2.661(6)	-2.196(2)	-3.405(3)	-3.450(0)	-4.334(0)	-3.765(5)
	T_B	[2008:11:05]	[2008:10:15]	[2008:02:27]	[2008:10:15]	[2008:11:26]	[2010:05:26]
	λ	λ=0.3	λ=0.3	λ=0.1	λ=0.3	λ=0.3	λ=0.6
$R_{E,t}$	Stat	-6.032(5)***	-13.483(0)***	-7.104(6)***	-16.71(0)***	-12.42(1)***	-8.592(5)***
	T_B	[2010:06:16]	[2008:12:31]	[2010:10:27]	[2008:12:03]	[2009:12:02]	[2011:06:22]^{n}
	λ	λ=0.6	λ=0.3	λ=0.7	λ=0.3	λ=0.5	λ=0.9

Notes: s_t ($R_{S,t}$) indicates the log of stock market price (stock market return), whereas e_t ($R_{E,t}$) denotes the log of exchange rates (exchange rate changes); the lag length is selected on the basis of the general-to-specific approach, represented in parentheses (.), allowing for a maximum of 6 lags; and the estimated breakpoints (T_B) are reported in square brackets [.] with *n* indicates that the identified break point is not significant at the 10% level. The critical values for the minimum single break LM unit root test allowing for a shift in the intercept and a change in the trend slope (Model C), which depend (to some extent) on the location of the breakpoint ($\lambda=T_B/T$, where T is sample size) and are symmetric around λ and (1- λ), are displayed below.

*** indicates statistical significance at the 1% level.

The critical values displayed below come from Lee and Strazicich (2004).

Break point	Critical values	
$\lambda= (T_B/T)$	1%	5%
0.1	-5.11	-4.50
0.2	-5.07	-4.47
0.3	-5.15	-4.45
0.4	–5.05	-4.50
0.5	-5.11	-4.51

Table A3.5. Results of the Engle and Granger (1987) cointegration tests based on residuals.

Countries	Sample	s_t on e_t	e_t on s_t
US	Pre-crisis	-1.718(0)	-2.838(0)
	Crisis	-2.712(0)	-2.978(0)
UK	Pre-crisis	-1.138(5)	-2.306(4)
	Crisis	-1.809(0)	-2.246(0)
Euro area	Pre-crisis	-1.627(5)	-2.696(0)
	Crisis	-2.155(0)	-1.485(0)
Canada	Pre-crisis	-2.568(0)	-2.550(0)
	Crisis	-3.003(4)	-1.982(8)
Japan	Pre-crisis	-5.149(0)***	-4.789(0)***
	Crisis	-2.218(0)	-1.526(0)
Switzerland	Pre-crisis	-2.201(1)	-2.640(0)
	Crisis	-1.899(0)	-0.694(0)

Notes: The pairwise Engle and Granger (1987) test is conducted by regressing the log of stock market price (s_t) on the log of the exchange rate (e_t), as well as in the other way around. The proper lag order is chosen by considering the Akaike information Criterion (AIC), representing in parenthesis; the 1% and 5% critical values of the MacKinnon (1991) for the augmented Dickey-Fuller test statistics are -3.89644 and -3.33613, respectively.
*** denotes statistical significance at the 1% level.

CHAPTER FOUR

EXCHANGE RATES AND NET PORTFOLIO FLOWS: A MARKOV-SWITCHING APPROACH

4.1. Introduction

Financial markets deregulation has led to a dramatic increase in international capital mobility across most developed economies. To give an example of the magnitude of such a shift, gross cross-border portfolio investments in equities and bonds in the US were accounting for only 4% of GDP in 1975, this proportion surged to 100% in the early of 1990s and has reached 245% by 2000 (Hau and Rey, 2006). Global capital flows, by contrast, had steadily increased from about 2% of world GDP in 1975 to over 20% in 2007; however, this proportion declined sharply on the onset of the Lehman Brothers collapse in late 2008 before starting to increase again in 2009 (see Milesi-Ferretti and Tille, 2011). Not surprisingly the recent literature has focused on the impact (causal effect) that the increased capital mobility across-borders has had on exchange rate dynamics, especially in the light of the poor performance of the macroeconomic models to explain such dynamics (see Meese and Rogoff, 1983; Cheung et al., 2005).

Recent works have shown that the microstructure dynamics of exchange rates, currency order flows, explain a significant proportion of exchange rate variations (e.g., Evans and Lyons, 2002; 2005; 2008; Payne, 2003; Rime et al., 2010; Chinn and Moore, 2011; and Duffuor et al., 2012; among others). Furthermore, Hau and Rey (2006), in an influential study, argued that portfolio flows and order flows are closely aligned since both flows are driven by investors' behaviour; hence the former flows are also likely to contain information about exchange rate changes. Indeed, the ongoing empirical literature has produced convincing evidence that portfolio investment flows do affect the dynamics of exchange rates to a large extent using different data sets (e.g., Brooks et al., 2004; Hau and Rey, 2006; and Kodongo and Ojah, 2012; among others). However, most empirical studies on the effects of portfolio inflows on exchange rate changes have assumed linear dependence, using mainly ordinary least squares (OLS) and VAR models.

This chapter contributes to the existing literature by using a Markov-switching framework to propose an alternative way of measuring the relationship between net equity and net bond portfolio investment flows and exchange rate changes using bilateral quarterly data from the US *vis-à-vis* Canada, the euro area, Japan, and the UK over the period 1990:01-2011:04. To the best of our knowledge, the nonlinear dependence between both variables has not been explored in the literature yet, and this chapter aims to fill this gap.

We argue that the linear analysis between portfolio flows and exchange rate changes used in the literature is quite limited and does not allow capturing the dynamics observed in the exchange rates over the last few decades.[48] Indeed, the swings of exchange rates and volatility regimes have been widely known by now. For example, the theoretical study by Jeanne and Rose (2002) showed that the exchange rates of two countries under floating exchange rate regimes may have different levels of volatility even though they have analogues macroeconomic fundamentals. The multiple equilibria

[48] The empirical studies on the linkage between portfolio flows and exchange rate changes have used linear techniques. However, these studies have primarily two limitations. First, the nonlinearity in the relationship between the two variables is yet to be explored in the literature and this chapter aims to fill this gap. Second, the empirical studies do not address the heteroscedasticity in the variance associated with the flows and exchange rate changes for which **Chapter 5** aims to fill this gap.

of the exchange rate and its volatility have also been associated with the currency crises and the self-fulfilling view of such crises (see, for examples, Jeanne and Masson, 2000; Chen, 2006, among others). More recently, Lovcha and Perez-Laborda (2013) further pointed out that investors' reaction should not be the same in different situations of the market in analysing the relationship between exchange rate changes and customer order flows and confirmed that such non-linearity can be well captured by the Markov switching models.

The nonlinear model employed in this chapter separates periods of high and low states of the world for the mean and the variance of the endogenous variable (exchange rate changes). Therefore it allows us to estimate separate causal effects for periods of exchange rate appreciation and depreciation, and periods of high and low exchange rate volatility. As shown by Engle and Hamilton (1990), Bekaert and Hodrick (1993), Kaminsky (1993), Engel (1994), Caporale and Spagnolo (2004), Frömmel et al. (2005), Bazdresch and Werner (2005), Chen (2006), Brunetti et al. (2008), and Lovcha and Perez-Laborda (2013) among others, the Markov regime-switching model is particularly appropriate to model exchange rate dynamics.

The chapter is organised as follows. Section 4.2 reviews the empirical literature on the relationship between portfolio flows and exchange rate changes; Section 4.3 describes the employed econometric model and outlines the hypotheses tested in the study. Section 4.4 provides details on the data set. Section 4.5 discusses the empirical results; and Section 4.6 offers some concluding remarks.

4.2. A review of the literature

The impact of equity and bond portfolio flows on exchange rate changes has been the matter of much attention over the recent years. Even though the empirical studies have focused on the linear dependence between the two variables using mainly

OLS and VAR models, the findings of these studies are inconclusive. For example, using quarterly data from 1980 to 1997 and based on Granger causality tests, Edwards (1998) observed the importance of capital flows as the leader or the existence of bidirectional causality between capital flows and real exchange rates in examining seven Latin American countries, namely Argentina, Brazil, Chile, Columbia, Mexico, Peru, and Venezuela. Froot et al. (2001), using daily data on cross-border flows for 44 countries, found a positive contemporaneous correlation between net portfolio investment inflows and both equity, expressed in US dollar, and currency returns, as well as a strong positive correlation between net portfolio inflows and lagged equity and currency returns.

Brooks et al. (2004), using quarterly data over the period 1988:01-2000:03, examined the impact of portfolio and foreign direct investment flows on the yen-dollar and the euro-dollar exchange rates. While the yen-dollar exchange rate was found to be driven by mainstream macroeconomic variables such as the interest rate differential and the current account, the euro-dollar exchange rate was shown to be driven primarily by bilateral net portfolio investment flows. More specifically, equity inflows from the euro area towards the US implied a depreciation of the euro against the US dollar.

Siourounis (2004), using monthly data for the US exchange rate *vis-à-vis* the exchange rates of the UK, Japan, Germany, and Switzerland over the period 1988-2000, provided evidence that equity flows rather than bond flows are tracing the evolution of exchange rates. The study was conducted using an unrestricted Vector Autoregressive (VAR) model in which net cross-border capital flows, equity return differentials, exchange rate changes, and interest rate differentials were set endogenously.

By employing an equilibrium framework in which exchange rate changes, stock returns, and capital flows are jointly set under incomplete foreign exchange risk trading feature, Hau and Rey (2006) showed that the correlation between equity flows and exchange rate changes is significant in 6 out of 17 OECD countries considered. When pooling the data across the countries, the correlation became highly significant.

Chaban (2009), using data from Canada, Australia, and New Zealand, argued that Hau and Rey' results were not specifically supported in commodity-exporting countries. While Chaban argued that commodity prices play a significant role in the transmission of shocks in these countries, Ferreira Filipe (2012) showed that differences in country-specific shocks volatility also play a role in these countries and should therefore be accounted for. However, examining the impact of various types of capital flows (foreign direct investment, portfolio investment, other capital flows) relative to traditional macroeconomic fundamentals on the dynamics of the Australian dollar, the Canadian dollar, and the US dollar exchange rates, Sun and An (2011) showed that, among macroeconomic variables, interest rate differential plays a notable role in the movement of the three exchange rates in question, by using Structural Vector Autoregressive (SVAR) model and quarterly data over the period 1980-2004. Furthermore, capital flows, primarily portfolio investment flows, were shown to explain the movements of the Australian dollar and the Canadian dollar, but not the US dollar exchange rate.

By using quarterly data from Mexico over the period 1988:01-2008:02, Ibarra (2011) showed that various types of capital inflows, including portfolio investment and foreign direct investment inflows, towards Mexico result in an appreciation of the real Mexican peso. This finding was obtained using the Autoregressive Distributed Lag (ARDL) bounds approach while controlling for a set of macroeconomic fundamentals (relative industrial production between Mexico and the US, relative government consumption as a percentage of GDP between both countries, and the world price of oil). Recently, Combes et al. (2012), employing a pooled mean group estimator for a sample of 42 emerging and developing economies over the period 1980-2006, also showed that public and private inflows, primarily portfolio investment inflows, result in an appreciation of the real effective exchange rate.

Kodongo and Ojah (2012), using monthly data over the period 1997:1-2009:12 for four African countries, namely Egypt, Morocco, Nigeria, and South Africa, showed that the dynamic relationship between real exchange rate changes and international portfolio flows is country-specific and time-dependent. For instance, the results of the full sample period showed the existence of no evidence of the dynamic interrelationship between real exchange rates and net portfolio inflows in Egypt, Morocco, and Nigeria. However, bidirectional causality between both variables is found in the case of South Africa. By considering their first sub-period (1997:01-2003:12), on the other hand, it is found that net portfolio inflows lead real exchange rates in Egypt, real exchange rates take the lead of net portfolio inflows in South Africa, whereas Nigeria and Morocco showed evidence of no causality in either way. Finally, the results of the second sub-period (2004:01-2009:12) revealed that causality is running from real exchange rates to net portfolio inflows in Morocco, causality in the opposite direction is found in Nigeria, whilst evidence of no causality is detected in Egypt and South Africa during the period.

4.3. The econometric model

We propose an alternative way of detecting the causal dynamics between net equity and net bond portfolio flows and exchange rate changes for the US *vis-à-vis* Canada, the euro area, Japan, and the UK. The regime-switching model considered in this chapter[49] allows for shifts in the mean, for periods of currency appreciation and depreciation, and in the variance, for periods of high volatility and low volatility; and is

[49] The model is based on the Markov switching representation proposed by Hamilton (1989; 1990).

given by:

$$r_t = \mu(s_t) + \sum_{i=1}^{4} \phi_i r_{t-i} + \alpha(s_t) nbf_{t-1} + \beta(s_t) nef_{t-1} + \sigma(s_t)\varepsilon_t, \qquad \varepsilon_t \sim N(0,1) \qquad (4.1)$$

where r_t = exchange rate changes, nbf_{t-1} = net bond flows, nef_{t-1} = net equity flows. Given that s_t is unobserved, estimation of Eq. (4.1) requires restrictions on the probability process governing s_t; it is assumed that s_t follows a first-order, homogeneous, two-state Markov chain. This means that any persistence in the state is completely summarised by the value of the state in the previous period.

Therefore, the regime indicators $\{s_t\}$ are assumed to form a Markov chain on $\mathbb{S}$ with transition probability matrix $\mathbf{P}' = [p_{ij}]_{2\times 2}$, where:

$$p_{ij} = \Pr(s_t = j \mid s_{t-1} = i), \quad i, j \in \mathbb{S}, \qquad (4.2)$$

and $p_{i1} = 1 - p_{i2}$ $(i \in \mathbb{S})$, where each column sums to unity and all elements are non-negative. The probability law that governs these regime changes is flexible enough to allow for a wide variety of different shifts, depending on the values of the transition probabilities. For example, the values of p_{ii} $(i \in \mathbb{S})$ that are not very close to unity imply that structural parameters are subject to frequent changes, whereas values near unity suggest that only a few regime transitions are likely to occur in a relatively short realization of the process. $\{\varepsilon_t\}$ are *i.i.d.* errors with $E(\varepsilon_t) = 0$ and $E(\varepsilon_t^2) = 1$. $\{s_t\}$ are random variables in $\mathbb{S} = \{1, 2\}$ that indicate the unobserved state[50] of the system at time t

[50] Regime 1 is labelled as the low regime, whereas regime 2 as the high regime.

. It is assumed that $\{\varepsilon_t\}$ and $\{s_t\}$ are independent. Also, note that the independence between the sequences $\{\varepsilon_t\}$ and $\{s_t\}$ implies that regime changes take place independently of the past history of $\{r_t\}$.

We are interested in documenting estimates of the low-high phase exchange rate changes, μ^l and μ^h, but mainly in investigating the extent to which net bond flows and net equity flows are associated with the low-high phase exchange rate changes. Autoregressive terms (up to four lags) are also sequentially included in the model to remove serial correlations; that is, to capture the adequate dynamics of exchange rate changes. Therefore, the parameter vector of the mean equation, Eq. (4.1), is defined by $\mu^{(i)}$ $(i = low, high)$ and $\sigma^{(i)}$ $(i = low, high)$, which are real constants, and the autoregressive terms, $\sum_{i=1}^{4}\phi_i$, up to four lags. $\alpha = (\alpha^l, \alpha^h)$ and $\beta = (\beta^l, \beta^h)$ measure the impact of net bond flows and net equity flows respectively on exchange rate changes. The parameter vector is estimated by maximum likelihood. The density of the data has two components, one for each regime, and the log-likelihood function is constructed as a probability weighted sum of these two components. The maximum likelihood estimation is performed using the EM algorithm described by Hamilton (1989; 1990).

Furthermore, we estimate the linear model commonly used in the literature and take it as a benchmark. This is given by:

$$r_t = \mu + \sum_{i=1}^{4}\phi_i r_{t-i} + \alpha nbf_{t-1} + \beta nef_{t-1} + \sigma\varepsilon_t, \qquad \varepsilon_t \sim N(0,1) \tag{4.3}$$

where the parameter vector of the mean equation, Eq. (4.1), is defined by the constant parameters $(\mu, \phi_i, \alpha, \beta, \sigma)$.

4.4. Data

The variables employed in this chapter are net bond flows (nbf_t), net equity flows (nef_t), and exchange rates (E_t) for the US *vis-à-vis* Canada, the euro area, Japan, and the UK. The currencies of these countries are the most tradable ones in the foreign exchange market. Hence, this study provides a good opportunity to examine the extent to what the US dollar exchange rate against the currencies of the four countries in question are driven by the corresponding bilateral net equity and net bond portfolio flows. Throughout, the US is considered the domestic economy. We use quarterly data from 1990:01 to 2011:04.[51] We employ quarterly data because regime shifts can be detected evidently in low frequency data (for an example, see Engle, 1994). Data on exchange rates are from the IMF's *International Financial Statistics (IFS)*, whereas portfolio investment flows are sourced from the USA *Treasury International Capital (TIC) System*.[52]

As Edison and Warnock (2008) pointed out the US TIC data have three main limitations. First, the data only cover transactions that involve the US residents. That is, these data are bilateral portfolio inflows and outflows of the US and do not include other cross countries portfolio investments. Second, transactions that take place via third countries lead to a financial centre bias in the bilateral flows data as the transaction is recorded against the foreign intermediary rather than where the issuer of the foreign security resides. Third, financing of cross-border mergers through stock swaps makes the analysis of equity flows rather difficult.[53] However, in spite of these limitations, the

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[51] Even though the chosen sample is small, the Markov switching models are likely to separate the two regimes in terms of appreciation/depreciation and low/high volatility because of the associated swings of the major currencies under observation over the last two decades.

[52] The data are obtained from the US Treasury Department website: http://www.treasury.gov/resource-center/data-chart-center/tic/Pages/country-longterm.aspx

[53] For further details of the US TIC data, the reader is directed to Edison and Warnock (2008).

TIC data have been widely used in the empirical literature as they are likely to contain information about bilateral portfolio investments between the US and the rest of the world.

Exchange rate changes are calculated as $r_t = 100 \times (E_t / E_{t-1})$. For the euro area exchange rate, the ECU is considered prior to the introduction of the euro in 1999. Net portfolio flows, on the other hand, are constructed as the difference between portfolio *inflows* and *outflows*. *Inflows* and *outflows* are measured as net purchases and sales of domestic assets (equities and bonds) by foreign residents, and net purchases and sales of foreign assets (equities and bonds) by domestic investors, respectively. The euro area portfolio flows are calculated aggregating the data for the individual EMU countries (Austria, Belgium-Luxemburg, Finland, France, Germany, Ireland, Italy, the Netherlands, Portugal, and Spain) (Heimonen, 2009). Furthermore, following Brennan and Cao (1997), Hau and Rey (2006), and Chaban (2009), flows were normalised by the average of their absolute values over the previous four quarters.

Summary statistics for the variables considered are displayed in Table 4.1. The mean quarterly changes in exchange rates are positive (US dollar depreciation) for Japan and Canada, while they are negative (US dollar appreciation) for the UK and the euro area. The averages of net bond flows and those of net equity flows, on the other hand, are positive in all cases, but for net equity flows in Japan, thereby indicating the existence of net bond and net equity inflows from all countries towards the US except Japan, where there exist net equity outflows from the US towards Japan. Exchange rate changes, as expected, exhibit higher volatility than net portfolio (bond and equity) flows. Furthermore, skewness and excess kurtosis characterise all the series.

Table 4.1. Summary of descriptive statistics.

	Mean	St. Dev	Skewness	Ex. Kurtosis	JB
Canada					
r_t	0.102	2.766	–0.736	8.1592	104.34***
nbf_t	0.355	1.897	3.358	23.535	1692.2***
nef_t	0.192	1.653	0.762	6.2477	46.660***
Euro area					
r_t	–0.035	4.773	–0.509	2.935443	3.7741
nbf_t	0.303	1.523	–0.072	4.1495	4.8655*
nef_t	0.032	1.547	0.128	3.40938	0.8451
Japan					
r_t	0.902	4.953	0.501	3.613	5.015*
nbf_t	0.972	1.307	–0.185	3.330	0.894
nef_t	–0.403	1.959	–1.186	9.737	184.9***
UK					
r_t	–0.282	4.812	–1.625	8.7961	160.0***
nbf_t	0.873	1.450	–2.685	17.829	901.6***
nef_t	0.0001	1.781	–0.141	5.8607	29.95***

Notes: r_t, nbf_t, and nef_t indicate exchange rate changes, net bond flows, and net equity flows, respectively; *JB* is the Jarque-Bera test for normality. *** and * indicate significance at the 1% and 10% levels, respectively.

The exceptions are net bond flows in Japan and variables of the euro area such as net equity flows, net bond flows (which exhibit only excess kurtosis) and exchange rate changes (which exhibit only skewness). Finally, the Jarque-Bera (JB) test statistics show evidence of no normality in all series but net bond flows in Japan and net equity flows and exchange rate changes in the euro area.

4.5. Empirical results

The first step of our analysis is to estimate the benchmark model, Eq. (4.3), by the standard OLS. Results, reported in Table 4.2, indicate that neither net equity flows, nor net bond flows have a statistically significant effect on exchange rate changes for all countries. These findings are consistent with those of Hau and Rey (2006) and Chaban (2009), but contradict those of Brooks et al. (2004), who found a statistically significant linkage in the euro area.

The null hypothesis of linearity against the alternative of a Markov regime switching cannot be tested directly using a standard likelihood ratio (LR) test. We properly test for multiple equilibria (more than one regime) against linearity using the Hansen (1992b)'s standardised likelihood ratio test. The values of the standardised likelihood ratio statistics and related *p*-values (Table 4.3) under the null hypothesis[54] (Hansen, 1992b) provide strong evidence in favour of a two-state Markov switching specification. This procedure requires the evaluation of the likelihood function across a grid of different values for the transition probabilities and for each state-dependent parameter. We also test for the presence of a third state (Table 4.3). The results provide strong evidence in favour of a two-state regime-switching specification.

Maximum likelihood (ML) estimates are reported in Tables 4.4 to 4.7. We estimate four nested Markov switching models. Model IV is the general model and allows for shifts in the mean, $\mu(s_t)$, the variance, $\sigma(s_t)$, bond, $\alpha(s_t)$, and equity, $\beta(s_t)$, flows parameters; Model III constraints the mean to be not regime dependent, whereas Model II constraints the variance to be constant. Finally Model I allows for switches in the bond and equity flows parameters only.

[54] The *p*-value is calculated according to the method described in Hansen (1996), using 1,000 random draws from the relevant limiting Gaussian processes and bandwidth parameter M = 0, 1, ..., 4. See Hansen (1992b) for details.

Table 4.2. Parameter estimates for linear models.

	Japan	Canada	UK	Euro area
μ_1	0.523 (0.635)	−0.014 (0.343)	−0.209 (0.583)	−0.0723 (0.484)
α_1	0.003 (0.437)	0.232 (0.177)	0.023 (0.426)	0.126 (0.323)
β_1	0.119 (0.235)	0.033 (0.210)	−0.042 (0.242)	0.329 (0.313)
ϕ_1	0.238^{**} (0.108)	0.384^{*} (0.110)	0.343^{*} (0.105)	0.246^{**} (0.109)
ϕ_2	-0.309^{*} (0.101)	-0.235^{**} (0.111)	-0.310^{*} (0.106)	-0.206^{**} (0.111)
ϕ_3	0.332^{*} (0.112)	−	−	−
σ_1	4.265	3.024	4.194	4.369
Log Likelihood	−237.933	−212.093	−242.764	−240.485
$LB_{(8)}$	0.685 [0.702]	5.622 [0.466]	6.222 [0.398]	8.143 [0.227]
$LB^2_{(8)}$	0.466 [0.875]	0.479 [0.866]	0.157 [0.995]	0.338 [0.947]
JB	19.68 [0.000]	31.34 [0.000]	33.47 [0.000]	3.470 [0.176]

Notes: Standard errors and *p*-values are reported in (.) and [.], respectively. $LB_{(8)}$ and $LB^2_{(8)}$ are respectively the Ljung and Box (1978) test of the significance of autocorrelations of eight lags in the standardised and standardised squared residuals. *JB* is the Jarque-Bera test for normality.
***, ** and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Following Psaradakis and Spagnolo (2003), we use the Akaike Information Criterion (AIC) to identify the best fitted model among the four estimated models for each case. The models selected are Model II (shift in the mean) for Canada and the UK, and Model III (shift in the variance) for the euro area and Japan.[55] The selected models appear to be well defined: the standardised residuals exhibit no signs of linear or nonlinear dependence.[56] The periods of high and low exchange rate changes (Canada

[55] We use the AIC to choose the best fitted model among the candidate models considered. That is, to choose the best fitted model in regard of switching in the parameters. For example, in the cases of US/Canada and US/UK, Model II is favoured according to the AIC. This implies that in both cases the switching in exchange rate changes is driven primarily by the mean, but not the variance.

[56] The lags of exchange rate changes are included sequentially to remove serial correlations in the models. That is, to capture the adequate dynamics of exchange rate changes, where lags found insignificant are excluded. For example, when only the second lag r_{t-2} is included in the model, it implies that two lags were adequate to remove serial

and UK) and of high and low exchange rate volatility (euro area and Japan) seem to be accurately identified by the filtered probabilities. Figures 4.1 to 4.4 show the plots of exchange rate changes, r_t, along with the corresponding estimated filtered probabilities for Canada, the euro area, Japan, and the UK, respectively.

More specifically, for Canada the mean value of -1.3 (1.9) in regime low (high) indicates a regime characterised by US dollar appreciation (depreciation) against the Canadian dollar. The probability of staying in regime low (high) is 0.94 (0.92) (see Model II in Table 4.4). The filtered probabilities (see Figure 4.1) show a relatively low number of switches, consistent with the high regime persistency. There are 52 quarters (61.18%) where the process is in the low regime and 33 quarters (38.82%) where the process is in the high regime. However, net equity flows and net bond flows are found to be insignificant in either regime. These findings are in line with previous studies; see Chaban (2009), and Ferreira Filipe (2012).

In the case of the UK (see Model II in Table 4.7), while the mean value associated with the low regime is insignificant, the statistically significant mean, 0.84, in the high regime indicates US dollar depreciation against the British pound. The probability of staying in a low (high) regime is 0.28 (0.96), with the number of regime switches, being quite frequent when the process is in the low state. There are 4 quarters (4.60% of the total) where the process is in the low regime and 83 quarters (95.40%) where the process is in the high regime (see Figure 4.4). Net equity flows and net bond flows are found to have a significant impact on exchange rate changes in the low (US dollar appreciation) regime. These results suggest that both net equity and net bond

correlation in the model and the first lag was excluded because it was insignificant. Furthermore, models with no lagged exchange rate changes included imply that exchange rate changes in these models have no dynamics to be captured. Models with only one lag of exchange rate changes indicate that only one lag was enough to remove serial correlation and capture the dynamics in these models.

Table 4.3. Markov switching state dimension: Hansen test.

Country	Standardised LR test	Linearity vs two-states	Two states vs three-states
Canada	LR	4.123	0.2731
	$M = 0$	(0.001)	(0.643)
	$M = 1$	(0.002)	(0.677)
	$M = 2$	(0.003)	(0.690)
	$M = 3$	(0.009)	(0.715)
	$M = 4$	(0.011)	(0.722)
Euro area	LR	4.021	0.2567
	$M = 0$	(0.001)	(0.612)
	$M = 1$	(0.002)	(0.654)
	$M = 2$	(0.004)	(0.686)
	$M = 3$	(0.008)	(0.701)
	$M = 4$	(0.010)	(0.715)
Japan	LR	3.985	0.2451
	$M = 0$	(0.001)	(0.599)
	$M = 1$	(0.003)	(0.611)
	$M = 2$	(0.004)	(0.671)
	$M = 3$	(0.007)	(0.689)
	0.474^{***} (0.078)	(0.012)	(0.699)
UK	LR	3.769	0.2113
	$M = 0$	(0.001)	(0.587)
	$M = 1$	(0.002)	(0.599)
	$M = 2$	(0.003)	(0.645)
	$M = 3$	(0.006)	(0.661)
	$M = 4$	(0.013)	(0.688)

Notes: The Hansen (1992b)'s standardised Likelihood Ratio (LR) test *p*-values, reported in brackets (.), are calculated according to the method described in Hansen (1992b), using 1,000 random draws from the relevant limiting Gaussian processes and bandwidth parameter M = 0, 1,…, 4. Test results for the presence of a third state are also reported.

Table 4.4. Parameter estimates for regime-switching models: Canada.

	Model I	Model II	Model III	Model IV
μ_1	-0.024 (0.288)	-1.326^{***} (0.456)	-0.242 (0.246)	2.202 (1.411)
μ_2	–	1.981^{**} (0.855)	–	-0.341 (0.240)
α_1	0.101 (0.127)	0.214 (0.171)	0.908 (0.778)	-0.254 (0.905)
α_2	5.394^{***} (1.832)	0.455 (0.559)	0.091 (0.113)	0.118 (0.113)
β_1	0.313^{*} (0.166)	-0.147 (0.276)	-0.678 (0.911)	-1.771^{*} (0.944)
β_2	-2.108^{**} (1.019)	-0.045 (0.312)	0.267^{*} (0.143)	0.278^{**} (0.138)
ϕ_1	–	–	0.235^{**} (0.112)	0.216^{**} (0.105)
ϕ_2	–	-0.267^{***} (0.111)	–	–
σ_1	2.328^{***} (0.231)	2.705^{***} (0.222)	4.667^{***} (0.786)	4.420^{***} (0.714)
σ_2	–	–	1.778^{***} (0.175)	1.781^{***} (0.165)
P	0.942^{***} (0.038)	0.946^{***} (0.040)	0.887^{***} (0.101)	0.874^{***} (0.088)
$1-Q$	0.584^{**} (0.235)	0.077 (0.065)	0.043 (0.033)	0.041 (0.029)
Log Likelihood	-211.096	-214.209	-203.051	-201.650
AIC	438.192	448.418	426.102	425.3
$LB_{(8)}$	9.061 [0.337]	22.811 [0.088]	11.364 [0.181]	8.542 [0.382]
$LB^2_{(8)}$	0.897 [0.524]	0.498 [0.852]	1.4344 [0.201]	1.073 [0.394]
JB	3.379 [0.184]	31.595 [0.000]	2.9871 [0.224]	3.411 [0.181]

Notes: See notes of Table 4.2.

Table 4.5. Parameter estimates for regime-switching models: Euro area.

	Model I	Model II	Model III	Model IV
μ_1	-0.173 (0.508)	-4.417^{***} (0.453)	-0.069 (0.464)	-3.346^{***} (0.944)
μ_2	$-$	3.393^{***} (0.474)	$-$	3.296^{***} (0.673)
α_1	-1.126^{*} (0.584)	0.165 (0.349)	-1.662^{***} (0.427)	0.405 (0.540)
α_2	1.106^{**} (0.469)	0.173 (0.215)	1.070^{**} (0.478)	0.250 (0.244)
β_1	0.147 (0.437)	-0.335 (0.276)	0.095 (0.451)	-0.003 (0.398)
β_2	0.380 (0.502)	0.145 (0.292)	0.359 (0.479)	0.158 (0.373)
ϕ_1	0.195^{*} (0.105)	$-$	$-$	$-$
ϕ_2	-0.321^{***} (0.110)	-0.431^{***} (0.086)	-0.225^{***} (0.109)	-0.260^{**} (0.102)
ϕ_3	$-$	0.253^{***} (0.077)	$-$	$-$
ϕ_4	-0.150^{*} 0.081	-0.185^{**} (0.072)	$-$	$-$
σ_1	3.893^{***} (0.323)	2.355^{***} (0.218)	2.891^{***} (0.462)	3.345^{***} (0.597)
σ_2	$-$	$-$	4.418^{***} (0.475)	2.468^{***} (0.404)
P	0.943^{**} (0.068)	0.699^{***} (0.085)	0.882^{***} (0.097)	0.757^{***} (0.106)
$1-Q$	0.078 (0.063)	0.236^{***} (0.067)	0.081 (0.062)	0.249^{***} (0.089)
Log Likelihood	-235.708	-229.687	-242.444	-239.090
AIC	493.416	483.374	504.888	500.18
$LB_{(8)}$	5.282 [0.382]	6.892 [0.228]	5.465 [0.706]	3.246 (0.918)
$LB^2_{(8)}$	0.862 [0.553]	0.758 0.640]	0.408 [0.911]	0.775 [0.626]
JB	6.827 [0.032]	8.251 (0.016)	5.321 [0.069]	4.548 [0.102]

Notes: See notes of Table 4.2.

Table 4.6. Parameter estimates for regime-switching models: Japan.

	Model I	Model II	Model III	Model IV
μ_1	0.902 (0.621)	-3.663^{***} (0.891)	0.819 (0.646)	-3.688^{***} (0.828)
μ_2	–	3.679^{***} (0.621)	–	3.627^{***} (0.644)
α_1	–1.378 (1.013)	–0.001 (0.595)	-2.158^{***} (0.652)	–0.002 (0.553)
α_2	0.898^{*} (0.546)	–0.140 (0.382)	0.736 (0.537)	–0.136 (0.395)
β_1	0.894^{**} (0.439)	0.137 (0.371)	0.544 (0.436)	0.133 (0.341)
β_2	–0.220 (0.341)	0.005 (0.199)	0.028 (0.280)	0.007 (0.206)
ϕ_2	-0.280^{***} (0.104)	-0.519^{***} (0.080)	-0.296^{***} (0.114)	-0.503^{***} (0.083)
σ_1	4.047^{***} (0.363)	3.125^{***} (0.255)	2.669^{***} (0.721)	2.880^{***} (0.422)
σ_2	–	–	4.496^{***} (0.447)	3.237^{***} (0.322)
P	0.828^{***} (0.128)	0.885^{***} (0.062)	0.815^{***} (0.117)	0.883^{***} (0.063)
$1-Q$	0.106 (0.069)	0.056^{*} (0.031)	0.057 (0.042)	0.056^{*} (0.031)
Log Likelihood	–246.704	–234.806	–246.208	–234.600
AIC	511.408	489.612	512.416	491.2
$LB_{(8)}$	3.425 [0.843]	10.555 [0.159]	4.105 [0.847]	5.156 [0.740]
$LB^2_{(8)}$	0.588 [0.783]	0.527 [0.831]	0.411 [0.909]	0.490 [0.858]
JB	8.989 [0.011]	10.080 [0.006]	6.335 [0.042]	7.706 [0.021]

Notes: See notes of Table 4.2.

Table 4.7. Parameter estimates for regime-switching models: UK.

	Model I	Model II	Model III	Model IV
μ_1	0.127 (0.387)	−1.932 (4.355)	0.026 (0.417)	10.317** (4.944)
μ_2	–	0.847* (0.454)	–	0.303 (0.363)
α_1	−15.231*** (1.430)	−14.173*** (4.976)	−2.179 (1.646)	−25.548** (5.634)
α_2	0.282 (0.291)	−0.219 (0.324)	0.266 (0.313)	0.076 (0.258)
β_1	−1.565* (0.790)	−2.119* (1.201)	−0.310 (1.143)	−3.783** (1.312)
β_2	−0.117 (0.150)	−0.160 (0.188)	−0.079 (0.153)	−0.143 (0.150)
ϕ_1	0.326*** (0.078)	–	0.417*** (0.098)	0.326*** (0.073)
ϕ_2	−0.369*** (0.074)	–	–	−0.256*** (0.091)
ϕ_3	0.175** (0.075)	–	–	–
ϕ_4	−0.146* (0.076)	–	–	–
ϕ_5	0.145** (0.068)	–	–	–
σ_1	2.399*** (0.190)	3.051*** (0.239)	7.244*** (1.250)	4.590*** (1.624)
σ_2	–	–	2.289*** (0.253)	2.410*** (0.214)
P	0.480*** (0.219)	0.280*** (0.245)	0.881*** (0.097)	0.330*** (0.284)
$1-Q$	0.026 (0.018)	0.037* (0.021)	0.025 (0.026)	0.042* (0.024)
Log Likelihood	−203.040	−237.161	−227.047	−216.853
AIC	432.08	492.322	474.094	457.706
$LB_{(8)}$	24.28 [0.060]	5.349 [0.719]	10.63 [0.155]	4.931 [0.552]
$LB^2_{(8)}$	0.900 [0.523]	0.717 [0.675]	0.652 [0.730]	0.256 [0.977]
JB	0.402 [0.817]	3.848 [0.146]	0.418 [0.811]	0.187 [0.910]

Notes: See notes of Table 4.2.

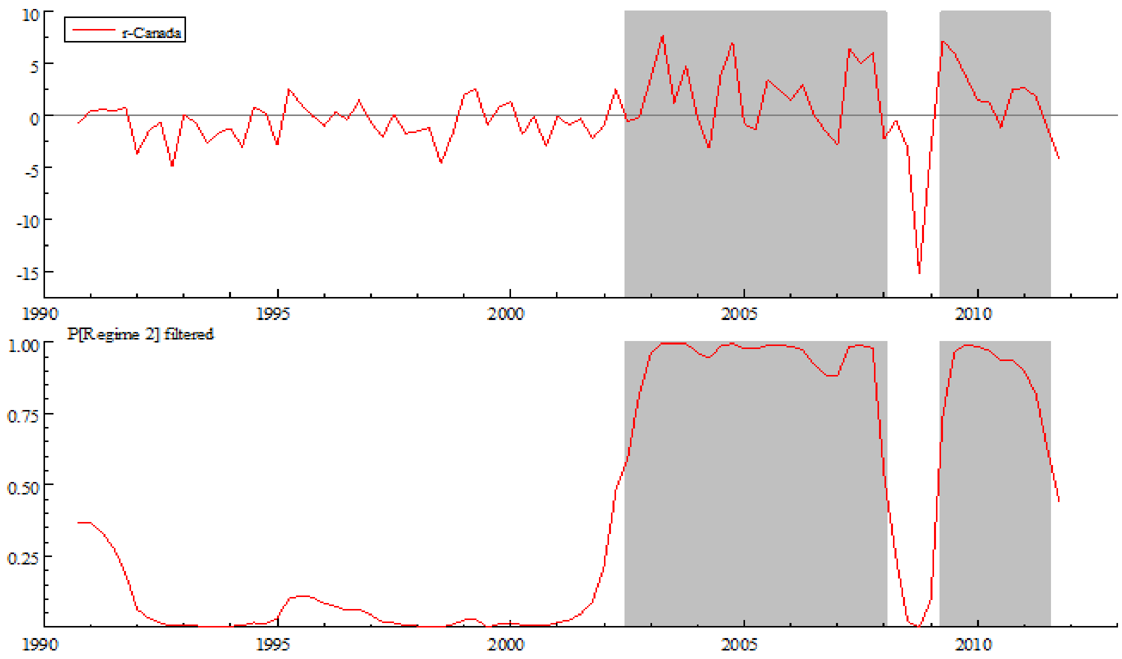

Figure 4.1. Exchange rate changes, r_t, and the transition probabilities for Canada. Regime 1 denotes the probability of staying in the low (appreciation) regime, while regime 2 indicates the probability of staying in the high (depreciation) regime.

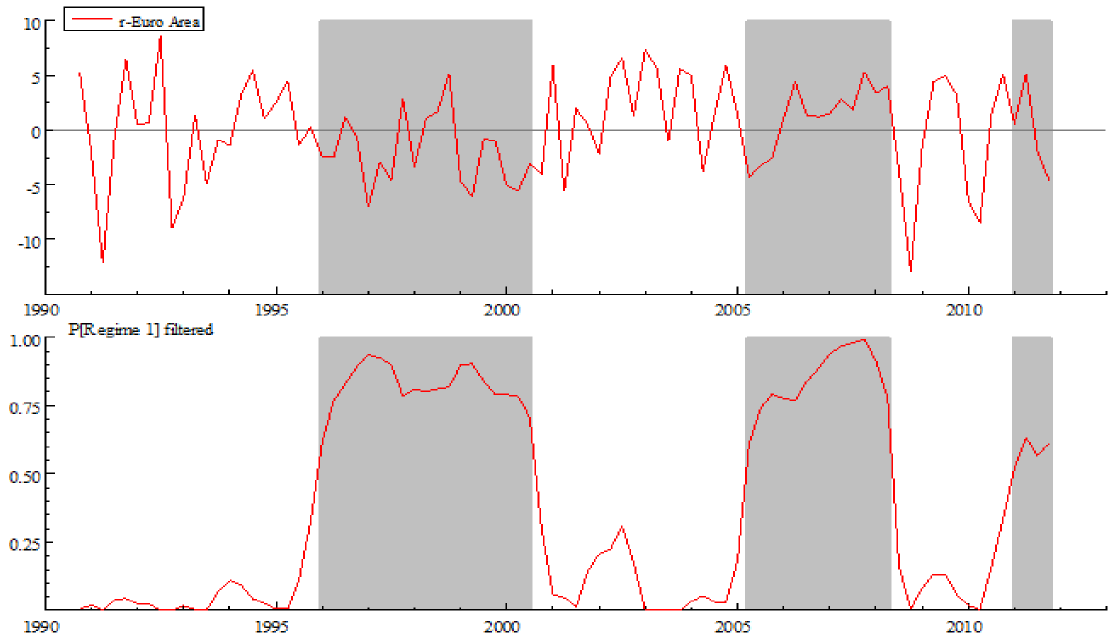

Figure 4.2. Exchange rate changes, r_t, and the transition probabilities for the euro area. Regime 1 denotes the probability of staying in the less volatile regime, while regime 2 indicates the probability of staying in the higher volatile regime.

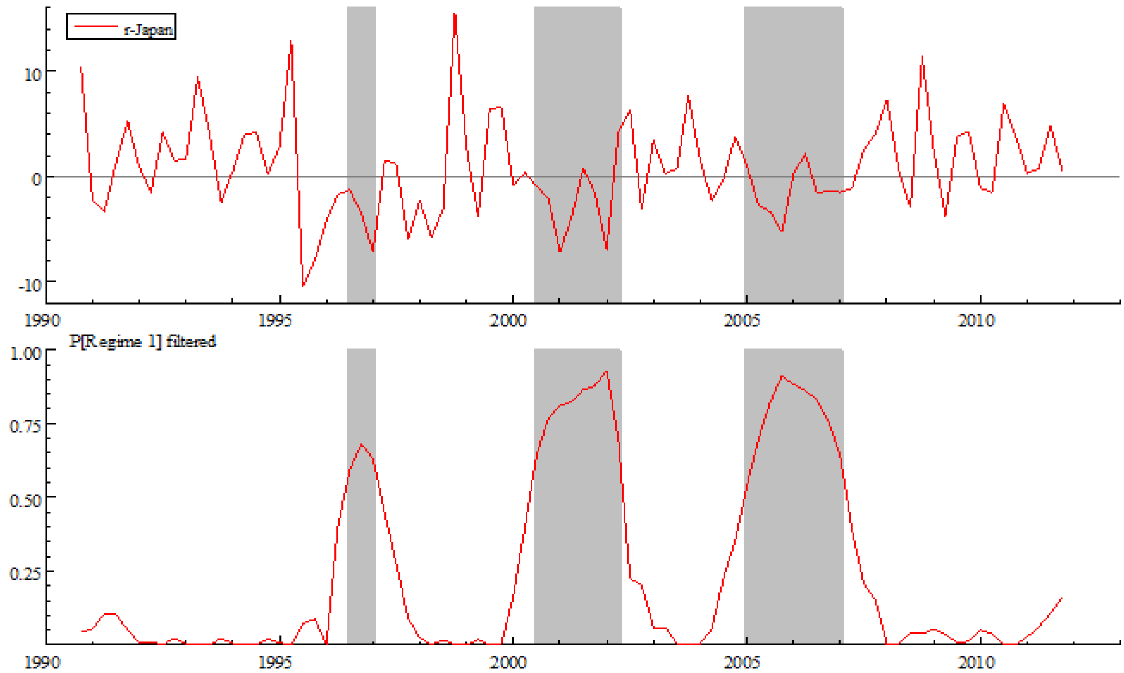

Figure 4.3. Exchange rate changes, r_t, and the transition probabilities for Japan. Regime 1 denotes the probability of staying in the less volatile regime, while regime 2 indicates the probability of staying in the higher volatile regime.

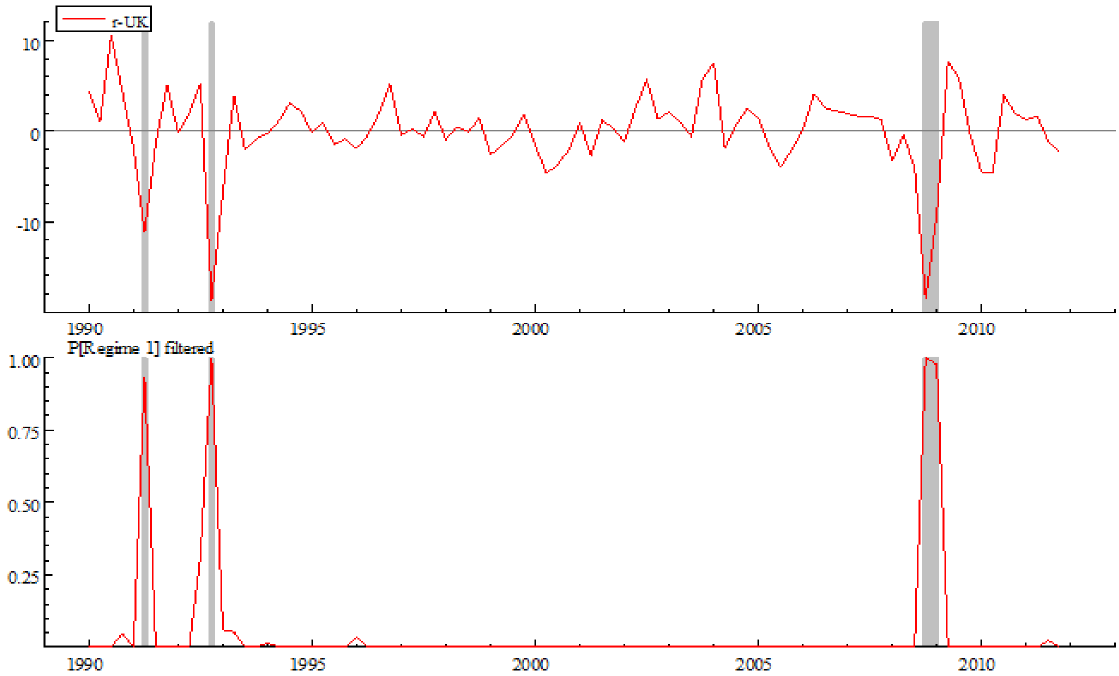

Figure 4.4. Exchange rate changes, r_t, and the transition probabilities for the UK. Regime 1 denotes the probability of staying in the low (appreciation) regime, while regime 2 indicates the probability of staying in the high (depreciation) regime.

inflows towards the US or net equity and net bond outflows from the UK result in an appreciation of the US dollar against the British pound.

The shift in the variance, in the case of the euro area (see Model III in Table 4.5), separates periods of low volatility $(\sigma^l = 2.9)$ from high volatility $(\sigma^h = 4.4)$. The probability of staying in the low (high) regime is 0.88 (0.92). Both regimes are quite persistent, well capturing the cluster effect. There are 36 quarters (42.35%) where the process is in the low regime and 49 quarters (57.65%) where the process is in the high regime (see Figure 4.2). While net equity flows appear to be insignificant in either regime, net bond flows are significant in both regimes. In particular, the results suggest that net bond inflows towards the US or net bond outflows from the euro area result in a US dollar appreciation (depreciation) against the euro in the low (high) volatility regime.

Finally, in the case of Japan (see Model III in Table 4.6), volatility in the high regime $(\sigma^h = 4.5)$ is 73% higher than in the low regime $(\sigma^l = 2.6)$, with the associated transition probabilities being equal to 0.81 and 0.94 for the low and high regimes, respectively. There are 20 quarters (23.53% of the total observations) where the process is in the low variance regime and 65 quarters (76.47%) where the process is in the high variance regime (see Figure 4.3). Net bond flows are found to have a significant impact on exchange rate changes only in periods of low volatility, in other words net inflows towards the US or net outflows from Japan result in a US dollar appreciation against the Japanese yen.

Our findings contradict those of Siourounis (2004), who found that equity flows rather than bond flows drive the dynamics of exchange rates. However, the evidence presented herein shows that the impact of equity flows compared to bond flows is rather limited. This applies to both the euro area and Japan. Furthermore, Brooks et al. (2004)

reported that the yen is likely to be driven by macroeconomic variables rather than portfolio flows. However, using the Markov switching specifications, net portfolio flows, primarily net bond flows, appear to impact on the US dollar-Japanese yen exchange rate in the less volatile regime. A possible explanation for the effect of bond flows may be due to the sterilised interventions exercised by the Bank of Japan in the foreign exchange market over the last two decades. The finding is also consistent with what was highlighted by Hau and Rey (2006) in that international portfolio flows with Japan involve mostly bonds as opposed to equities.

4.6. Conclusions

In this chapter, we have provided some empirical evidence on the causal relationship between net portfolio flows and exchange rate changes using quarterly data from 1990:01 to 2011:04 for the US against the UK, Canada, Japan, and the euro area. The focus is on the nonlinear causal dynamics and the methodology adopted differentiates this study from most other contributions to the literature. Our argument is that investors behave differently when the market is in an appreciation than in a depreciation and when it is highly volatile than in less volatile periods. The existence of multiple equilibria in the behaviour of the exchange rate and its volatility has been well known by now as a result of the associated swings driven by market anomalies, as well as the financial crises occurred over the last two decades (see, Jeanne and Rose, 2002; Chen, 2006; Lovcha and Perez-Laborda, 2013; among others). The linear models proposed in the literature are not rich enough to accommodate those different behaviours. Indeed, the recent study of Lovcha and Perez-Laborda (2013) showed that the relationship between exchange rate changes and customer order flows evolves over time and that such non-linearity can be well captured by the Markov switching models.

The empirical results can be summarised as follows. There is evidence of a nonlinear relationship between exchange rate changes and net portfolio flows in three (euro area, Japan and UK) of the four countries under examination. Canada was the only case where portfolio inflows were found not to impact on exchange rate dynamics, even by accounting for nonlinearities. Though, this result is in line with previous studies on commodity exporting countries (see Chaban, 2009; Ferreira Filipe, 2012).

The debate on the linkages between net portfolio inflows and exchange rate appreciation/depreciation is, clearly, still open, but our findings indicate that careful consideration should be given to the often neglected nonlinearities involved. The results presented in this chapter are a first cut and further analyses are no doubt needed. Since the focus of the paper is on the in-sample model comparisons, future work might also conduct an out-of-sample forecasting to choose among the candidate models in addition to the AIC. Such candidate models might also be compared to the random walk benchmark, even though studies employed the Markov switching specifications have shown that such specifications are outperformed by the random walk benchmark in out-of-sample forecasting (see Engle and Hamilton, 1990; Engle, 1994; Frörmmel et al., 2005; among others).

CHAPTER FIVE

EXCHANGE RATE UNCERTAINTY

AND INTERNATIONAL PORTFOLIO FLOWS

5.1. Introduction

The macroeconomic effects of exchange rate uncertainty, especially on trade flows, have attracted considerable attention since the collapse of the Bretton Woods system in 1971 and the adoption of floating exchange rates in March 1973; both in the theoretical and empirical literature (see McKenzie, 1999, for a comprehensive review). By contrast, the impact at the micro level on equity and bond portfolio flows has drawn less attention in the literature.

Also, there is a substantial literature examines the determinants of international transactions in assets, but they are very few empirical papers analysing the impact of exchange rate uncertainty. For example, Bohn and Tesar (1996) found that investors tend to move to markets where the returns are expected to be high. This ‘return chasing’ hypothesis has also been confirmed by Bekaert et al. (2002), who found that positive returns shocks lead to an increase in short-term equity capital flows using data from 20 emerging countries. Portes et al. (2001) and Portes and Rey (2005), by contrast, showed that transactions in financial assets are explained by the gravity model at least as well as those of trade in goods. Controlling for push and pull factors, Edison and Warnock (2008) showed that the cross-listing of an emerging Asian or Latin American market equity on a US exchange leads to sharp short-horizon equity inflows, while the reduction of capital controls increases equity inflows over longer horizons in emerging Asia, but not in Latin America. More recently, Milesi-Ferretti and Tille (2011) found that capital flows were heterogeneous throughout the recent financial crisis period and across types of flows, with bank flows have been the hardest hit during the period as a result of the marked decline in cross-border lending. Fratzscher (2012), instead, finds evidence that push factors were important drivers of net capital flows during the recent financial crisis, but not during the recovery period (2009-2010) for which domestic pull factors were important during such a period, especially for countries in Emerging Asia and Latin America.

The underlying idea behind the effects of exchange rate uncertainty on international transactions in assets is that exchange rate volatility increases transaction costs and reduces potential gains from international diversification by making the acquisition of foreign securities such as bonds and equities more risky, which in turn affects negatively portfolio flows across borders. Indeed, Eun and Resnick (1988) had previously shown that exchange rate uncertainty is non-diversifiable and has an adverse impact on the performance of international portfolios. This finding is also consistent with the evidence presented in the study by Levich et al. (1999), who found, by surveying 298 US institutional investors, that foreign exchange risk hedging constitutes only 8% of total foreign equity investment. Choi and Rajan (1997) also found, by using data from seven major developed countries outside of the US, that foreign exchange risk has a significant effect on asset returns and ignoring such a factor induces misspecification when analysing the integration or segmentation of international capital markets. By considering a wide range of developed and emerging market economies, Fidora et al. (2007) and Borensztein and Loungani (2011) further found that exchange rate volatility is an essential factor for bilateral equity and bond portfolio home bias.

However, Eun and Resnick (1988) suggest that hedging through forward exchange contracts and multicurrency diversification are effective ways to reduce exchange rate risk. Glen and Jorion (1993) and Eun and Resnick (1994) further provide evidence that hedging in the forward exchange markets improves the performance of diversified portfolios of equities and bonds. Jorion (1991), in a related vein, also found that the exchange rate risk is diversifiable in which such risk appears to be not priced in the US stock market. Gehrig (1993), instead, argued that exchange rate risk, purchasing power risk, and capital market restrictions are insufficient factors for equity portfolio home bias, where the informational segmentation to be of significance for such behaviour instead.

Furthermore, while Hau and Rey (2006) provided a theoretical framework for analysing the implications of incomplete foreign exchange risk trading for the correlation structure of exchange rate changes and equity returns, as well as exchange rate changes and net portfolio flows,[57] they did not include statistical tests for the impact of exchange rate uncertainty on portfolio flows across borders.

The present chapter makes a fourfold contribution to the existing literature. First, it analyses empirically whether exchange rate uncertainty affects international portfolio flows and their variability using a bivariate VAR GARCH (1, 1)-in-mean framework. It is in fact the first empirical investigation of this kind, based on bilateral monthly data for the US *vis-à-vis* six developed economies, namely Australia, Canada, the euro area, Japan, Sweden, and the UK over the period 1988:01-2011:12. It follows that our analysis is based on longer monthly time series and differs from previous studies which focus on the determinants of home bias and international transactions in assets using panel and cross-sectional techniques (see, for examples, Portes and Rey, 2005; Fidora et al., 2007; Bekaert and Wang, 2009; Batten and Vo, 2010; Borensztein and Loungani,

[57] Their analysis was motivated by the recent microstructure approach to exchange rate determination which had been shown to improve remarkably the performance of exchange rate models, with currency order flows explaining a substantial proportion of exchange rate changes (see, e.g., Evans and Lyons, 2002; 2005; 2008; Payne, 2003; Rime et al., 2010; Chinn and Moore, 2011; and Duffuor et al., 2012 among others).

2011; Mishra, 2011; Mercado, 2013; Daly and Vo, 2013). Moreover, we use the most appropriate measure of volatility (GARCH) in the context of time series analysis, as opposed to other measures of exchange rate volatility used in the literature on home bias and portfolio flows determinants, such as the continuous volatility measure in Portes and Rey (2005), the stochastic deviation from PPP in Fidora et al. (2007) and Mishra (2011), the standard deviation of exchange rate changes in Bekaert and Wang (2009) and the coefficient of variation of real exchange rate in Mercado (2013).

Second, unlike Hau and Rey (2006) who assume that the supply of bonds is infinitely elastic, thereby simplifying the dynamics of bond acquisitions in their model, we examine the impact of exchange rate uncertainty on bond and equity flows (as well as their variability) in turn. In this way, we are able to evaluate the impact of uncertainty on the individual components of portfolio flows across borders. According to Hau and Rey (2006), exchange rate uncertainty should affect equity, but not bond flows. Fidora et al. (2007) and Borensztein and Loungani (2011), by contrast, found evidence that bond flows exhibit stronger home bias compared with equity flows. We provide some relevant empirical evidence on this issue.

Third, existing empirical studies on the relationship between exchange rate changes and portfolio flows investigate short-run dynamic interactions only with linear dependence techniques (i.e., first moment analysis). For example, Brooks et al. (2004) and Hau and Rey (2006) use simple correlations and regression analysis for the US *vis-à-vis* the euro area and Japan, and 17 OECD countries, respectively; Siourounis (2004), Chaban (2009), and Kodong and Ojah (2012) estimate VAR models respectively for four developed countries (the UK, Japan, Germany, and Switzerland), three oil-exporting countries (Canada, Australia, and New Zealand), and four African countries (Egypt, Morocco, Nigeria, and South Africa) *vis-à-vis* the US. Their results are characterised by significant deviations from normality and conditional heteroscedasticity, i.e., volatility clustering or the so-called ARCH effects (see Engle, 1982) that are not captured by their setup. By contrast, we model first and second moments simultaneously to analyse the dynamic interactions between exchange rate changes and portfolio flows, in this way avoiding the potential pitfalls of earlier studies.

Fourth, since volatility is a measure of the information flow (see Ross, 1989), it is of paramount importance to understand how the stochastic information arrivals in the form of simple portfolio investment shifts in bonds and equities are transmitted to the foreign exchange market, and vice versa. Our analysis sheds light on this mechanism and thus provides important information to policy-makers and regulators to formulate appropriate policies based on imposing or relaxing credit controls on these flows depending on the state of the economy, with the aim of achieving economic and financial stability.

The remainder of the chapter is organised as follows. Section 5.2 outlines the econometric model used in the study. Section 5.3 describes the data and reports some descriptive statistics. Section 5.4 discusses the empirical results, and finally Section 5.5 concludes the chapter.

5.2. The econometric model

We employ a bivariate VAR-GARCH (1, 1) in the BEKK specification (Engle and Kroner, 1995) allowing for in-mean effects in order to examine the impact of exchange rate uncertainty on net equity and net bond flows as well as the dynamic linkages in the first and second moments of these variables over the period 1988:01-2011:12. In addition to the contemporaneous effect, various lags of exchange rate volatility affecting the conditional mean of net equity and net bond flows are included in the specification to avoid the potential pitfalls of models allowing only for contemporaneous interactions. The economic interpretation for the inclusion of the lags of the exchange rate volatility is that the investors' response to exchange rate volatility might take some time to be incorporated into their strategies. Therefore the conditional mean equation is specified as follows:

$$y_t = \mu + \sum_{i=1}^{p} \psi_i y_{t-i} + \sum_{i=0}^{p} \eta_i h_{t-i} + \varepsilon_t$$

$$\mu = \begin{bmatrix} \mu_1 \\ \mu_2 \end{bmatrix}; \psi_i = \begin{bmatrix} \psi_{11}^{(i)} & \psi_{12}^{(i)} \\ \psi_{21}^{(i)} & \psi_{22}^{(i)} \end{bmatrix}; \eta_i = \begin{bmatrix} \eta_{11}^{(i)} & \eta_{12}^{(i)} \\ \eta_{21}^{(i)} & \eta_{22}^{(i)} \end{bmatrix}; \varepsilon_t = \begin{bmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \end{bmatrix} \tag{5.1}$$

where $y_t = [E_t, EF_t(BF_t)]$, E_t and EF_t (BF_t) indicate respectively exchange rate changes and net equity (bond) flows. $h_t = [h_{11,t}, h_{22,t}]$, $h_{11,t}$ and $h_{22,t}$ represent respectively the conditional variances of exchange rate changes and net flows depending on whether equities or bonds are considered. The parameters $\psi_{11}^{(i)}$ and $\psi_{22}^{(i)}$ measure the response of exchange rate changes and net flows to their own lags, whilst $\psi_{21}^{(i)}$ and $\psi_{12}^{(i)}$ represent Granger causality from exchange rate changes to net flows, and vice versa (i denotes the lagged time-period). If the parameter $\eta_{21}^{(i)}$ is significantly different from zero, it implies that exchange rate uncertainty affects net equity flows and/or net bond flows (the lag length in this case is defined as i = 0, 1,.., p, with 0 indicating the contemporaneous effect). The innovations vector is assumed to be normally distributed $\varepsilon_t | \Omega_{t-1} \sim (0, H_t)$ with its corresponding variance-covariance matrix given by H_t; Ω_{t-1} is the information set available at time t-1. Lags are included sequentially in Eq. (5.1) until serial correlation is removed by employing the Hosking (1981) multivariate Q-statistic to the standardised residuals, $z_{s,t} = \varepsilon_{s,t}/\sqrt{h_{s,t}}$ for s = 1, 2.

Having specified the conditional mean equation, we then estimate the multivariate GARCH model in its BEKK representation, this being a straightforward generalisation of the univariate GARCH model of Bollerslev (1986). The BEKK specification has advantages compared to other multivariate GARCH specifications such as the VEC-GARCH model of Bollerslev et al. (1988) because of its quadratic forms ensuring that the conditional covariance matrices in the system are positive definite.[58] Unlike the dynamic conditional correlation model of Engle (2002), which estimates the time-varying correlations directly, the BEKK specification allows for time-varying correlations and also for interactions between the variances in a lead-lag framework. Furthermore, the curse of dimensionality highlighted by Caporin and McAleer (2012) is not a serious issue in the present case with only two variables. The model can be represented as follows:

$$H_t = C'C + A'\varepsilon_{t-1}\varepsilon'_{t-1}A + B'H_{t-1}B\,. \tag{5.2}$$

In matrix form, it can be specified as:

$$\begin{bmatrix} H_{11,t} & H_{12,t} \\ H_{21,t} & H_{22,t} \end{bmatrix} = C'C + A'\begin{bmatrix} \varepsilon_{1,t-1}^2 & \varepsilon_{1,t-1}\varepsilon_{2,t-1} \\ \varepsilon_{2,t-1}\varepsilon_{1,t-1} & \varepsilon_{2,t-1}^2 \end{bmatrix}A + B'\begin{bmatrix} H_{11,t-1} & H_{12,t-1} \\ H_{21,t-1} & H_{22,t-1} \end{bmatrix}B\,, \tag{5.3}$$

$$C = \begin{bmatrix} c_{11} & 0 \\ c_{21} & c_{22} \end{bmatrix}, A = \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{bmatrix}, B = \begin{bmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{bmatrix}$$

where C is constrained to be a lower triangular matrix and A and B are respectively ARCH and GARCH parameter matrices. Eq. (5.3) shows that in the BEKK model each conditional variance and covariance in H_t is modelled as a function of lagged conditional variances and covariances, lagged squared innovations and the cross-product of the innovations. Volatility is transmitted between exchange rate changes and

[58] For a survey on multivariate GARCH models, see Bauwens et al. (2006) or Silvennoinen and Teräsvirta (2009).

net equity/bond flows through two channels represented by the off-diagonal parameters in the ARCH and GARCH matrices: a symmetric shock $\varepsilon_{ii,t-1}$ and the conditional variance $H_{ii,t-1}$. More specifically, volatility transmission from exchange rate changes to net equity/bond flows can be analysed by testing the null hypothesis $a_{12} = b_{12} = 0$, and $a_{21} = b_{21} = 0$ in the opposite direction. These causality-in-variance tests within the multivariate GARCH-BEKK models have superior power to the cross correlation function (CCF) of Cheung and Ng (1996) which is a two-step approach (see Hafner and Hewartz, 2008). Causality-in-variance is tested using the following likelihood ratio test statistic:

$$LR = -2(L_r - L_{ur}) \sim x^2{}_{df} \tag{5.4}$$

where L_r and L_{ur} indicate respectively the restricted and unrestricted log-likelihood functions; LR follows the chi-squared distribution with degrees of freedom equal to the number of the restricted coefficients (df).

Given that the innovations are assumed to be normally distributed, as stated earlier, the log-likelihood function for such a model is given by:

$$L(\theta) = \frac{-Tn}{2}\ln(2\pi) - \frac{1}{2}\sum_{t=1}^{t}(\ln|H_t| + \varepsilon'_t H^{-1}{}_t \varepsilon_t) \tag{5.5}$$

where n is the number of equations, two in our case; T is the number of observations, which is 287; and θ is a vector of unknown parameters to be computed. More specifically, we use the Quasi-Maximum Likelihood (QML) method of Bollerslev and Woolbridge (1992) to calculate the standard errors that are robust to deviations from normality.[59] As a final check of the adequacy of the estimated model we employ the Hosking (1981) multivariate *Q*-statistic for the standardised squared residuals to evaluate whether or not the ARCH and GARCH dynamics have been appropriately captured in the conditional variance-covariance equation, Eq. (5.3).

5.3. Data description

We examine the impact of exchange rate uncertainty on different components of net portfolio flows, namely net equity and net bond flows, as well as the dynamic linkages between these flows and exchange rate changes for the US *vis-à-vis* the UK, Japan, Canada, Australia, Sweden, and the euro area. Throughout, the US is considered the domestic or home economy. Since the data on portfolio investment flows, obtained from the US *Treasury International Capital (TIC) System*,[60] are sampled at a monthly frequency, we employ monthly data from 1988:01 to 2011:12 for all series. For the limitations of the TIC data, the reader is directed to **Chapter 4** or Edison and Warnock (2008). The reason for the chosen start date is that portfolio flows for the period preceding 1988 are known to be insignificant (see Brooks et al., 2004). Net equity (bond) flows are calculated as equity (bond) *inflows* minus *outflows*. While *inflows* are measured as net purchases and sales of domestic (US) assets (equities and bonds) by foreign residents, *outflows* are measured as net purchases and sales of foreign assets (equities and bonds) by domestic (US) residents. With regard to the euro area, we aggregate the data for the individual EMU countries (Austria, Belgium-Luxemburg, Finland, France, Germany, Ireland, Italy, the Netherlands, Portugal, Spain) to extract

[59] We use the SIMPLEX free-derivative method, which is useful to improve the initial values, and then the BFGS standard algorithm to obtain the standard errors (see Engle and Kroner, 1995; Kearney and Patton, 2000; among others). This procedure was implemented in RATS 8.1 with a convergence criterion of 0.00001.

[60] The data are retrieved from the US Treasury Department website: http//www.treasury.gov/resource-center/data-chart-center/tic/Pages/country-longterm.aspx

cross-border net bond and net equity flows between the US and this region (Heimonen, 2009).

Positive numbers imply net equity and net bond inflows (in millions of US dollars) towards the US or net outflows from the counterpart countries. Following Brennan and Cao (1997), Hau and Rey (2006), and Chaban (2009) among others, we normalise these flows using the average of their absolute values over the previous 12 months, since without scaling model convergence is difficult to achieve. The exchange rates are end of period data, defined as US dollars per unit of foreign currency; the source is the IMF's *International Financial Statistics (IFS).* Exchange rate changes are calculated as $E_t = 100 \times P_{E,t}/P_{E,t-1}$ where $P_{E,t}$ represents the log of the exchange rate at time *t*. For the period preceding the inception of the euro, i.e. before 1999, we use US dollar per ECU as the euro area's exchange rate.

Descriptive statistics are displayed in Table 5.1. The mean of monthly exchange rate changes is positive (US dollar depreciation) for Japan and Canada, and negative (US dollar appreciation) for the rest of the countries. On the other hand, the monthly mean of net equity flows is positive for Sweden and Canada and negative for the remaining countries, indicating equity inflows from Sweden and Canada towards the US and outflows from the US towards the other countries. The monthly mean of net bond flows is negative for Australia and positive for the other countries. This indicates the existence of bond inflows from all countries towards the US except Australia for which there is evidence of bond outflows from the US towards Australia. Exchange rate changes are found to exhibit higher volatility than the two net flows. Furthermore, net equity flows appear to be characterised by higher volatility than net bond flows (although their volume is very small).

Table 5.1. Summary of descriptive statistics for the normalised net portfolio flows and exchange rate changes.

		Australia	Canada	Euro area	Japan	Sweden	UK
Mean	E_t	-0.122	0.083	-0.002	0.160	-0.047	-0.066
	EF_t	-0.200	0.068	-0.051	-0.432	0.020	-0.017
	BF_t	-0.106	0.191	0.222	0.718	0.260	0.848
St. Dev	E_t	3.270	2.148	3.080	3.088	3.439	2.855
	EF_t	1.599	1.443	1.487	1.552	1.729	1.414
	BF_t	1.467	1.394	1.358	1.251	1.638	1.136
Skewness	E_t	0.790	-0.692	-0.375	0.221	-0.554	-0.738
	EF_t	-1.129	0.144	0.028	-0.631	-1.333	-0.342
	BF_t	-0.446	-0.202	-0.365	0.634	0.379	-0.385
Ex. kurtosis	E_t	6.226	9.417	4.119	4.958	5.410	5.634
	EF_t	10.619	4.301	4.157	6.103	8.363	3.607
	BF_t	4.988	3.830	3.665	7.905	7.914	9.786
JB	E_t	154.3***	515.3***	21.71***	48.19***	84.17***	109.0***
	EF_t	755.3***	21.26***	16.06***	134.2***	429.0***	10.02***
	BF_t	56.83***	10.20***	11.69***	306.9***	295.6***	557.8***

Notes: E_t, EF_t, and BF_t indicate exchange rate changes, net equity flows, and net bond flows, respectively; *JB* is the Jarque-Bera test for normality.
*** indicate significance at the 1 % level.

As for the third and fourth moments, exchange rate changes, net equity flows, and net bond flows all exhibit skewness and excess kurtosis in most cases. The Jarque-Bera (JB) test statistics imply a rejection at the 1% level of the null hypothesis that exchange rate changes and the two net flows are normally distributed in all countries in question.

Figure 5.1 shows monthly exchange rate changes, net equity flows and net bond flows in all countries over the period under investigation. Volatility clustering is clearly present in all cases, suggesting that an ARCH model might be required to capture it. All series also appear to be covariance stationary.

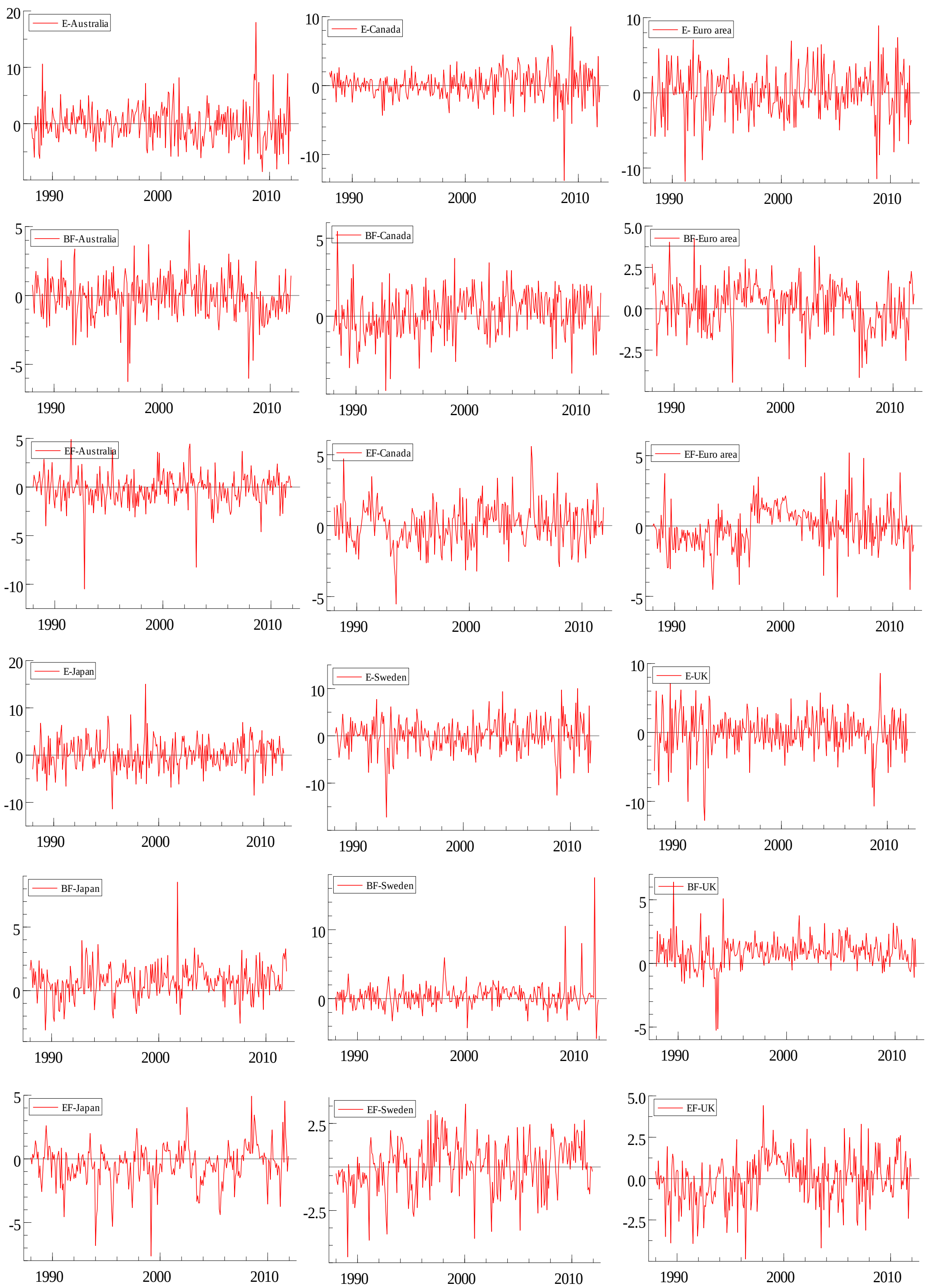

Figure 5.1. Time series of exchange rate changes (*E*), net bond flows (*BF*), and net equity flows (*EF*) of the six advanced economies over the period 1988:01–2011:12.

5.4. Empirical results

The objective of our analysis is to establish whether exchange rate uncertainty affects net equity and net bond flows across borders, and also whether there is a volatility transmission (hence information flows) between these flows and exchange rate changes and, if so, in what direction causality runs.[61]

The QML estimates of the bivariate GARCH **(**1, 1)–BEKK–in mean parameters as well as the associated multivariate *Q*-statistics (Hosking, 1981) are displayed in Tables 5.2–5.7 for Australia, Canada, the euro area, Japan, Sweden, and the UK, respectively. Panel A and B in each Table concern the bivariate model of exchange rate changes against net equity and net bond flows, respectively. The reported Hosking multivariate *Q*-statistic of order (6) and (12) for the standardised residuals in the exchange rate changes-net equity flows equation indicate the existence of no serial correlation at the 5% level, when the conditional mean equations are specified with p=1 for Japan, p=2 for Sweden and p=3 for the other countries (the insignificant parameters in the mean equations have been dropped). With regard to the exchange rate changes-net bond flows relationship, whilst no dynamic terms appear to be necessary for Sweden, setting p=1 for the UK, p=2 for the euro area, p=3 for Australia and Canada, and p=5 for Japan is required to capture adequately the dynamic structure in these cases.

[61] The results are relatively the same across the considered countries by examining the impact of exchange rate uncertainty on net equity and bond portfolio flows as an aggregate. However, here the attention is paid to the individual components of net portfolio flows such as net equity and net bond flows. Knowledge of the impact of exchange rate uncertainty on the exact component(s) of portfolio flows can help financial regulators and policy makers to target the exact market(s) to achieve economic and financial stability.

Table 5.2. The estimated bivariate GARCH–BEKK–in mean model for Australia.

	Panel A. Exchange rates (E_t) and equity flows(EF_t)			*Panel B. Exchange rates (E_t) and bond flows (BF_t)*	
	E_t (s=1)	EF_t (s=2)		E_t (s=1)	BF_t (s=2)
Conditional Mean Equation					
μ_s	-0.015 (0.159)	-0.348^{***} (0.120)	μ_s	-0.128 (0.157)	0.158 (0.129)
$\psi_{2s}^{(2)}$	–	0.157^{***} (0.079)	$\psi_{2s}^{(3)}$	–	0.129^{***} (0.049)
$\psi_{1s}^{(3)}$	0.110^{*} (0.062)	–	$\eta_{2s}^{(0)}$	-0.026^{**} (0.010)	–
$\eta_{2s}^{(5)}$	0.014^{*} (0.007)	–			
Conditional Variance Equation					
c_{1s}	0.496^{***} (0.496)	0	c_{1s}	-0.103 (0.281)	0
c_{2s}	-0.042 (1.760)	1.352^{**} (0.545)	c_{2s}	0.753^{***} (0.129)	-0.00008 (1.148)
a_{1s}	0.363^{***} (0.087)	0.027 (0.058)	a_{1s}	0.254^{***} (0.046)	-0.011 (0.030)
a_{2s}	-0.133 (0.241)	-0.205 (0.311)	a_{2s}	0.380^{***} (0.076)	0.076 (0.152)
b_{1s}	0.920^{***} (0.037)	0.014 (0.039)	b_{1s}	0.949^{***} (0.010)	0.001 (0.004)
b_{2s}	0.062 (0.785)	0.472^{*} (0.256)	b_{2s}	0.033 (0.071)	0.849^{***} (0.065)
Loglik	-1254.54		*Loglik*	-1225.38	
$Q(6)$	27.654[0.274]	$Q^2(6)$ 9.823[0.981]	$Q(6)$	12.073 [0.979]	$Q^2(6)$ 26.041[0.204]
$Q(12)$	49.470[0.414]	$Q^2(12)$ 30.46[0.952]	$Q(12)$	31.67 [0.966]	$Q^2(12)$ 48.899[0.319]

Tests of No Volatility Transmission:		Tests of No Volatility Transmission:	
(i) Bidirectional between E_t and EF_t		(i) Bidirectional between E_t and BF_t	
$H_0: a_{12}=a_{21}=b_{12}=b_{21}=0$	*LR*=1.74[0.781]	$H_0: a_{12}=a_{21}=b_{12}=b_{21}=0$	*LR*=11.6 [0.020]
(ii) From E_t to EF_t		(ii) From E_t to BF_t	
$H_0: a_{12}=b_{12}=0$	*LR*=0.12[0.939]	$H_0: a_{12}=b_{12}=0$	*LR*=0.13 [0.934]
(iii) From EF_t to E_t		(iii) From BF_t to E_t	
$H_0: a_{21}=b_{21}=0$	*LR*=1.63[0.440]	$H_0: a_{21}=b_{21}=0$	*LR*=10.3 [0.005]

Notes: E_t, EF_t, and BF_t indicate exchange rate changes, net equity flows, and net bond flows, respectively; while *LR* indicates likelihood ratio test statistics. Heteroscedasticity-consistent standard errors are in parentheses (.), whereas *p*-values are reported in [.]. The superscripts of the ψ parameters denote the lagged time periods, with zero being the contemporaneous effect. $Q(p)$ and $Q^2(p)$ are multivariate Hosking (1981) tests for p^{th} order serial correlation on the standardised residuals z_{it} and their squares z_{it}^2, respectively where i = 1 (for exchange rate changes (E_t)), 2 (for net equity flows (EF_t) and net bond flows (BF_t)). The covariance stationarity condition is satisfied by all the estimated models, all the eigenvalues of ($A_{11}\otimes A_{11}$ + $B_{11}\otimes B_{11}$) being less than one in modulus.

***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 5.3. The estimated bivariate GARCH–BEKK–in mean model for Canada.

Panel A. Exchange rates (E_t) and equity flows(EF_t)

	E_t (s=1)	EF_t (s=2)
Conditional Mean Equation		
μ_s	-0.034 (0.097)	0.065 (0.081)
$\psi_{2s}^{(1)}$	–	0.249^{***} (0.061)
$\psi_{2s}^{(3)}$	–	0.143^{***} (0.053)
Conditional Variance Equation		
c_{1s}	0.060 (0.164)	0
c_{2s}	1.270^{***} (0.224)	0.00102 (3.260)
a_{1s}	0.328^{***} (0.050)	-0.017 (0.061)
a_{2s}	-0.001 (0.097)	0.260^{**} (0.131)
b_{1s}	0.921^{***} (0.034)	-0.097 (0.103)
b_{2s}	-0.274^{*} (0.158)	-0.242 (0.603)
Loglik	-1079.47	
$Q(6)$	16.201 [0.880]	$Q^2(6)$ 13.294 [0.897]
$Q(12)$	29.301 [0.984]	$Q^2(12)$ 37.210 [0.788]

Panel B. Exchange rates (E_t) and bond flows (BF_t)

	E_t (s=1)	BF_t (s=2)
Conditional Mean Equation		
μ_s	-0.035 (0.099)	0.160^{*} (0.084)
$\psi_{1s}^{(2)}$	–	0.136^{*} (0.067)
$\psi_{2s}^{(3)}$	–	0.121^{*} (0.070)
$\eta_{2s}^{(0)}$	0.026^{*} (0.013)	–
Conditional Variance Equation		
c_{1s}	0.230^{**} (0.108)	0
c_{2s}	0.0005 (0.063)	-0.0000007 (0.012)
a_{1s}	0.314^{**} (0.047)	-0.070^{**} (0.031)
a_{2s}	0.0002 (0.038)	-0.109^{***} (0.036)
b_{1s}	0.947^{***} (0.018)	0.017^{**} (0.008)
b_{2s}	0.004 (0.013)	0.989^{***} (0.006)
Loglik	-1075.08	
$Q(6)$	13.329 [0.960]	$Q^2(6)$ 8.539[0.992]
$Q(12)$	31.505 [0.968]	$Q^2(12)$ 30.70[30.70]

Tests of No Volatility Transmission:

(i) Bidirectional between E_t and EF_t

$H_0:\ a_{12}=a_{21}=b_{12}=b_{21}=0$ LR=2.01[0.733]

(ii) From E_t to EF_t

$H_0:\ a_{12}=b_{12}=0$ LR=1.23[0.538]

(iii) From EF_t to E_t

$H_0:\ a_{21}=b_{21}=0$ LR=0.79[0.670]

Tests of No Volatility Transmission:

(i) Bidirectional between E_t and BF_t

$H_0:\ a_{12}=a_{21}=b_{12}=b_{21}=0$ LR=8.69 [0.069]

(ii) From E_t to BF_t

$H_0:\ a_{12}=b_{12}=0$ LR=8.11 [0.017]

(iii) From BF_t to E_t

$H_0:\ a_{21}=b_{21}=0$ LR=7.77 [0.020]

Notes: See notes to Table 5.2.

Table 5.4. The estimated bivariate GARCH–BEKK–in mean model for the euro area.

Panel A. Exchange rates (E_t) and equity flows (EF_t)	E_t (s=1)	EF_t (s=2)	*Panel B. Exchange rates (E_t) and bond flows (BF_t)*	E_t (s=1)	BF_t (s=2)
Conditional Mean Equation					
μ_s	-0.065 (0.178)	1.818^{**} (0.916)	μ_s	0.023 (0.194)	0.627^{***} (0.274)
$\psi_{1s}^{(1)}$	–	0.229^{**} (0.101)	$\psi_{2s}^{(1)}$	–	0.142^{**} (0.058)
$\psi_{2s}^{(2)}$	–	0.314^{***} (0.058)	$\psi_{2s}^{(2)}$	–	0.171^{***} (0.059)
$\psi_{2s}^{(3)}$	–	0.129^{**} (0.057)	$\eta_{2s}^{(2)}$	-0.049^{*} (0.027)	–
$\eta_{2s}^{(0)}$	-0.202^{*} (0.105)	–			
Conditional Variance Equation					
c_{1s}	0.480^{***} (0.113)	0	c_{1s}	0.294 (0.252)	0
c_{2s}	-0.819^{***} (0.069)	-0.0001 (0.181)	c_{2s}	0.402^{***} (0.096)	-0.000005 (0.056)
a_{1s}	0.115^{***} (0.030)	0.021 (0.027)	a_{1s}	0.174^{***} (0.066)	0.010 (0.027)
a_{2s}	0.001 (0.074)	0.382^{***} (0.073)	a_{2s}	0.313^{***} (0.120)	-0.159^{***} (0.067)
b_{1s}	0.980^{***} (0.007)	0.003 (0.007)	b_{1s}	0.968^{***} (0.020)	0.018^{***} (0.008)
b_{2s}	0.038 (0.027)	0.910^{***} (0.030)	b_{2s}	-0.134^{***} (0.049)	0.936^{***} (0.021)
Loglik	-1185.16		*Loglik*	-1193.43	
$Q(6)$	20.615 [0.661]	$Q^2(6)$ 24.614 [0.264]	$Q(6)$	18.292 [0.788]	$Q^2(6)$ 11.580[0.950]
$Q(12)$	43.803 [0.645]	$Q^2(12)$ 40.661 [0.656]	$Q(12)$	40.470 [0.771]	$Q^2(12)$ 40.514[0.662]

Tests of No Volatility Transmission:

(i) Bidirectional between E_t and EF_t

$H_0:\ a_{12}=a_{21}=b_{12}=b_{21}=0$ LR=9.35[0.052]

(ii) From E_t to EF_t

$H_0:\ a_{12}=b_{12}=0$ LR=1.82[0.401]

(iii) From EF_t to E_t

$H_0:\ a_{21}=b_{21}=0$ LR=7.86[0.019]

Tests of No Volatility Transmission:

(i) Bidirectional between E_t and BF_t

$H_0:\ a_{12}=a_{21}=b_{12}=b_{21}=0$ LR=12.8 [0.011]

(ii) From E_t to BF_t

$H_0:\ a_{12}=b_{12}=0$ LR=3.08 [0.213]

(iii) From BF_t to E_t

$H_0:\ a_{21}=b_{21}=0$ LR=12.8 [0.001]

Notes: See notes to Table 5.2.

Table 5.5. The estimated bivariate GARCH–BEKK–in mean model for Japan.

Panel A. Exchange rates (E_t) and equity flows (EF_t)			*Panel B. Exchange rates (E_t) and bond flows (BF_t)*		
	E_t (s=1)	EF_t (s=2)		E_t (s=1)	BF_t (s=2)
Conditional Mean Equation					
μ_s	0.112 (0.190)	-0.199^{**} (0.082)	μ_s	0.286 (0.211)	1.472^{***} (0.426)
$\psi_{1s}^{(1)}$	0.100^{*} $0.062)$	–	$\psi_{1s}^{(1)}$	–	-0.390^{***} (0.129)
$\psi_{2s}^{(1)}$	–	0.530^{***} (0.046)	$\psi_{2s}^{(1)}$	0.077^{***} (0.026)	0.126^{*} (0.073)
			$\psi_{2s}^{(2)}$	0.065^{***} (0.020)	-0.084^{*} (0.046)
			$\psi_{1s}^{(3)}$	0.104^{**} (0.048)	0.037^{*} (0.021)
			$\psi_{2s}^{(3)}$	0.037^{*} (0.021)	0.101^{**} (0.049)
			$\psi_{1s}^{(4)}$	-0.098^{*} (0.059)	–
			$\psi_{1s}^{(5)}$	-0.125^{**} (0.055)	–
			$\eta_{2s}^{(0)}$	-0.091^{*} (0.050)	–
Conditional Variance Equation					
c_{1s}	2.192^{***} (0.300)	0	c_{1s}	1.600^{***} (0.214)	0
c_{2s}	0.012 (0.266)	-0.000002 (0.156)	c_{2s}	-0.243 (0.196)	0.743^{***} (0.089)
a_{1s}	0.356^{***} (0.098)	0.031 (0.032)	a_{1s}	0.265^{***} (0.073)	-0.047 (0.032)
a_{2s}	0.357 (0.315)	0.327^{**} (0.133)	a_{2s}	-0.259 (0.343)	0.528^{**} (0.229)
b_{1s}	0.542^{***} (0.132)	-0.231^{***} (0.031)	b_{1s}	0.799^{***} (0.030)	0.124^{***} (0.038)
b_{2s}	0.624^{*} (0.349)	0.753^{***} (0.081)	b_{2s}	-0.241 (0.265)	0.439^{*} (0.284)
Loglik	-1195.794		*Loglik*	-1157.405	
$Q(6)$	31.611 [0.136]	$Q^2(6)$ 15.878 [0.776]	$Q(6)$	23.606 [0.484]	$Q^2(6)$ 12.521 [0.924]
$Q(12)$	64.352 [0.057]	$Q^2(12)$ 28.645 [0.972]	$Q(12)$	57.582 [0.161]	$Q^2(12)$ 28.743 [0.971]

Tests of No Volatility Transmission:		Tests of No Volatility Transmission:	
(i) Bidirectional between E_t and EF_t		(i) Bidirectional between E_t and BF_t	
$H_0: a_{12}=a_{21}=b_{12}=b_{21}=0$	LR=16.9 [0.001]	$H_0: a_{12}=a_{21}=b_{12}=b_{21}=0$	LR=5.82 [0.212]
(ii) From E_t to EF_t		(ii) From E_t to BF_t	
$H_0: a_{12}=b_{12}=0$	LR=10.5 [0.005]	$H_0: a_{12}=b_{12}=0$	LR=1.45 [0.482]
(iii) From EF_t to E_t		(iii) From BF_t to E_t	
$H_0: a_{21}=b_{21}=0$	LR=9.66 [0.007]	$H_0: a_{21}=b_{21}=0$	LR=4.14 [0.126]

Notes: See notes to Table 5.2.

Table 5.6. The estimated bivariate GARCH–BEKK–in mean model for Sweden.

	Panel A. Exchange rates (E_t) and equity flows (EF_t)		*Panel B. Exchange rates (E_t) and bond flows (BF_t)*		
	E_t (s=1)	EF_t (s=2)		E_t (s=1)	BF_t (s=2)
Conditional Mean Equation					
μ_s	0.118 (0.179)	0.045 (0.196)	μ_s	0.066 (0.165)	0.597^{***} (0.130)
$\psi_{2s}^{(1)}$	–	0.275^{***} (0.059)	$\eta_{2s}^{(0)}$	-0.028^{***} (0.024)	–
$\psi_{2s}^{(2)}$	–	0.137^{**} (0.069)			
$\eta_{2s}^{(5)}$	-0.013^{*} (0.008)	–			
Conditional Variance Equation					
c_{1s}	1.128 (0.810)	0	c_{1s}	1.174^{***} (0.308)	0
c_{2s}	-0.567 (0.757)	1.183^{***} (0.421)	c_{2s}	0.881^{***} (0.172)	0.000001 (.382)
a_{1s}	0.502^{***} (0.094)	0.023 (0.047)	a_{1s}	0.422^{***} (0.093)	0.017 (0.041)
a_{2s}	-0.427^{*} (0.255)	0.506^{**} (0.251)	a_{2s}	-0.433^{***} (0.097)	0.116 (0.106)
b_{1s}	0.740^{***} (0.079)	0.013 (0.030)	b_{1s}	0.792^{***} (0.083)	0.002 (0.023)
b_{2s}	0.680^{*} (0.382)	0.382^{**} (0.185)	b_{2s}	-0.445^{***} (0.103)	0.828^{***} (0.061)
Loglik	-1274.35		*Loglik*	-1277.11	
$Q(6)$	17.970 [0.804]	$Q^2(6)$ 10.660 [0.968]	$Q(6)$	24.507 [0.432]	$Q^2(6)$ 16.166[0.760]
$Q(12)$	34.809 [0.922]	$Q^2(12)$ 30.903 [0.945]	$Q(12)$	39.705 [0.797]	$Q^2(12)$ 37.887[0.764]

Tests of No Volatility Transmission:		Tests of No Volatility Transmission:	
(i) Bidirectional between E_t and EF_t		(i) Bidirectional between E_t and BF_t	
$H_0: a_{12} = a_{21} = b_{12} = b_{21} = 0$	*LR*=5.61 [0.230]	$H_0: a_{12} = a_{21} = b_{12} = b_{21} = 0$	*LR*=13.44[0.009]
(ii) From E_t to EF_t		(ii) From E_t to BF_t	
$H_0: a_{12} = b_{12} = 0$	*LR*=0.62 [0.732]	$H_0: a_{12} = b_{12} = 0$	*LR*=0.36 [0.831]
(iii) From EF_t to E_t		(iii) From BF_t to E_t	
$H_0: a_{21} = b_{21} = 0$	*LR*=4.22 [0.120]	$H_0: a_{21} = b_{21} = 0$	*LR*=12.9 [0.001]

Notes: See notes to Table 5.2.

Table 5.7. The estimated bivariate GARCH–BEKK–in mean model for the UK.

Panel A. Exchange rates (E_t) and equity flows (EF_t)	E_t (s=1)	EF_t (s=2)	*Panel B. Exchange rates (E_t) and bond flows (BF_t)*	E_t (s=1)	BF_t (s=2)
Conditional Mean Equation					
μ_s	-0.060 (0.202)	0.239^{*} (0.139)	μ_s	-0.370^{*} (0.194)	1.334^{***} (0.185)
$\psi_{2s}^{(1)}$	–	0.186^{***} (0.054)	$\psi_{1s}^{(1)}$	–	0.342^{***} (0.109)
$\psi_{2s}^{(2)}$	–	0.096^{*} (0.051)	$\eta_{2s}^{(2)}$	-0.052^{**} (0.025)	–
$\psi_{2s}^{(3)}$	–	0.156^{***} (0.048)			
$\eta_{2s}^{(3)}$	-0.028^{*} (0.017)	–			
Conditional Variance Equation					
c_{1s}	0.659^{***} (0.146)	0	c_{1s}	0.290^{*} (0.160)	0
c_{2s}	-1.133^{***} (0.052)	0.00002 (0.272)	c_{2s}	0.173^{***} (0.041)	0.000002 (0.063)
a_{1s}	0.294^{***} (0.070)	0.032 (0.040)	a_{1s}	0.265^{**} (0.139)	0.070^{*} (0.040)
a_{2s}	-0.074 (0.154)	-0.226^{**} (0.097)	a_{2s}	-0.039 (0.067)	-0.001 (0.036)
b_{1s}	0.899^{***} (0.027)	0.023 (0.040)	b_{1s}	0.968^{***} (0.038)	-0.066^{***} (0.009)
b_{2s}	0.502^{***} (0.056)	0.468^{***} (0.003)	b_{2s}	0.324^{***} (0.088)	0.922^{***} (0.022)
Loglik	-1172.15		*Loglik*	-1078.10	
$Q(6)$	16.962 [0.850]	$Q^2(6)$ 8.996 [0.989]	$Q(6)$	21.022 [0.637]	$Q^2(6)$ 27.405[0.157]
$Q(12)$	40.318 [0.776]	$Q^2(12)$ 24.90 [0.993]	$Q(12)$	38.397 [0.837]	$Q^2(12)$ 39.612[0.698]

Tests of No Volatility Transmission:

(i) Bidirectional between E_t and EF_t

$H_0: a_{12} = a_{21} = b_{12} = b_{21} = 0$ *LR*=4.18 [0.381]

(ii) From E_t to EF_t

$H_0: a_{12} = b_{12} = 0$ *LR*=1.16 [0.559]

(iii) From EF_t to E_t

$H_0: a_{21} = b_{21} = 0$ *LR*=2.86 [0.238]

Tests of No Volatility Transmission:

(i) Bidirectional between E_t and BF_t

$H_0: a_{12} = a_{21} = b_{12} = b_{21} = 0$ *LR*=20.1[0.000]

(ii) From E_t to BF_t

$H_0: a_{12} = b_{12} = 0$ *LR*=33.7[0.000]

(iii) From BF_t to E_t

$H_0: a_{21} = b_{21} = 0$ *LR*=6.74[0.034]

Notes: See notes to Table 5.2.

As can be seen from the Tables, the dynamic interactions between exchange rate changes and net equity and net bond flows, captured by $\psi_{12}^{(i)}$ and $\psi_{21}^{(i)}$, suggest that there exist limited dynamic linkages between the first moments compared to the second ones. The results in the mean equation indicate the existence of bidirectional causality between exchange rate changes and net bond flows in Japan, causality from net bond flows to exchange rate changes in Canada and the UK, and from net equity flows to exchange rate changes in the euro area.

With regard to the impact of exchange rate uncertainty on net equity flows, the results suggest that exchange rate volatility affects net equity flows negatively in the euro area, Sweden, and the UK, and positively in Australia, and has no effect in Canada and Japan. Its impact on net bond flows, on the other hand, appears to be negative in all countries except Canada for which it is positive.

The observed negative impact on net equity as well as net bond flows has important implications. First, it indicates that risk averse market participants, especially those of the counterpart countries to the US, respond to exchange rate uncertainty by reducing their financing activities, hence favouring domestic rather than foreign securities in their portfolios to reduce their exposure to exchange rate volatility. In a broad sense, this finding is line with the evidence of Bayoumi (1990), Iwamoto and van Wincoop (2000), and Bacchetta and van Wincoop (1998; 2000). While Bayoumi (1990) showed that net capital flows as a share of GDP are lower during the floating exchange rate period (1965-1986) than during the gold standard (1880-1913), Iwamoto and van Wincoop (2000) provided evidence that net capital flows as a share of GDP are much larger across regions of a country, which use the same currency, than across countries. Bacchetta and van Wincoop (1998; 2000), on the other hand, showed that exchange rate uncertainty dampens net international capital flows by developing a two-period general equilibrium model.

Second, in contrast to Hau and Rey (2006) who assume that bonds are usually hedged instruments not affected by exchange rate uncertainty, it appears that uncertainty in fact affects bond as well as equity flows, and the former more widely, since a negative impact is found in five of the six countries considered. This is consistent with the results of Fidora et al. (2007), who found in a wide set of industrialised and emerging economies that exchange rate volatility is an important factor for bilateral portfolio home bias, this being higher for bonds than for equities. This finding has recently been confirmed by Bekaert and Wang (2009) and Borensztein and Loungani (2011), but the former study finds evidence that the effect of exchange rate volatility on home bias is economically insignificant. The rationalisation of Fidora et al. (2007) of the higher home bias for bonds compared to equities is that it is consistent with Markowitz-type international CAPM specifications in which less volatile financial assets should show larger home bias.

However, the above findings indicate that exchange rate volatility does not induce any type of home bias in Australia and Japan for equities and in Canada for both equities and bonds. Though the finding that exchange rate volatility positively affects net equity flows in Australia is consistent with that of Batten and Vo (2010) and Daly and Vo (2013), who found that exchange rate volatility reduces equity home bias in Australia as opposed to Mishra (2011), who provided evidence that the Australian investors invest less in the equity market of a country if the real exchange rate of that country is volatile. A possible explanation for the findings of Australia and Canada may be due to the fact that these countries are commodity-exporting countries and that financial market developments in these countries are driven by the terms of trade shocks. Chaban (2009) and Ferreira Filipe (2012) indeed found that the portfolio rebalancing motive of Hau and Rey (2006) in these countries is weak. Chaban (2009) argued that commodity prices play a significant role in the transmission of shocks in these countries, while Ferreira Filipe (2012) found that differences in country-specific shocks volatility are also at play in these countries. Japan is also a special case, its finding may be due to what was highlighted by Hau and Rey (2006) in that international portfolio flows with Japan regards only bonds as opposed to equities, even though a high percentage of Japanese debt is finance internally.

The estimates of the conditional variance equations indicate that exchange rate changes (net equity/bond flows) exhibit conditional heteroscedasticity: the diagonal elements of the ARCH matrices are significant at the 10% level in all cases except for net equity flows in Australia and net bond flows in Australia, Sweden, and the UK. Furthermore, the conditional variances exhibit persistence in all cases except for net equity flows in Canada. While the persistence of the conditional variances of exchange rate changes ranges from 0.54 (Japan) to 0.98 (euro area), the persistence of the corresponding conditional variances of the flows ranges from 0.38 (Sweden) to 0.91 (euro area) for net equity flows and from 0.43 (Japan) to 0.98 (Canada) for net bond flows.

The ARCH, a_{11}, and GARCH, b_{11}, estimates of exchange rate changes in the bivariate GARCH–BEKK–in mean models are rather similar, regardless of whether the relationship with net bond or net equity flows is considered (see Panels A and B respectively in all Tables). More specifically, the change in a_{11} is less than 10% and this also applies to b_{11}, except for Japan where the change is around 26%. Furthermore, the off-diagonal elements of the ARCH and GARCH matrices indicate that shocks to exchange rate changes (net equity flows) affect the conditional variance of net equity flows (exchange rate changes) at the 10% level in the euro area and Japan. The results also show that shocks to exchange rate changes (net bond flows) affect the conditional variance of net bond flows (exchange rate changes) at the 10% level in all cases except Japan.

More specifically, the causality-in-variance (i.e., the information flows) tests based on likelihood ratio test statistics provide evidence of strong causality-in-variance from net equity flows to exchange rate changes in the case of the euro area and bidirectional causality-in-variance in the case of Japan. There is also causality-in-variance from net bond flows to exchange rate changes in Australia, the euro area, and Sweden, as well as bidirectional causality in Canada and the UK.

A possible explanation for the existence of stronger dynamic linkages in terms of the second moment between exchange rate changes and net bond flows rather than net equity flows is that foreign exchange dealers usually follow bond yields in their trading behaviour, with such yields, in turn, driving cross-border bond acquisitions, which results in volatile exchange rates. Spillovers from the exchange rates, on the other hand, may due to the fact that investors adjust their portfolios on the basis of their volatility. Also, the limited linkage in the second moment between exchange rate changes and net bond flows in Japan can be explained by the fact that a high percentage of Japanese debt is financed internally, primarily by Japanese pension funds, hence the volatility of the net bilateral bond flows between the US and Japan has no impact on exchange rate volatility, and vice versa.

Finally, the Hosking multivariate *Q*-statistics of order (6) and (12) for the squared standardised residuals suggest that the multivariate GARCH (1, 1) structure is sufficient to capture the volatility in the series.

5.5. Conclusions

In this chapter, we have analysed the impact of exchange rate uncertainty on net bond and net equity flows, as well as the dynamic linkages between exchange rate volatility and the variability of these flows, using data for the US *vis-à-vis* six advanced economies, namely Australia, the UK, Canada, Japan, Sweden, and the euro area over the period 1988:01-2011:12. By estimating bivariate GARCH–BEKK–in mean models, we find evidence that exchange rate volatility impacts on net equity flows negatively in the euro area, Sweden, and the UK and positively in Australia. Furthermore, in contrast to Hau and Rey (2006), it also affects net bond flows negatively in all countries except Canada where the effect is positive. The general conclusion that can be drawn from these results is that exchange rate volatility induces risk averse investors, especially those of the counterpart countries to the US, to reduce their international financing activities and to favour domestic to foreign assets in their portfolios in order to minimise their exposure to volatility. This evidence is strong in the cases of the UK, the euro area and Sweden as opposed to Canada, Australia and Japan. Though, the results of Australia, Canada and Japan may be due to the specific characteristics of these economies, consistent with earlier studies documented in the literature (e.g., Hau and Rey, 2006; Chaban, 2009; and Ferreira Filipe, 2012).

The causality-in-variance analysis suggests the existence of strong causality-in-variance from net equity flows to exchange rate changes in the euro area and bidirectional causality-in-variance in Japan. As for the linkages between exchange rate changes and net bond flows, causality-in-variance from net bond flows to exchange rate changes is found for Australia, the euro area, and Sweden, and bidirectional causality-in-variance is observed for Canada and the UK. These findings have important policy implications, since they suggest that policy-makers and economic and financial regulators could use the exchange rate or credit controls on equity as well as bond flows as instruments to achieve economic and financial stability.

CHAPTER SIX

CONCLUDING REMARKS

This thesis is a contribution to the literature on exchange rates and international finance. The relevant literature is very extensive and has evolved in several directions. One line of literature includes the macroeconomic models of exchange rate determination, which have been proposed ever since the floating exchange rates in 1973. See, for examples, the flexible-price monetary model (e.g., Frenkel, 1976; Bilson, 1978), the sticky-price monetary model (Dornbusch, 1976), the real interest rate differential monetary model (Frankel, 1979), the portfolio balance models (e.g., Branson, 1976; Dooley and Isard, 1979), and the general equilibrium models of Stockman (1980) and Lucas (1982). This literature also includes the ongoing empirical studies that examine the empirical validity of these models, the PPP, the UIP, among many other specifications of exchange rate determination.

Another line of literature considers the firms' foreign exchange exposure, as well as the dynamic linkages between stock prices and exchange rates. While the firms' foreign exchange exposure has a unified theoretical appeal in that exchange rate movements are an essential source of macroeconomic uncertainty and hence are likely to have a significant effect on the value of the firm, the empirical evidence is mixed (see Muller and Verschoor, 2006b, for a thorough review). As far as the dynamic linkages between stock market prices and exchange rates are concerned, both the theoretical and empirical literature are inconclusive. Though, the empirical research shows that it is likely that stock prices lead exchange rates during turbulent periods (see Granger et al., 2000; Caporale et al., 2002; Tsagkanos and Siriopoulos, 2013; among others).

The literature on international finance also includes the micro-founded models, which have focused on the impact of the currency order flows on exchange rate variations (e.g., Evans and Lyons, 2002; 2005; 2008; Payne, 2003; Rime et al., 2010; Chinn and Moore, 2011; and Duffuor et al., 2012; among others). Prompted by the view that currency order flows and portfolio flows represent the investors' behaviour, and hence they are closely interrelated, a growing literature has focused on the impact of equity and bond portfolio flows on exchange rate dynamics (e.g., Hau and Rey, 2006; Brooks et al., 2004; Siourounis, 2004; Chaban, 2009; and Kodong and Ojah, 2012; among others).

An ongoing strand of literature also focuses on the impact of exchange rate volatility. This literature has mainly examined the impact of exchange rate volatility on the macroeconomic variables, such as trade flows and economic growth. See McKenzie (1999) or Bahmani-Oskooee and Hegerty (2007) for a comprehensive review of the theoretical and empirical literature.

In this thesis, we consider the exchange rate determination issue, with a particular focus on the monetary and portfolio choice models. In particular, we examine the monetary approach of exchange rates, the dynamic linkages between stock prices and exchange rates during the recent financial crisis, and the role of net portfolio flows in exchange rate changes. The thesis also addresses the impact of exchange rate uncertainty. However, rather than on trade flows, we focus on the impact of exchange rate uncertainty on cross-border equity and bond portfolios flows. Throughout the thesis, we use a wide range of time series models. These models include cointegration techniques, multivariate GARCH specification, multivariate GARCH-in-mean specification, and the Markov regime-switching specifications.

Chapter 2 puts under econometric scrutiny two models of exchange rate determination: the flexible-price monetary model of Bilson (1978) and the real interest rate differential model of Frankel (1979). Although the models have important analytical applications, they have provided limited support when examined empirically. As the US dollar-Japanese yen exchange rate was much debated over last few decades, we consider such an exchange rate in the analysis using quarterly data over a period of high international capital mobility between the US and Japan, 1980:01-2009:04.

Using the Johansen cointegration technique, we trace both money demand instability and real exchange rate persistence arguments to authenticate, *inter alia*, that the limited success of the monetary models of the dollar-yen exchange rate is the result of the breakdown of their underlying building blocks. In particular, adjusting the models to factors that are at play in capturing the stability of money demands such as real stock prices, on the one hand, and factors explaining the persistence of the real exchange rate such as productivity differential, relative government spending and real oil price, on the other, result in a notable improvement of such models. The importance of the modified monetary models is also pinpointed by the results of the long-run weak exogeneity tests. Considering the real interest rate differential model and its counterpart modified version, it is shown that the cumulated shocks to the nominal exchange rate are sourced from the shocks of factors affecting the conventional monetary model's building blocks (e.g., relative real equity prices and productivity differential). The results further confirmed that the modified monetary models outperform the random walk benchmark in out-of-sample forecasting in the medium- and long-term, but not the short-term.

Chapter 3 investigates the interrelationship between stock prices and exchange rates during the banking crisis of 2007-2010. Weekly data from six developed countries (the US, the UK, Canada, Japan, the euro area and Switzerland) are analysed over two sub-periods: the pre-crisis period (August 6, 2003-August 8, 2007) and the crisis period (August 15, 2007-December 28, 2011). The analysis is conducted by using a bivariate VAR-GARCH (1, 1) in the BEKK representation of Engle and Kroner (1995). The adopted specification accounts for the existence of cointegration between the levels of stock prices and exchange rates. Also, it allows for the time-varying conditional correlation and for causation in the variances in a lead-lag framework.

The empirical evidence shows the existence of Granger causality from stock returns to exchange rate returns in the US and the UK, in the reverse direction in Canada, and of bidirectional causality in the euro area and Switzerland during the recent financial crisis. The causality-in-variance results, on the other hand, show the existence of strong causality-in-variance from stock returns to exchange rate returns in Japan and in the reverse direction in the euro area and Switzerland, whereas the US and Canada show evidence of bidirectional causality-in-variance during the period. The findings are in line with those of Tsagkanos and Siriopoulos (2013) who examined the dynamic linkages between stock prices and exchange rates during the recent financial crisis and also with those of Granger et al. (2000) and Caporale et al. (2002) in examining the linkages between the two variables during the 1997 Asian financial crisis.

The findings imply that capital flows into and out of the considered economies seem to have played an important role in the interactions between the two variables during the recent financial crisis. The results also indicate important practical implications for investors. Since stock and foreign exchange markets show strong linkages during the financial crisis, investors have limited opportunities for portfolio diversification during the period.

Even though the linkages between the financial markets during the recent financial crisis have drawn much attention in the literature, as stated earlier, further investigations are in fact of paramount interest. Future works might examine the linkages between different types of financial markets such as stock, foreign exchange, and bond markets during the global financial crisis to provide evidence how these markets have been interrelated during the period. This topic is left for future research.

Chapter 4 investigates the impact of such different components of net portfolio flows as net equity and net bond flows on exchange rate dynamics. In particular, the chapter contributes to the existing literature by examining the nonlinear relationship between exchange rate changes and these two types of flows, using quarterly data for the US against the UK, the euro area, Japan, and Canada over the period 1990:01-2011:04. Using two-state Markov switching models, the empirical findings show the existence of state-dependent linkages between exchange rate changes and net portfolio flows for all countries, except Canada.

In particular, the results provide evidence that net equity and net bond inflows from the UK towards the US lead to an appreciation of the US dollar against the British pound in the appreciation regime. Furthermore, the results suggest that net bond inflows from the euro area towards the US result in a US dollar appreciation (depreciation) against the euro in the low (high) volatility regime. By contrast, net bond inflows from Japan towards the US imply an appreciation of the US dollar against the Japanese yen in the less volatile state. The results also show that neither net equity flows nor net bond flows have statistically significant effects on the US dollar-Canadian dollar exchange rate, consistent with previous studies on commodity-exporting economies (see Chaban, 2009; Ferreira Filipe, 2012).

The results presented in the chapter are a first cut and further analyses are no doubt needed. Suggestions for future works might include examining the out-of-sample forecasting ability of the estimated models in the chapter and comparing the ability of the competing nonlinear models. Investigating the role of portfolio flows in exchange rate changes of developing and emerging countries is also of paramount interest. These issues are beyond the scope of the chapter, but constitute interesting topics for future research.

Chapter 5 contributes to the existing literature by examining the effect of exchange rate volatility on net equity and net bond portfolio flows across borders. The idea is that exchange rate uncertainty reduces international portfolio flows as a result of high transaction costs and low potential gains from international diversification. Hence, holding foreign securities such as bonds and equities becomes more risky (Eun and Resnick, 1988). The empirical analysis is conducted by using a bivariate GARCH-BEKK-in mean models and monthly bilateral data for the US *vis-à-vis* Australia, the UK, Japan, Canada, the euro area, and Sweden over the period 1988:01-2011:12.

The findings show that the effect of exchange rate volatility on net equity flows is negative in the euro area, the UK and Sweden, and positive in Australia, whilst two countries (Japan and Canada) show insignificant responses to volatility. With regard to net bond flows, the results show that net bond flows are affected negatively in all countries except Canada, where such flows are affected positively. A possible explanation for these results is that, presuming investors are risk averse, US dollar exchange rate volatility induces investors, primarily those of the counterpart countries to the US, to lower their international financing activities and favour domestic to foreign assets in their portfolios in order to maximise their returns and minimise their exposure to volatility. This evidence is strong in the cases of the UK, the euro area and Sweden as opposed to Canada, Australia and Japan. Though, the results of Australia, Canada and Japan may be due to the specific characteristics of these economies, consistent with earlier studies in the literature (e.g., Hau and Rey, 2006; Chaban, 2009; and Ferreira Filipe, 2012).

Furthermore, the results show the existence of strong linkages between exchange rate volatility and the flows variability. Hence, economic and financial regulators can use the exchange rate or credit controls on these two types of flows as intervention channels to pursue the economic and financial stability.

As the results presented in the chapter are the newest addition to the relevant literature, further analyses are no doubt required. Future works might examine the impact of exchange rate uncertainty on equity and bond portfolio inflows and outflows separately using data not only from developed countries but also from developing and emerging ones. These issues constitute interesting topics and are left for future research.

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