

This work is used with the permission of Bangladesh Institute of Development Studies.

© 2003, Bangladesh Institute of Development Studies.

**MIMAP-Bangladesh
Micro Impacts of Macroeconomic and Adjustment
Policies in Bangladesh**

**Technical Paper No. 05
The Effect of Exchange Rate Changes on Output:
An Empirical Analysis on Bangladesh**

Mohammad A. Razzaque 100713

Lecturer, Department of Economics, University of Dhaka, Bangladesh.

MIMAP Technical Papers contain preliminary material and research results and are circulated provisionally in order to stimulate discussion and critical comment. It is expected that the content of the Technical Papers may be revised prior to their eventual publication in some other form.

ARCHIVE
11/11/03

MIMAP-Bangladesh Technical Paper No. 05

The Effect of Exchange Rate Changes on Output: An Empirical Analysis on Bangladesh

This work was carried out with the aid of a grant from the International Development Research Centre, Ottawa, Canada.

The materials presented and the opinion expressed in this Publication are those of the authors and do not necessarily reflect those of BIDS and the International Development Research Centre (IDRC).

January 2003
BIDS

Published by:
Bangladesh Institute of Development Studies (BIDS)
E-17 Agargaon, Sher-e-Bangla Nagar
Dhaka – 1207
Bangladesh.
Tel: 9143441-8
Fax: 880-2-8113023
E-mail: dg_bids@sdbnd.org
Web site: www.bids-bd.org

Contents

	<u>Page</u>
1. Introduction	1
2. Theoretical Framework	4
3. Construction of RER for Bangladesh	7
3.1 Real Exchange Rate: Definition and Measurement	7
3.2 Bilateral and Multilateral RERs for Bangladesh	9
4. Empirical Specification and Data	16
4.1 The Empirical Model	16
4.2 Data	17
5. Estimation Strategy	18
5.1 Time series properties of the variables	18
5.2 Cointegration and Error Correction Modelling	20
5.2.1 The Engle-Granger Procedure	20
5.2.2 The Phillips-Hansen Fully Modified OLS	21
5.2.3 Existence of a Long-run Relationship	24
5.2.3.1 Testing for Cointegration	24
5.2.3.2 Testing Cointegration for Variables with Different Order of Integration: The Pesaran et al. Test	24
6. Estimation Results	27
6.1 Examining the Time Series Properties of Variables	27
6.2 Estimating the Long-run Relationship	31
7. Summary of Findings and Implications	38
References	42

List of Tables

	<u>Page</u>
Table 1	13
Table 2	27
Table 3	32
Table 4	35
Table 5	35
Table 6	36
Table 7	37

List of Figures

Figure 1	14
Figure 2	14
Figure 3	15
Figure 4	29
Figure 5	31
Figure 6	34
Figure 7	34

The Effect of Exchange Rate Changes on Output: An Empirical Analysis on Bangladesh

Mohammad A. Razzaque^{*}

1. Introduction

Devaluation of the domestic currency has been an important component of the orthodox stabilisation programme leading to trade policy reforms. By raising the domestic currency price of foreign exchange devaluation increases the price of traded goods relative to non-trade ones. This causes a reallocation of resources resulting in increased production of exports and items of import competing sectors.¹ Devaluation is also believed to contribute to the enhancement of external competitiveness of the country allowing exporters to cut their product prices in foreign currency in overseas markets. Increased competitiveness should stimulate production in the export sector further. On the other hand, as a direct consequence of nominal devaluations import prices go up, which is likely to depress the demand for imports in the domestic economy. Increased exports and reduced imports are expected to improve the external trade balance of the country in question and many developing countries have relied upon devaluation to correct fundamental disequilibria in their balance of payments.² It is argued that by expanding the production of the traded sector in general, and exports in particular, devaluation should have an expansionary effect on the overall economy.

However, although nominal devaluations help achieve the goal of relative price adjustment along with an improvement in trade balance, they might do so at a high cost. There are concerns that

^{*} Lecturer, Department of Economics, University of Dhaka, Dhaka 1000, Bangladesh. For comments and suggestions please e-mail at: razzaque_duecon@yahoo.com.

¹ This is also known as expenditure-switching effect of devaluation.

² According to the economic theory devaluation will improve the balance of payments if the Marshall-Lerner condition holds, i.e., the devaluing economy's demand for imports and demand for its exports are elastic. In fact, Marshall-Lerner condition specifies that the sum of both elasticities, in absolute term, must be greater than one.

indirect costs of devaluation can actually outweigh its benefits adversely affecting the overall output growth. This is what is known as the contractionary effect of devaluation.

There are a number of theoretical reasons for a contractionary effect of devaluation. First, devaluation increases the price of traded goods, which feeds into the general price level rendering a negative real balance effect. This, in turn, will result in lower aggregate demand and output (Edwards, 1986). Second, the contractionary effect might also result from income distributional effect of devaluation. This point was first mentioned by Diaz-Alejandro (1963) who argued that devaluation could lead to a redistribution of income from people with high marginal propensity to consumption to high propensity to save rendering a negative effect on the aggregate demand. Third, if the demand for imported goods is inelastic due to the dominance of capital and essential intermediate and consumers' goods in a country's import basket, then devaluation may be contractionary (Upadhaya and Upadhaya, 1999). Apart from these demand side channels, contractionary effects can also arise from the supply side (Edwards [1986], Upadhaya and Upadhaya [1999]). The increased cost of imported inputs might affect production and output adversely. Thus, while Hanson (1983) emphasized on the importance of imported inputs even in the production of non-traded goods, Lizondo and Montiel (1989) maintained that reduced profits in the non-traded sector caused by increased costs of imported inputs (e.g., oil) led to contraction in aggregate supply after devaluation. Besides, the real balance effect of devaluation might raise the interest rate thus reducing the demand for working capital by the firms. Krugman and Taylor (1978), using a Keynesian framework, have identified certain conditions under which devaluations are found to be contractionary, viz., (1) if initially imports exceed exports, (2) consumption propensity out of profits and wages are different, and (3) if government revenues are increased as a result of devaluation.

Empirical findings on the consequences of devaluations on output are mixed. While Gylfason and Schmid (1983), Connolly (1983) provided some support for expansionary devaluations, Gylfason and Radetzki (1985) and Atkins (2000) encountered with contractionary effects of devaluation. Some interesting results are reported in Edwards (1986) and Rhodd (1993) where the authors found negative short-run effects but in the long-run the output response to devaluation appeared to be positive. Finally, there are other studies that do not find any

significant effect on devaluation (e.g., Bahmani-Oskooee, 1998 and Upadhyaya and Upadhyay (1999).

Like many other developing countries, devaluation has been a major component of trade policy instruments in Bangladesh. Since the initiation of trade liberalisation programmes in the early 1980s, the country has adopted a policy of frequent but small doses of devaluation at a time. Every time the government devalues the taka, it emphasises on enhancing the competitiveness of exports as one of the most important reasons for justifying the action.³ However, in light of the above discussions, it is legitimate to ask what has been the net effect of exchange rate changes on Bangladesh's overall economic activity? As gathered from various studies, it is not possible to generalise the experience of developing countries and consequently nothing can be inferred about Bangladesh in the absence of an empirical research. While the issue of devaluation attracts so much attention in Bangladesh, discussions surrounding it are usually uninformed in nature due to lack of an in-depth study. The present paper therefore contributes to the macro policy discourse on Bangladesh more effectively by carrying out an empirical investigation into the impact of devaluation on output. Another important aspect of this study is the use of the new national income estimate recently compiled by the Bangladesh Bureau of Statistics (BBS, 2001). The new national income accounting system provides a wider coverage than the old method of estimating the GDP and hence is expected to capture such macroeconomic shocks as exchange rate changes more adequately.

The paper is organized as follows: after this introductory discussion, Section 2 provides the analytical framework specifying the output-exchange rate link. While Section 3 defines and constructs multilateral real exchange rate series for Bangladesh, Section 4 provides the empirical specification of the theoretical model and elaborates the sources of data. In Section 5 we devise our estimation strategy and estimation results are presented in the subsequent section. Finally, Section 7 contains a summary of findings of the study and some ensuing implications.

³ Containing trade deficit and encouraging increased inflow of remittances through official channels are two other major objectives of nominal devaluations.

2. Theoretical Framework

Edwards (1986), who builds on the analytical framework of Khan and Knight (1981) that attempted to analyse the effects of stabilization programmes on aggregate production in developing countries, has been the most influential work in guiding the empirical analysis on the effects of devaluation on output. Another useful framework is due to Rhodd (1993) who uses a simple three market Keynesian model to illustrate the relationship between output and the exchange rate. For the present paper, we will consider this model as the basis for theoretical understanding and subsequently empirical estimation will be based on the reduced form equation for real output derived thereof. In Rhodd's model the goods market is represented by:⁴

$$Y = C + I + G + X - M \quad (2.1)$$

Or, $Y - C - G = I + X - M \quad (2.2)$

$$S = I_d + I_f \quad (2.3)$$

$$S = S(Y, r); S_Y > 0, S_r > 0 \quad (2.4)$$

$$I_d = I_d(Y, r); I_{dY} > 0, I_{dr} < 0 \quad (2.5)$$

$$I_f = I_f(Y, e); I_{fY} < 0, I_{fe} > 0 \quad (2.6)$$

$$S_Y = \frac{\partial S}{\partial Y}, \quad S_r = \frac{\partial S}{\partial r}, \quad I_{dY} = \frac{\partial I_d}{\partial Y}, \quad I_{fY} = \frac{\partial I_f}{\partial Y}, \quad I_{fe} = \frac{\partial I_f}{\partial e}$$

where, total expenditure, consumption expenditure, domestic investment expenditure, savings, government spending, net exports or foreign investment (I_f), domestic interest rate and exchange rate are represented respectively by Y , C , I , S , G , $X-M$, r , and e . Equation (2.3) shows the equilibrium between aggregate demand and aggregate supply. Equations (2.4), (2.5) and (2.6) show how S , I_d and I_f are determined in the model. Foreign investment (I_f), which defines the net build-up of claims on the rest of the world or $(X-M)$ is expected to vary inversely with domestic income, Y , and directly with the exchange rate (e). As Y increases, imports increase and $X-M$ worsens. An increase in e or nominal devaluation causes the trade balance to increase.

⁴ This section draws on Rhodd (1993).

Considering the money market, the equilibrium requires the balancing of money demand and money supply. Money supply is to be determined by monetary policy, while money demand is determined by income and interest rate.

$$M_{s0} = M_d \quad (2.7)$$

$$M_d = L(Y, r); \quad L_Y = \frac{\partial M_d}{\partial Y} > 0, \quad L_r = \frac{\partial M_d}{\partial r} < 0 \quad (2.8)$$

The third and the final market in the model is the foreign exchange market, which gives the equilibrium of the demand for foreign exchange against its supply. Under a fixed exchange rate regime, the balance of payments is to be influenced by trade flows and financial flows where the former are determined by Y while the latter by r . According to Rhodd (1993), the greater the level of income the worse the trade balance. Although capital flows can improve trade balance in the short-run, the long run effect is not known due to loan repayment and repatriation of dividends and interest.

$$B = T(Y) + F(r) \quad (2.9)$$

$$\frac{\partial B}{\partial Y} < 0, \quad \frac{\partial B}{\partial r} = ? \quad (2.10)$$

To facilitate the solution of the model algebraically the equilibrium conditions can be written in linear form as given in (2.11-2.13).

$$S_0 + S_1 Y + S_2 r - I_{d0} - I_{d1} Y - I_{d2} r - I_{f0} - I_{f1} Y - I_{f2} R = 0 \quad (2.11)$$

$$L_0 + L_1 Y + L_2 r = M_s \quad (2.12)$$

$$T_0 + T_1 Y + T_2 e + F_0 + F_1 Y + F_2 r - B = 0 \quad (2.13)$$

Equations (2.11) – (2.13) can be written in matrix form to give:

$$\begin{bmatrix} (S_1 - I_{d1} - I_{f1}) & (S_2 - I_{f2}) & 0 \\ L_1 & L_2 & 0 \\ (T_1 + F_1) & F_2 & -1 \end{bmatrix} \begin{bmatrix} Y \\ r \\ B \end{bmatrix} = \begin{bmatrix} -S_0 + I_{d0} + I_{f0} + I_{f2}e \\ M_s - L_0 \\ T_0 - T_2e - F_0 \end{bmatrix} \quad (2.14)$$

From (2.14) Y can be determined, which is given by:⁵

$$Y = \frac{(L_2 S_0 - L_2 I_{d0} - L_2 I_{f0} - L_2 I_{f2} e + M_s S_2 + I_{f2} M_s) - (S_2 L_0 + I_{f2} M_s)}{D} \quad (2.15)$$

Where $D = (S_1 - I_{d1} - I_{f1})(L_2)(-1) - (-1)(L_2)(S_2 - I_{f2}) > 0$

$$\frac{\partial Y}{\partial e} = -\frac{L_2 I_{f2}}{D} > 0 \quad (2.16)$$

$(L_2 < 0, I_{f2} > 0, D > 0)$

Therefore, the empirical model will be based on equation (1.15) that shows the relationship between real output, a measure of monetary policy as indicated by M_s , the exchange rate, and government expenditure, which is included in the saving-investment identity. By including fiscal and monetary measures Rhodd's model shows that a devaluation is not undertaken by itself but is associated with other policy measures.

⁵ Cramer's rule can be used to determine Y.

3. Construction of RER for Bangladesh

3.1. Real Exchange Rate: Definition and Measurement

Since the discussions on the effects of devaluation on output usually consider the changes in the ‘real’ rather than the ‘nominal’ exchange rate, at the outset, it is useful to define the term *RER*. Despite its importance there is a great deal of confusion over the definition and measurement of *RER*. The two main strands of *RER* are purchasing power parity (*PPP*) and trade theory definitions. The *PPP* theory is based on the observation that exchange rate movements are determined by the difference between the domestic and foreign rates of inflation. If domestic inflation is higher than that of the foreign rate, the exchange rate will appreciate, and vice versa. Thus, the *RER* is defined as the ratio of foreign prices (P_f) to domestic prices (P_d) adjusted for the nominal exchange rate (local currency per unit of foreign currency) (E), that is, $RER^* = E(P_f/P_d)$, where RER^* is the *PPP RER*.⁶ On the other hand, the trade theory definition is derived from the dependent economy type model (e.g., Salter-Swan model), where *RER* is defined as the ratio of price of tradables (P_T) to non-tradables (P_{NT}), or, $RER = (P_T/P_{NT})$. A fall in *RER*, or a real appreciation, indicates an increase in the domestic cost of producing tradables reflecting the worsening of a country’s competitiveness. Conversely, an increase in *RER*, or a real depreciation, represents an improvement in a country’s international competitiveness.

Modern theoretical and empirical studies, however, have most frequently used the trade theory definition of *RER* (e.g., Dornbusch, 1980; Edwards, 1988, 1989 and 1994; Elbadawi, 1994;

⁶ In fact, the *PPP* theory is often postulated in terms of nominal exchange rate, which has two variants. The strong or absolute hypothesis postulates that the exchange rate between two countries should equal the ratio of the price levels in these countries. If R is the nominal exchange rate, P_d and P_f are the price levels in the home and foreign country respectively, the strong version of *PPP* can simply be written as: $R = P_f/P_d$. On the other hand, the relative or weak version of *PPP* specifies that the exchange rate should bear a constant proportionate relationship to the ratio of national price levels. That is: $R = \theta(P_f/P_d)$, where θ is a constant scalar. Both versions of *PPP* suggest that any change in price levels should result in equi-proportional change in exchange rate (Isard, 1995). For example, logarithmic transformation of relative version would yield: $\ln R = \theta + \ln P_f - \ln P_d$, and under the absolute version θ will be equal to zero. However, if we consider a change in the exchange rate then $\Delta \ln R = \Delta \ln P_f - \Delta \ln P_d$. That is, under either version of *PPP* a change in price ratio will call for an equi-proportionate change in exchange rate. In estimates of equations of the form: $\ln R = \theta + \beta_1 \ln P_f + \beta_2 \ln P_d + \nu_t$, a test of the restrictions $\beta_1 = 1, \beta_2 = -1$ would be interpreted as a test of the relative *PPP*, while the test of the same restrictions applied to the equation with the variables in first differences would be interpreted as a test of the absolute *PPP*. The *PPP* theory is, however, frequently restated in terms of the real exchange rate.

Elbadawi and Soto, 1996; and Montiel, 1999a). This is because although both RER^* and RER provide indices of competitiveness, the latter also gives information on the domestic incentive structure and the consequent resource allocation between tradable and non-tradable sectors. The other important reason for wider acceptance of the trade theory definition is that developing countries are likely to fit the Salter-Swan model quite well since these countries have a significant number of non-traded goods and they are typically ‘small’ countries that can hardly affect the world price of the traded goods (White and Wignaraja, 1991). Following the seminal contribution of Edwards (1988 and 1989), the trade theory definition of RER has become most popular in empirical research on developing countries. In this chapter we will also use the trade theory definition.

Even though the trade theory definition is useful for analytical purposes, it might be difficult to compute the RER for a number of reasons. Firstly, data on prices of tradables and non-tradables are virtually non-existent and the construction of such indices for developing countries is extremely difficult. As a result, Edwards (1989) suggested using the world price of tradables (P_T^f) as a proxy for P_T , and domestic price of non-tradables (P_{NT}^d) for P_{NT} . Thus, equation (1) can be considered as the operational definition of RER .

$$RER = E \left[\frac{P_T^f}{P_{NT}^d} \right] \quad (3.1)$$

Equation (3.1), however, does not really solve the problem, as it is not possible to find an exact empirical counterpart to the above analytical construct. This means we will still have to use proxies for the two price indices. Following Harberger (1986) and Edwards (1989) it has now become an established practice to use the foreign whole sale price index (WPI^f) as a proxy for P_T^f and the domestic consumers’ price index (CPI^d) for P_{NT}^d . Studies that have used these two proxies include, Baffes, *et al.* (1999), Cottani *et al.* (1990), Domac and Shabsigh (1999), Dorosh and Valdes (1990), Elbadawi (1994), Elbadawi and Soto (1996), and White and Wignaraja

(1992). In the present chapter these two proxies will also be used for constructing a RER series for Bangladesh.⁷

The next issue is whether to construct a bilateral or multilateral *RER*. It can be argued that a multilateral *RER* provides a better index of competitiveness. This is because even if the taka depreciates against one currency it might appreciate against others. Edwards (1989) provides evidence that the bilateral and multilateral indexes can move in opposite directions and he advocates for constructing a broad multilateral index of real exchange rate.

Since we need to construct a multilateral *RER* for Bangladesh, two further issues are: which countries are to be considered and what weighting system should we use? The usual practice in this regard is to consider the important trade partners with weights (α) equal to the share of each partner in the country's trade transactions (i.e. either exports or imports or both). Hence, (2) gives the basic formulation of the multilateral *RER*, where subscript t denotes time.

$$RER_t = \left[\sum_{i=1}^k \alpha_i \frac{E_{it} WPI_t^f}{CPI_t^d} \right] \text{ with } \sum_{i=1}^k \alpha_i = 1 \quad (3.2)$$

3.2. Bilateral and Multilateral RERs for Bangladesh

In order to construct the RER, initially 25 most important countries in terms of total value of trade transactions (i.e. exports plus imports) with Bangladesh were chosen.⁸ However, due to unavailability of the data on foreign prices China, Hong Kong, and the United Arab Emirates had to be dropped.⁹ Indonesia and Iran were excluded because of abrupt and dramatic changes in

⁷ Note that in practical terms the difference between RER^* and RER is then minimal. The RER^* uses the ratio of the same foreign to domestic price index (usually CPI), whereas the RER uses two different indices. However, as the WPI and CPI indices are usually highly correlated, RER^* and RER are likely to be very similar.

⁸ Data on Bangladesh's exports to and imports from different countries have been obtained from various issues of the Direction of Trade Statistics of the IMF.

⁹ For China and Hong Kong prices were available only for a short period of time. For UAE no information on either CPI, WPI or GDP deflators could be found.

their nominal exchange rates with respect to the US dollar.¹⁰ Appendix 1 gives Bangladesh's bilateral formal trade with the remaining 20 partner countries for 1980-2000.

There is no clear-cut rule about whether to use the trade share of partners in a reference year or the average share over a longer time horizon. Whilst Edwards (1989) has worked out weights on the basis of one reference year, Dorosh and Valdes (1990) have used average share for a relatively longer time span. In this study, we compute the *RER* for Bangladesh on the basis of partners' average trade shares during 1995-2000.¹¹ The weights thus obtained are: the US (0.218), India (0.142), Japan (0.093), Germany (0.080), the UK (0.069), Singapore (0.064), Korea (0.050), France (0.046), Italy (0.037), Netherlands (0.036), Belgium (0.024), Australia (0.022), Canada (0.021), Malaysia (0.017), Pakistan (0.017), Thailand (0.017), Denmark (0.012), Switzerland (0.012), Spain (0.010) and Sweden (0.010).

It must be stressed here that when weights are used on the basis of the recorded or official trade, the share of India is actually underestimated, as there is an astronomical amount of informal (illegal) border trade between Bangladesh and India. A survey of all border check-posts by Bakht (1996) found that the informal border trade with India in 1996 was about 50 percent of official trade with India. Again, based on a study by Rahman and Razzaque (1998), World Bank (1999) has suggested that the value of all such trade could be at least as high as official trade with India (i.e., about US\$ 0.8 billion). Despite the fact that reliable estimates of informal trade are hard to obtain for every year, considering the widespread availability of Indian goods it might not be unrealistic to assume that the volume of informal trade with India is as big as the volume of formal trade with her. Since the *RER* is also used as an indicator of competitiveness, it is important not to underestimate the role of such a large neighbour country as India. Therefore, we decided to compute another series of *RER* with weights derived from an enhanced bilateral trade with India by 100 per cent. At this, the weights were re-estimated to be: India (0.248), the US (0.191), Japan (0.082), Germany (0.04), the UK (0.061), Singapore (0.056), Korea (0.044), France (0.04), Italy (0.03), the Netherlands (0.032), Belgium (0.021), Australia (0.020), Canada

¹⁰ In the case of Iran, the exchange rate (rials per US dollar) rose sharply from 92.3 in 1992 to 2415 in 1993. On the other hand, Indonesia staged massive devaluations of its domestic currency in response to the so-called 'East Asian crisis'.

(0.019), Malaysia (0.015), Pakistan (0.015), Thailand (0.015), Denmark (0.011), Switzerland (0.011), Spain (0.009) and Sweden (0.009).

The data on *WPI* of partners and *CPI* of Bangladesh were gathered from IMF (2001).¹² However, there is no information available on Bangladesh's nominal exchange rates vis-à-vis all partners. This required us to modify the *RER* formula so that each partners' exchange rate with the US dollar could be used to construct the index (Dorosh and Valdes, 1990; Sadoulet and deJanvry, 1995). Precisely, P_T^f is being computed as:

$$P_T^f = \sum_{i=1}^k \alpha_i \left(\frac{WPI_i}{e_i} \right) \quad (3.3)$$

where, e_i is the period average nominal exchange rate expressed in units of a country's own currency per US dollar. Equation (3) was divided by Bangladesh's *CPI* and was then multiplied by the end period exchange rate of Bangladesh taka vis-à-vis the US dollar.¹³

Table 1 reports the 4 different RER series for Bangladesh along with the bilateral nominal exchange rate between the taka and the US dollar. The multilateral RER computed using the weights based on the trade partners' formal trade is given by RERM while RERIND is the series, which assigns a greater weight to India due to the existence of the informal border trade. Figure 1 exhibits the graphical plots of RERM and RERIND along with the nominal exchange rate, *NER*, with respect to the US dollar (i.e., taka per US dollar) – all indexed to 1985=100. It shows that while the two RER series, in general, move in the same direction, the movement is not always uniform. Especially in the 1990s the relationship between RERM and RERIND has changed

¹¹ Initial experiments showed that significantly different RERs would not have been obtained by using the weights for different periods.

¹² Note that the data on *WPI* are not available for all partners. In that case, for some countries, if available, producers' prices were used. When either of the *WPI* and producers' prices was not reported consumers' prices (*CPI*) were used. As correlations between *WPI* and *CPI* are usually high, the choice of price series should not make major differences.

¹³ The exchanges rate between dollar and other partners' currencies are the period average rate as reported in the *IFS*. However, we used the end period exchange rate between Bangladesh taka and US dollar to capture total devaluation of the taka in one year from the preceding year. Our experiments showed that there would not have been any significant difference had we used the period average nominal exchange rate between the taka and the US dollar.

noticeably compared to their co-movement until the 1980s. This can be attributable to the frequent and large downward adjustments of the Indian rupee with respect to the US dollar.¹⁴

The most striking feature of Figure 1 is the two contrasting trends in the RERs: prior to the early-1980s the *RERs* (both RERM and RERIND) are increasing along with the nominal exchange rate (NER) but since the mid-1980s they have remained at the same level although during the same period the nominal rate increased by about 75 per cent. According to the definition given above, a rise in the index will imply depreciation of the RER and thus improved competitiveness. Therefore, from Figure 1 it can be concluded that since the mid-1980s nominal devaluations in Bangladesh have not resulted in any significant improvement in competitiveness. In other words, changes in the exchange rate have largely been offset by the fall in the foreign to domestic price ratio.

Figure 2 provides a simple relationship between the growth of NER (nominal devaluations) and the changes in RERM. The left figure shows that during 1980-2000 only on four instances were the changes in the RERM greater than those of NER.¹⁵ The scatter plot on the right panel depicts a positive relationship between the changes in NER and RERM and the regression equation shows that a one percentage point change in the nominal rate is associated with about 0.68 percentage point rise in the RER index. However, this relationship must be considered with a great caution, as Figure 1 shows the relationship is most likely to have changed significantly between 1980s and 1990s.¹⁶

¹⁴ During 1990-95 India undertook a massive devaluation of its currency by about 95 per cent against the US dollar, while the comparable figure for Bangladesh during the same period was only about 14 per cent.

¹⁵ This happened in 1987, 1990, 1994 and 2000.

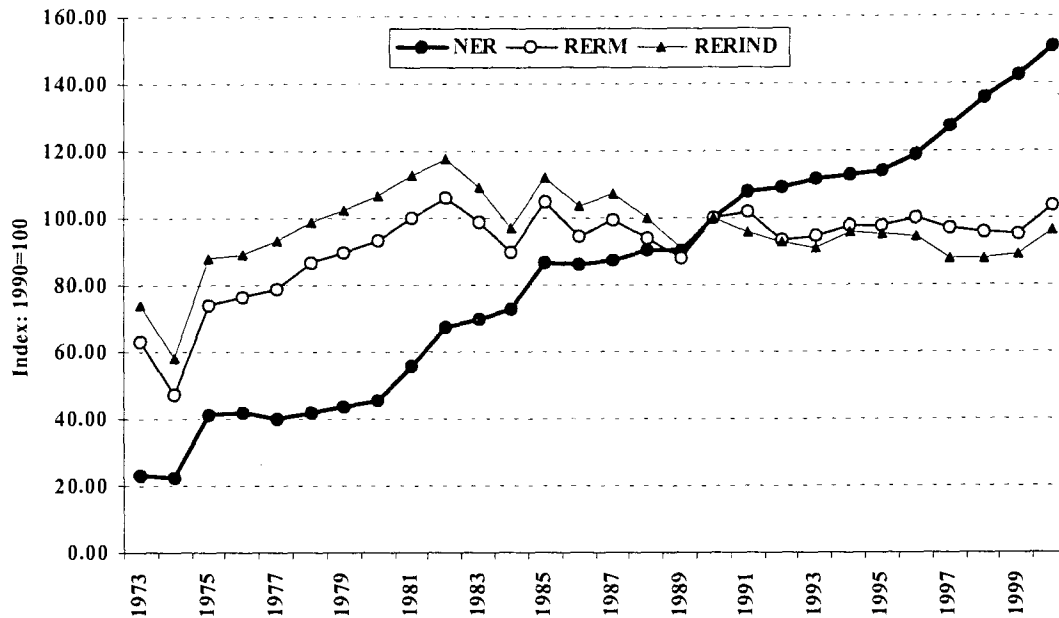
¹⁶ Indeed, for 1980-90 the estimated relationship is: $\Delta \ln RERM = -0.062 + 0.88 \Delta \ln NER$, with the coefficient on the explanatory variable being significant at the one percent level. But for 1991-2000 the estimated relationship turned out to be $\Delta \ln RERM = -0.10 + 0.33 \Delta \ln NER$ and the coefficient on $\Delta \ln NER$ was highly insignificant. Note that these regressions do not suffer from the problems associated with the non-stationarity of the time series data, as the two first differenced variables are stationary.

Table 1: Estimated RERs for Bangladesh

Year	NER	RERM	RERIND	RERB	BRERIN
1973	8.16	63.04	74.16	67.44	102.56
1974	8.07	47.10	57.95	51.38	85.35
1975	14.83	73.79	88.00	84.59	122.06
1976	14.95	76.60	89.19	87.15	118.01
1977	14.40	79.07	93.07	85.03	126.91
1978	14.93	86.67	98.95	90.17	124.73
1979	15.64	89.97	102.61	92.75	131.85
1980	16.25	93.58	106.75	96.89	144.32
1981	19.85	100.25	112.61	111.35	148.56
1982	24.07	106.02	117.44	122.20	154.54
1983	25.00	98.53	109.33	117.58	145.57
1984	26.00	89.40	96.74	113.44	123.41
1985	31.00	104.76	111.87	121.42	142.28
1986	30.80	94.67	103.43	105.61	124.81
1987	31.20	99.45	107.28	100.21	125.80
1988	32.27	94.04	99.89	98.59	111.49
1989	32.27	87.71	90.83	95.07	94.89
1990	35.79	100.00	100.00	100.00	100.00
1991	38.58	101.64	95.80	100.77	79.86
1992	39.00	93.40	92.59	98.19	85.41
1993	39.85	94.70	91.00	101.83	78.26
1994	40.25	97.53	95.69	100.55	84.44
1995	40.75	97.74	95.38	99.72	78.78
1996	42.45	100.06	94.80	102.09	80.66
1997	45.45	96.78	87.73	103.91	78.48
1998	48.50	95.52	87.83	99.78	75.77
1999	51.00	95.31	88.79	99.73	75.75
2000	54.00	103.55	96.48	109.17	77.76

Note: NER is the nominal exchange rate: taka per US dollar. RERM is the multilateral real exchange rate based on 20 most important trade partners' currencies; RERIND is the multilateral RER but with an enhanced weight for India considering the existence of informal border trade; RERB is the bilateral RER with respect to the US dollar only; and finally BRERIN is the bilateral RER with respect to Indian rupees only. Except the NER all are indexed with 1985=100.

Figure 1: Movements in Nominal and Real Exchange Rates



Note: While RERM is the multilateral real exchange rate using the weights based on Bangladesh's bilateral formal trade with 20 most important partners', RERIND is the multilateral RER when a greater weight is assigned to India to take account of the existence of the widespread informal border trade between the two neighbouring countries. NER is the nominal exchange rate: taka per US dollar.

Source: Authors' own computation as explained in the text.

Figure 2: Relationship Between Growth in Nominal Exchange Rate and Multilateral Real Exchange Rate

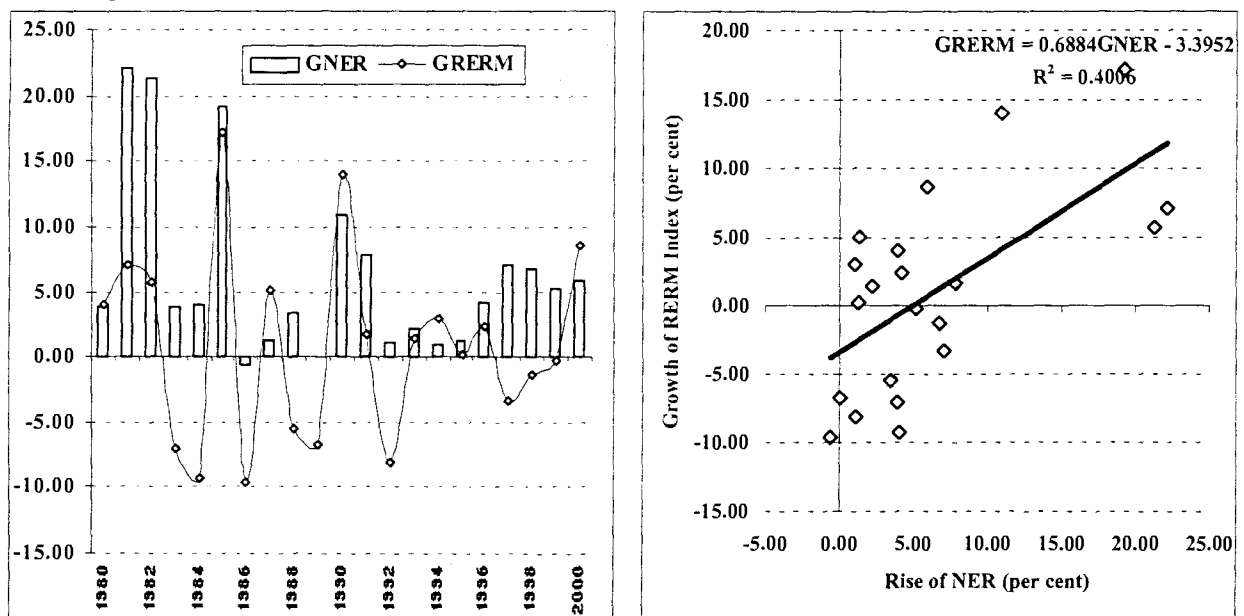
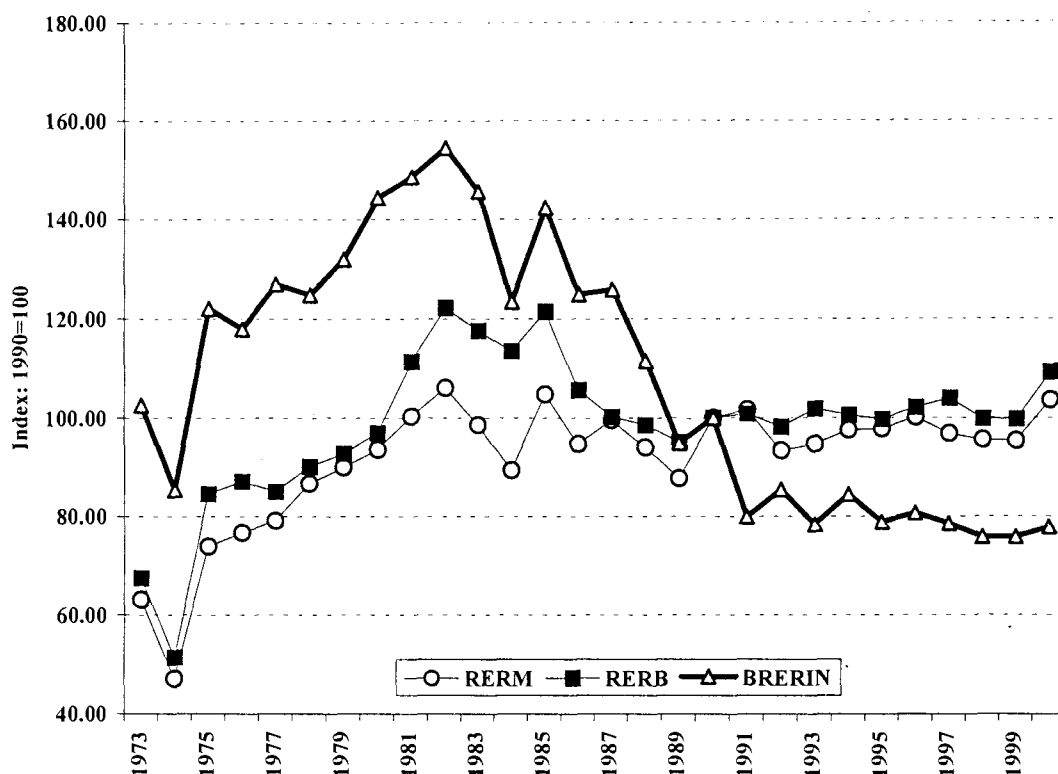


Figure 3: Bilateral vis-à-vis Multilateral RERs



Note: RERM is the multilateral exchange rate as explained in Figure 3.1. RERB is the bilateral exchange rate with respect to the US dollar only while BRERIN is the bilateral real exchange rate with respect to Indian rupees only.
 Source: Authors' own estimates.

Table 1 also provides the two bilateral real exchange rates with respect to the US dollar (RERB) and to the Indian rupee (BRERIN). Both the bilateral series depict significant variation from the RERM; particularly the one with respect to India differs a lot. It is found that in comparison with India Bangladesh's competitiveness underwent a massive decline in the 1980s. In the 1990s Bangladesh has just managed to prevent further deterioration in her competitiveness vis-à-vis India. In contrast, RERB moves rather closely to RERM except for the period of 1982-87 when the US dollar depreciated considerably against other currencies.¹⁷ On the whole, Figure 3 lends support to Edwards' (1989) observation that bilateral and multilateral RER can move differently. In the following, therefore, we will concentrate only on the multilateral RERs.

4. Empirical Specification and Data

4.1. The Empirical Model

The theoretical model, as presented in Section 3, posits a long-run relationship between aggregate output and a vector of other variables comprising RER, a measure of fiscal policy, and an indicator of monetary policy. But mainly following Edwards (1986), most empirical studies (such as Atkins, 2000; Rhodd, 1993; and Upadhyaya and Upadhyay, 1999) also include the terms of trade (TOT) of the country, which is considered to render important influence on the growth of aggregate output. For a small open economy TOT is truly exogenous and if not controlled explicitly in the experiment some of its unaccounted for influence could be transmitted through the indicator of external competitiveness, RER. However, Atkins (2000) and Bahmani-Oskooee (1998) are studies that have ignored the incorporation of TOT in the model. It is true that on the one hand the theoretical model does not call for the inclusion of it, on the other hand many do not treat TOT to be an independent policy variable distinct from the RER (Dornbusch, 1986; Atkins, 2000).¹⁸ For the present paper, while we keep TOT in our main estimating equations, results without it will also be reported.

In the empirical literature there seems to be a consensus with regard to the use of government expenditure as a measure of fiscal policy but to represent the monetary policy two different indicators have been used. While Edwards (1986) used a ‘money surprise’ or unexpected money growth term, Atkins (2000) and Rhodd (1993) have considered total domestic credit instead.¹⁹ For this study the estimation of money surprise function, as specified by Edwards (1986), was not satisfactory and therefore it was decided to use total domestic credit (DC) to represent the

¹⁷ For example, during 1982-87 the yen-dollar exchange rate fell from 259 to 175 resulting in more than 30 per cent appreciation in the Japanese yen.

¹⁸ This is because RER is often considered to be the terms of trade of the country. This is conceivable if it is assumed that the country in question cannot influence the world price of tradables, and consequently the ratio of partners’ tradable goods’ price to her own domestic non-traded goods is the terms of trade.

¹⁹ Edwards defines money surprise as the actual rate of growth of nominal money ($\Delta \log M$) less the expected rate of growth of nominal money ($\Delta \log M^*$) where it is assumed that expectations are formed rationally. This requires estimation of an expected money supply equation, which Edwards specified as

$$\Delta \log M_t = b_0 + b_1 \Delta \log M_{t-1} + b_2 \log M_{t-2} + b_3 DEF$$
, where M is the broad money, DEF is the budget deficit of the government and t is time subscript. Having estimated the money creation equation the estimated values of $\Delta \log M$ are subtracted from the actual money growth to arrive at the surprise money growth series.

monetary policy measure in the empirical model.²⁰ The use of DC can be justified because of its impact on income through domestic investment, and because the control of total bank credit (to government as well as to the private sector) represents one of main instruments of monetary policy in many developing countries including Bangladesh. Using the logarithmic transformation of the variables the empirical specification of the model thus can be written as:²¹

$$\ln Y_t = \beta_0 + \beta_1 \ln(TOT_t) + \beta_2 \ln(GE_t) + \beta_3 \ln(DC_t) + \beta_4 \ln(RER_t) + \nu_t \quad (4.1)$$

where, \ln stands for natural logarithm, time is denoted by subscript t , Y , TOT , GE , and DC , stand respectively for real GDP, terms of trade, real government expenditure, domestic credit in real terms, Real Exchange Rate, and ν is the error term. In the above equation it is expected that β_2 and β_3 are positive while the sign of β_1 cannot be determined a priori. The coefficient β_4 captures the effect of real devaluation on real output growth and is the primary interest of this study.

4.2. Data

We use the revised national income estimates by BBS (2000 and 2001), which is compiled by improving the old national income accounting methodology and widening the coverage. This revision has resulted in an increase in Bangladesh's GDP (in current prices) by 26-43 per cent. Under the new accounting system the BBS provides comparable data for 1980-2000. As a result, our sample will be limited to only 21 annual observations. The TOT index has been estimated from the quoted unit value indices for exports and imports in the BBS (2000 and 2001). The TOT is defined here as the unit value index for exports divided the unit value index for imports. Data on total domestic credit are taken from Bangladesh Bank (2002). Government expenditure comprises government consumption expenditure (recurrent expenditure) as well as public investment expenditure allocated via the annual development plans. These data are taken from Chowdhury (1995) and from the *Bangladesh Economic Survey 2002* published by the Ministry of Finance of the Government of Bangladesh. The data on government expenditure and domestic credit are initially given in current prices but using the implicit GDP deflator (for the revised

²⁰ Rhodd (1993) reports that money surprise functions also does not work out satisfactorily in his empirical investigation. Furthermore, in most regressions of Upadhyaya and Upadhyay (1999) money surprise terms were not significant.

²¹ All empirical studies on the subject use log-linear models. One advantage of logarithmic transformation is that the estimated coefficients can directly be interpreted as elasticities with respect to the relevant variables.

GDP) corresponding figures in real prices have been obtained. Finally, the variable, RER, has already been constructed in Section 3.

5. Estimation Strategy

5.1. Time series properties of the variables

Recent developments in econometrics have emphasised a lot on the characteristics of time series data. Central to this is the distinction made between the stationary and non-stationary time series in contrast to the traditional practice of assuming all variables in the regression model are to be stationary. A time series is said to be stationary if its mean, variance and auto-covariance are independent of time. By now there is compelling evidence that many macroeconomic time series are non-stationary in nature and as a consequence the ordinary least squares (OLS) regressions using these data might produce not only inconsistent and inefficient estimates but also spurious results. In other words, one could obtain a highly significant correlation between variables although in reality there may not exist any such relationship.²² In order to avoid such problems of estimating non-sense relationship the integrating properties of the variables should be examined carefully by testing for the existence of unit roots in variables under consideration.

The two most popular tests for unit roots, which we also intend to use for the present study, are the Dickey Fuller (DF) and Augmented Dickey Fuller (ADF) tests. The DF test is based on equation (5.1) where Y is the variable under consideration, Δ is the first difference operator, subscript t denotes time period, T is the time trend and e is the error term. The null hypothesis for this test is that $(\psi - 1) = 0$ (i.e., Y_t is non-stationary) against the alternative of $(\psi - 1) < 0$ (i.e., Y_t is stationary). The 't' test on the estimated coefficient of Y_{t-1} provides the DF test for the presence of a unit root. The ADF test, on the other hand, is a modification of the DF test and involves augmenting equation (5.1) by lagged values of the dependent variables.²³ This is done to ensure that the error process in the estimating equation is residually uncorrelated.²⁴ More

²² One interesting example of spurious regression is due to Hendry (1980) who found a very strong positive relationship between inflation rate and the accumulated annual rainfall for the United Kingdom.

²³ Note that the DF and ADF tests are usually carried out with and without the time trend term (T) in the regression. If the variable is trended the insertion of the term is required. However, if the variable is not trended DF-ADF regressions can be applied without it.

²⁴ In the case of the annual data incorporation of the first lag of the dependent variable most often overcomes the problem of residual correlation. Higher order of lags would be necessary for quarterly and other high frequency data.

precisely, the ADF version of the test is based on the equation (5.2). As in the case of the DF test, the t-ratio on $(\psi - 1)$ provides the ADF test statistic.

$$\Delta Y_t = \tau + (\psi - 1)Y_{t-1} + \chi T + e_t \quad (5.1)$$

$$\Delta Y_t = \tau + (\psi - 1)Y_{t-1} + \chi T + \delta \Delta Y_{t-1} + e_t \quad (5.2)$$

In both the equations (5.1) and (5.2) the estimated t-ratios on $(\psi - 1)$ are non-standard requiring the computed test statistics to be compared with the corresponding critical values to infer about the stationarity of the variables.²⁵ The DF and ADF tests can, however, provide contrasting evidence and there appears to be a consensus in the literature that the ADF test is preferable to the DF test. It is quite common to find that macroeconomic time series data are non-stationary on their levels but stationary on their first or higher order differences. Following Engle and Granger (1987) a time series is said to be integrated of order d [usually denoted as $\sim I(d)$] with d is the number of times the series needs to be differenced in order to become stationary.

It needs to be mentioned that in small sample the testing procedure for unit roots might be very complicated. Not only that the results emanating from different unit root test regressions can be inconclusive but also the critical values for such tests may prove to be very demanding. Apart from these, it is well known that the low power of the DF and ADF tests is an unavoidable fact as Harris (1995) points out that the most important problem faced when applying the unit root test is their probable poor size and power properties.²⁶ This is often reflected in the tendency to over-reject the null when it is true and under-reject the null when it is false. In a small sample the problem is likely to be even worse. Thus in the case of small sample Hall (1986) suggests the inspection of the autocorrelation function and correlogram as an important tool in determining whether the variables are stationary or not. The autocorrelation function for any variable at any lag k is defined by the ratio of covariance at lag k divided by variance.²⁷ When the estimated autocorrelation coefficients at different lags are plotted against k , population correlogram is

²⁵ These critical values were first computed by Dickey and Fuller (1981). If the computed test statistics exceed the critical values, the null hypotheses underlying the DF-ADF tests are rejected. Computed t-ratios and the corresponding critical values are compared on their absolute levels.

²⁶ Engle and Granger (1987) also highlighted the low power of the DF and ADF tests.

²⁷ The autocorrelation coefficient like any ordinary correlation coefficient lies between -1 and $+1$.

obtained.²⁸ For non-stationary variables correlograms die down slowly giving rise to a secular declining trend in the graph of autocorrelation coefficients while in the case of stationary variables they damp down almost instantly and then show random movement.

For the present paper, therefore, we will employ the DF-ADF tests, autocorrelation coefficients and correlograms to determine the integrating properties of the variables.

5.2. Cointegration and Error Correction Modelling

5.2.1. The Engle-Granger Procedure

Once it is determined that the variables in the model are non-stationary, the only way to infer about the long-run relationship is to employ some kind of cointegration technique. There are several cointegration methodologies in the literature – the simplest one being the Engle-Granger two step procedure. The basic idea behind the Engle-Granger technique is that if two variables say Z_t and X_t are both $\sim I(d)$, a linear combination of these two variables such that $V_t = X_t - \theta Z_t$, in general, will also be $\sim I(d)$. Engle and Granger, however, showed that in an exceptional case if the constant θ yields an outcome where $V_t \sim I(d-a)$ and $a > 0$, then X_t and Z_t will be cointegrated. Usually the linear combination represented by the residuals from the OLS regression is tested for stationarity. Thus, if Z_t and X_t are both $\sim I(1)$, they will be cointegrated and have a valid long-run relationship if residuals from the OLS regression of X_t on Z_t is $\sim I(0)$. This is what is known as the first step of Engle-Granger procedure.

One important contribution of Engle and Granger (1987) was to find that if variables were cointegrated, there would have existed an error-correction model (ECM) of that cointegrating relationship. The ECM will then capture the short-run dynamics of the long-run behaviour, which is known as the second step of Engle-Granger procedure. The ECM is constructed by regressing the dependent variable in stationary form, onto its own lagged values and the current and lagged values of the stationary forms of the dependent variables, and the lagged error term

²⁸ Note that in practice we only have a realisation of a stochastic process and therefore can only compute sample autocorrelation function, which is defined as:

$$\frac{\sum (Y_t - \bar{Y})(Y_{t-k} - \bar{Y})}{\sum (Y_t - \bar{Y})^2}$$

from the cointegrating relationship. If we assume that both Z_t and X_t are $\sim I(1)$ such that ΔZ_t and ΔX_t are $\sim I(0)$, the ECM can be represented as:

$$\Delta Z_t = \pi_0 + \sum_{i=0}^m \pi_{1i} \Delta X_t + \sum_{i=1}^n \pi_{2i} \Delta Z_t + \pi_3 \hat{v}_{t-1} + \varepsilon_t \quad (5.3)$$

Equation (5.3) gives a very general representation of the ECM. Since all variables in (5.3) are $\sim I(0)$, the problem of spurious regression is overcome. It is worth noting that the ECM is not a mere regression of the stationary variable rather it includes \hat{v}_{t-1} , the deviation from the steady-state long-run path, which basically contains the long-run information. Thus the ECM captures the short-run relationship taking into consideration of the long-run information. A valid representation of the ECM will require $0 > \pi_3 \geq -1$. The usual practice with the error correction modeling is to follow the “general to specific” methodology by constructing a general model in the beginning and subsequently reduce it to a parsimonious form after dropping all the insignificant variables step-by-step.

5.2.2. The Phillips-Hansen Fully Modified OLS

In estimating the equations as specified in Section 4 we can employ the Engle-Granger cointegration procedure to test for a valid long-run relationship. However, although this procedure can test for cointegration, it yields standard errors that do not provide the basis for valid inferences. In equations with more than two explanatory variables this can be problematic in the sense that even if the variables are found to be cointegrated we cannot be certain whether any particular explanatory variable is significant or not.²⁹ We propose to handle this problem by using the Phillips-Hansen Fully Modified OLS (PHFMOLS) technique (Phillips and Hansen, 1990). The Phillips-Hansen method is an optimal single-equation technique, which is asymptotically equivalent to maximum likelihood procedure. It makes a semi-parametric correction to the OLS estimator to eliminate dependency of the nuisance parameters and provides standard errors that follow standard normal distribution asymptotically and thus are valid for drawing inferences. Due to this particular advantage the use of PHFMOLS has become

²⁹ That is, for example, in a three variable, say Y, X and Z, regression model cointegration does not necessarily suggest statistically significant influence of all both the explanatory variables, X and Z. It might be that only X is significant but not Z and vice-versa. Since the computed standard errors in the first step of the Engle-Granger procedure is not valid, correct statistical inference from the estimated model is not possible.

quite popular in international trade and macroeconomic modeling.³⁰ The PHFMOLS procedure can be described by the following.³¹

Consider the data generating mechanism for Y_t following the cointegration system:

$$Y_{1t} = \alpha_0 + \alpha_1 + \beta'Y_{2t} + u_{1t} = \lambda R_t + u_{1t} \quad (5.4)$$

$$\Delta Y_{2t} = u_{2t} \quad (5.5)$$

$$u_t = \begin{pmatrix} u_1 \\ u_2 \end{pmatrix} = \psi(L)\varepsilon_t, \quad E(\varepsilon_t \varepsilon_t') = PP' \quad (5.6)$$

where, Y_{1t} and Y_{2t} are scalar and mxt vector of $I(1)$ stochastic processes, $\lambda' = (\alpha_0 + \alpha_1 + \beta')$ and $R_t = (Y_{1t}, Y_{2t})$. We define:

$$\Omega = \psi(1)P, \quad \Sigma = \Omega\Omega' = \begin{pmatrix} \Sigma_{11} & \Sigma_{12} \\ \Sigma_{21} & \Sigma_{22} \end{pmatrix} \quad (5.7)$$

Σ is the long-run covariance matrix of u_t . As mentioned earlier, the PHFMOLS estimator is an optimal single equation method based on the use of OLS on equation (5.4) with semi-parametric corrections for serial correlation and potential endogeneity of the right hand side variables.

Consider the OLS estimator of the cointegrating equation (5.4) by $\hat{\lambda} = (R_t'R_t)^{-1} R_t'Y_{1t}$, where R_t and Y_{1t} are respectively $T \times (m+2)$ and $T \times 1$ matrices of observations on R_t and Y_{1t} . Due to serial correlation in u_{1t} and endogeneity of Y_{1t} , $\hat{\lambda}$, in general, is consistent but biased. The FM procedure modifies the OLS estimator $\hat{\lambda}$ to correct for serial correlation and endogeneity bias. The FM estimator is given by:

$$\hat{\lambda}^* = \begin{pmatrix} \hat{\alpha}_0^* \\ \hat{\alpha}_1^* \\ \hat{\beta}^* \end{pmatrix} = \begin{pmatrix} T & \sum_{t=1}^T Y_{2t}' & \sum_{t=1}^T t \\ \sum_{t=1}^T t & \sum_{t=1}^T t Y_{2t}' & \sum_{t=1}^T t^2 \\ \sum_{t=1}^T Y_{2t} & \sum_{t=1}^T Y_{2t} Y_{2t}' & \sum_{t=1}^T Y_{2t} t \end{pmatrix}^{-1} \begin{pmatrix} \sum_{t=1}^T \hat{Y}_{1t}^+ \\ \sum_{t=1}^T t Y_{1t}^+ \\ \sum_{t=1}^T Y_{2t} Y_{1t}^+ - T \hat{V}_T^+ \end{pmatrix} \quad (5.8)$$

³⁰ Amongst others, Athukorala and Riedel (1995) and (1996), Muscatelli, et al. (1992), Senhadji (1998) and Senhadji and Montenegro (1998) have used the Phillips-Hansen procedure to modeling trade for various countries, while Mallick (1999) has applied the procedure to macroeconomic modelling for India.

$$\hat{Y}_{1t}^+ = Y_{1t} - \hat{\Sigma}_{21}' \hat{\Sigma}_{22}^{-1} \Delta Y_{2t} \quad (5.9)$$

$$\hat{\Sigma} = \begin{pmatrix} \hat{\Sigma}_{11} & \hat{\Sigma}_{21}' \\ \hat{\Sigma}_{21} & \hat{\Sigma}_{22} \end{pmatrix} = \hat{\Gamma}_0 + \sum_{\nu}^q \left(1 - \frac{\nu}{q+1}\right) (\hat{\Gamma}^{\nu} + (\hat{\Gamma}^{\nu})')$$
(5.10)

$$\hat{\Gamma}_{\nu} = T^{-1} \sum_{t=\nu+1}^T \begin{pmatrix} \hat{u}_{1t} \hat{u}_{1t-\nu} & \hat{u}_{1t} \hat{u}_{2t-\nu}' \\ \hat{u}_{2t} \hat{u}_{1t-\nu}' & \hat{u}_{2t} \hat{u}_{2t-\nu}' \end{pmatrix} = \begin{pmatrix} \hat{\Gamma}_{11}^{\nu} & \hat{\Gamma}_{12}^{\nu} \\ \hat{\Gamma}_{21}^{\nu} & \hat{\Gamma}_{22}^{\nu} \end{pmatrix} \quad (5.11)$$

$$\hat{V}_T^+ = \sum \left(1 - \frac{\nu}{q+1}\right) [(\hat{\Gamma}_{12}^{\nu})' + (\hat{\Gamma}_{22}^{\nu})'] \quad (5.12)$$

Where q is the bandwidth parameter in the Bartlett window used in the estimation of the long-run covariance matrix. The difference between the OLS and FM estimators is highlighted in the last vector of (5.8) where Y_{1t} is replaced by \hat{Y}_{1t}^+ (which corrects for the potential endogeneity of Y_{2t}) and the factor TV_T^+ (which corrects for the potential autocorrelation of the error term). The FM estimator $\hat{\lambda}^*$ has the same asymptotic behaviour as the full information system maximum likelihood estimators.

A problem with the Engle-Granger and PHFMOLS is that they ignore the possibility of multiple cointegrating vectors. This problem can be tackled by Johansen's (1988) Full Information Maximum Likelihood (FIML) procedure. However, there are two important problems associated with this approach. First of all, the results from the Johansen procedure can be very sensitive to the choice of lag-length (Hall, 1991; Banerjee *et al.* 1993). Although there are statistical tests for choosing the appropriate lag-lengths, in a small sample such tests may not be feasible. Moreover, severe problem of collinearity among the regressors may also arise when a considerable size of VAR is used (Athukorala and Riedel, 1996). As in the present study we will be dealing with a small sample size (annual observations for 20 years), the Johansen procedure may not be an appropriate one. As a result, we decided to rely on the single equation procedures, such as those of Engle-Granger and PHFMOLS.

³¹ This is based on the illustrations in Senhadji (1998).

5.2.3. Existence of a Long-run Relationship

5.2.3.1. Testing for Cointegration

From the above our estimation strategy can be summarized as follows. First, the time series properties of the variables will be analysed and in the case of equations containing non-stationary variables PHFMOLS method will be used, which would provide standard errors for valid inferences. The estimation by PHFMOLS itself does not guarantee cointegration needing one to check for residual stationarity. In the literature the standard practice of testing for cointegration has been the use of ADF test, which is given in equation (5.13). Note that in contrast to the regular ADF regressions, the test for residual does not include any intercept term.³²

$$\Delta \hat{v}_t = \rho \hat{v}_{t-1} + \kappa \Delta \hat{v}_{t-1} + \tau \quad (5.13)$$

The null hypothesis for the test is that $\rho=1$ (non-cointegration) against the alternative of $\rho<1$ (cointegration). Like the regular ADF test statistics the estimated standard errors in (5.13) are non-standard and hence they will have to be compared with the appropriate critical values as estimated by Engle and Granger (1987) and Mackinnon (1991).³³

Despite its widespread use, the low power of the ADF test is considered to be a serious shortcoming for cointegration test. Engle and Granger showed that when $\rho=0.9$ the ADF test for cointegration has about 28 per cent chance of not rejecting the null of no cointegration even when it is false.³⁴ In small sample, testing for cointegration is more troublesome as apart from the low power critical values for such tests become more demanding. One effective way of tackling this problem will be to follow Hall (1986) and examine the autocorrelation coefficients and the resultant correlograms of the estimated error term from the static long-run equations.

5.2.3.2 Testing Cointegration for Variables With Different Order of Integration: The Pesaran et al. Test

³² This is because the by definition the mean of the residual should be zero.

³³ These days many econometric software routinely compute such critical values.

³⁴ Razzaque and Ahmed (2000) illustrates a case when the ADF test on the residuals falls into the trap of its low power.

It is very well possible to have an equation with variables of different integrated orders – for example, a mixture of $I(1)$ and $I(0)$ variables. Then the question is whether the $\sim I(0)$ regressors play a role in determining the $I(1)$ variable. In a study Holden and Perman (1994) have considered a model with two $I(1)$ and one $I(0)$ variables. The authors used the Johansen rank cointegration procedure to determine a valid long-run relationship between the two $I(1)$ variables and then included the $I(0)$ variable only in the short-run error-correction model. The procedure, thus, assumes that the $I(0)$ variable does not have any role to play in the long-run disregarding the economic theory behind it. On the other hand, Pesaran *et al.* (2001) observe that “the strict precondition for the same order of integration of the variables in a model involves a certain degree of pre-testing, thus introducing a further degree of uncertainty into the analysis of a long run relationship”. They have strongly argued that the fact that the variables in the estimating equation have different orders of integration does not necessarily mean that they are unlikely to have any long-run impact.³⁵ Pesaran, *et al.* (2000) have also devised a strategy, which tests the existence of a long-run relationship when the variables are a mixture of $I(0)$ and $I(1)$. This procedure is based on an OLS estimation of unrestricted error correction model, a general specification of which can be written as:

$$\Delta \ln Z_t = \alpha + \gamma \ln X_{t-1} + \xi Z_{t-1} + \sum_{i=1}^p \pi_i \Delta \ln X_{t-i} + \sum_{i=0}^g \delta_i \Delta Z_{t-i} + \varepsilon_i \quad (5.14)$$

Estimation of (11) in itself is not interesting since the existence of a long-run relationship can only be tested by examining the joint null hypothesis that $\gamma = \xi = 0$ with the help of either a *Wald* or an *F* test. The presence of a long run relationship requires the rejection of this null. However, as the asymptotic distribution of these statistics is non-standard, Pesaran *et al.* provide the necessary critical upper (F_U) and lower (F_L) bound for the *F* test.³⁶ The F_U are derived under the assumption that all variables are $I(1)$ and the F_L considers all of them to be $I(0)$. If the computed *F* statistic (F), which is obtained by restricting that $\gamma = \xi = 0$, is greater than the critical upper value, i.e. $F > F_U$, we reject the null and conclude that there is a valid long-run relationship among the variables. If $F < F_L$ then no long-run relationship exists, and finally, if $F_L < F < F_U$ the test is inconclusive. Pesaran *et al.* clearly point out that “[I]f the computed *Wald* or

³⁵ See, Pesaran *et al.* (1999), particularly page 19, first paragraph.

³⁶ Pesaran *et al.* give both the critical values for *Wald* and *F* Statistics. In this paper we will only consider the *F*-statistics.

F-statistic falls outside the critical value bounds a conclusive inference can be drawn without needing to know the integration/cointegration status of the underlying regressors.”³⁷

³⁷ Pesaran *et al.* (2001), p.1.

6. Estimation Results

6.1. Examining the Time Series Properties of Variables

As a first step toward the estimation of equation (4.1), we test all variables to determine whether they can be represented as a stationary or non-stationary process by employing the unit root tests and examining the correlograms and autocorrelation functions. Table 2 provides the results of DF and ADF tests on level and first difference of the variables both with and without the trend term in the regressions, while Figure 3 presents the graphical plot of variables along with their correlograms.

Table 2: Unit Root Test of the Variables

Variables	Without Trend		With Trend	
	DF	ADF	DF	ADF
Ln Y	2.76	2.82	-0.67	-0.35
Δ ln Y	-3.87	-2.46	-4.82	-3.63
Ln TOT	-1.05	-0.94	-2.79	-1.61
Δ ln TOT	-7.00	-2.63	-6.81	-2.53
Ln GE	-1.00	-0.76	-2.94	-3.14
Δ ln GE	-3.49	-4.12	-3.67	-4.79
Ln DC	-0.15	-0.039	-1.82	-2.17
Δ ln DC	-3.66	-3.77	-3.56	-3.69
Ln $RERM$	-4.69	-4.64	-4.50	-4.51
Δ ln $RERM$	-6.40	-6.13	-6.41	-6.05
Ln $RERIND$	-2.25	-1.68	-3.75	-3.48
Δ ln $RERIND$	-6.07	-4.78	-6.19	-4.79

Note: The 95 per cent critical value for DF and ADF test statistics without the trend term is -3.02 . The comparable statistic for DF-ADF regressions with the trend term is -3.67 .

In the case of aggregate output it is found from Table 2 that while all DF and ADF tests on $\ln Y$ cannot reject the null hypothesis of non-stationarity, such tests on $\Delta \ln Y$ are inconclusive as the DF tests reject the null hypothesis but ADF test cannot. In light of the fact that the ADF test is usually considered to be preferable to DF, without further testing we cannot be certain about the order of integration of $\ln Y$. The correlograms of $\ln Y$ do not show any tendency of damping down while those on $\Delta \ln Y$ tail off on the first lag just like any stationary variable. The examination of the sample autocorrelation coefficients (ACFs) revealed that such coefficients of $\ln Y$ were individually statistically significant upto the second order of lag but for $\Delta \ln Y$ none of

them was significant.³⁸ If a series is stationary, sample autocorrelation coefficients are approximately normally distributed with zero mean and standard error $1/\sqrt{n}$, where n is the sample size. Since we have a sample of 20 observations, it will imply a standard error of 0.2236. Now, following the properties of the standard normal distribution, the 95 per cent confidence interval for any of the sample autocorrelation coefficient will be $\pm 1.96(0.2236) = \pm 0.4382$. If all autocorrelation coefficients fall inside this interval one cannot reject the null hypothesis that the true autocorrelation coefficient is zero. It was found that up to the 4th lag length the null hypothesis could be rejected for $\ln Y$ but not for $\Delta \ln Y$. Besides, one can test the joint hypothesis that all the autocorrelation coefficients are simultaneously equal to zero by employing the Box-Pierce and Ljung-Box statistics.³⁹ Both these statistics provided the evidence that the joint hypothesis of all the autocorrelation coefficients are simultaneously equal to zero can be rejected on the level of the variable $\ln Y$ but not on its first difference $\Delta \ln Y$.⁴⁰ Thus, probably it would not be inappropriate to conclude that $\ln Y$ is $\sim I(0)$.⁴¹

The unit root tests on terms of trade variable (TOT) are also inconclusive. The tests cannot reject non-stationarity of $\ln TOT$ but the test statistics on $\Delta \ln TOT$ offer contrasting evidence. The graphical plot of $\Delta \ln TOT$, as shown in Figure 4, seems to be stationary, correlograms show random movement but the sample autocorrelation coefficient at the first lag order is found to just exceed the 95 per cent error bar. The second difference of the variable also could not reject the non-stationarity. As mentioned earlier that in small sample the unit root test might prove to be problematic and given the graphical plots we will consider $\ln TOT$ to be $\sim I(1)$.

³⁸ Since we have a sample of 20 observations, the examination of autocorrelations coefficients upto the 5th lag order should be sufficient.

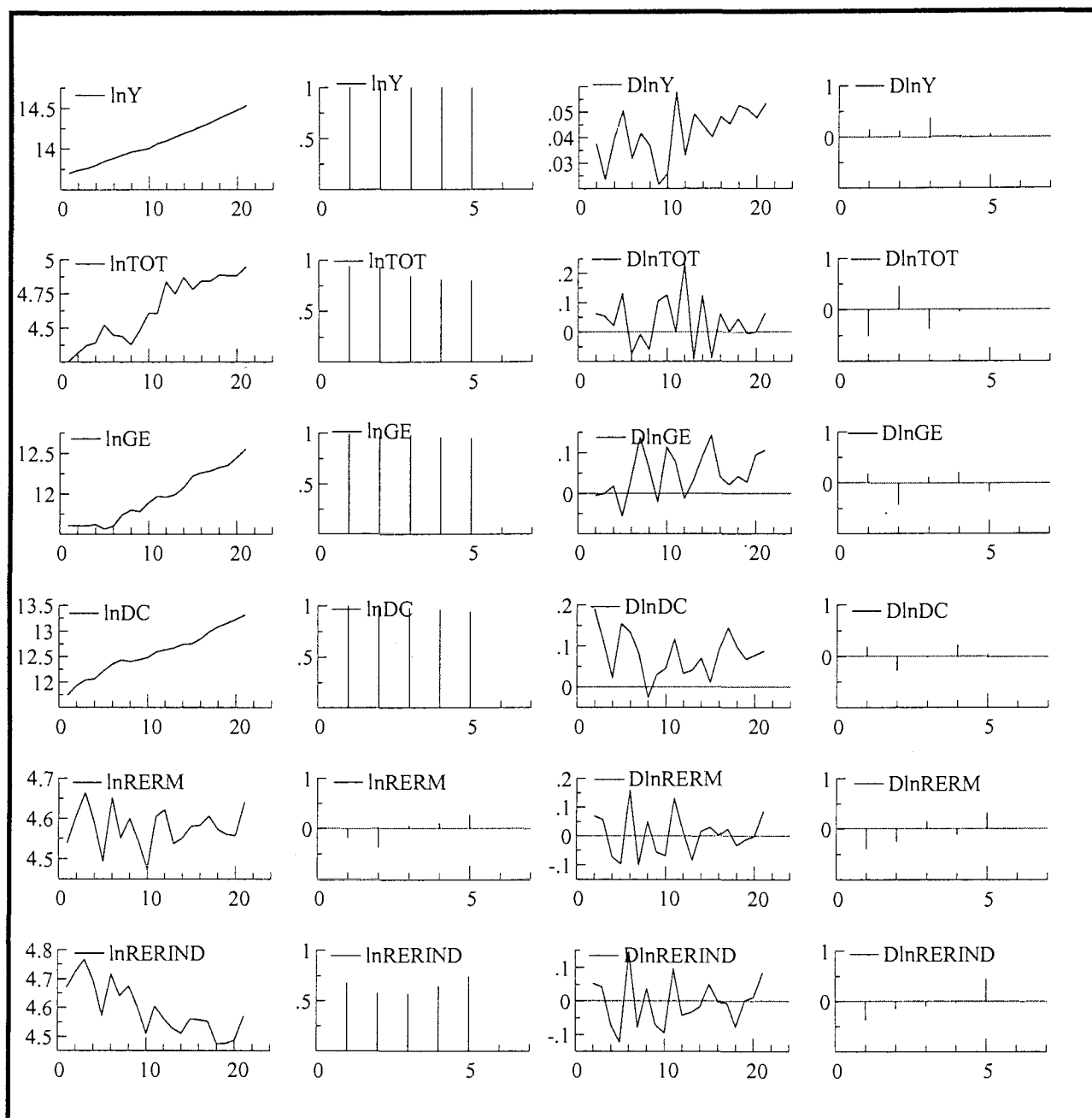
³⁹ The Box-Pierce statistic (also known as the Q statistic) is given by $Q = n \sum_{k=1}^m \hat{p}_k^2$, where n is the sample size, m is the lag length and \hat{p} is the sample autocorrelation coefficient. On the other hand, the Ljung-Box statistic (LB) is

derived as: $LB = n(n+2) \sum_{k=1}^m \left(\frac{\hat{p}_k^2}{n-k} \right)$.

⁴⁰ Such a conclusion remains valid irrespective of the order of lag chosen.

⁴¹ Note that $\Delta \ln Y$ is the growth of real GDP and it does not make sense to consider that the growth rate of GDP can be non-stationary over a long period of time.

Figure 4: Plot of Variables on Their Levels and First Differences and Correlograms



Note: The above graphs are produced by using PcFiml, version 9.10 (Doornik and Hendry, 1997).

Turning to government expenditure and domestic credit we have rather strong evidence that these two variables are non-stationary on their levels but stationary on their first differences. In the case of $\ln GE$, none of the DF and ADF tests can reject the unit root hypothesis, while similar tests on $\Delta \ln GE$ decisively reject non-stationarity. Similarly, on the level of $\ln DC$ the computed

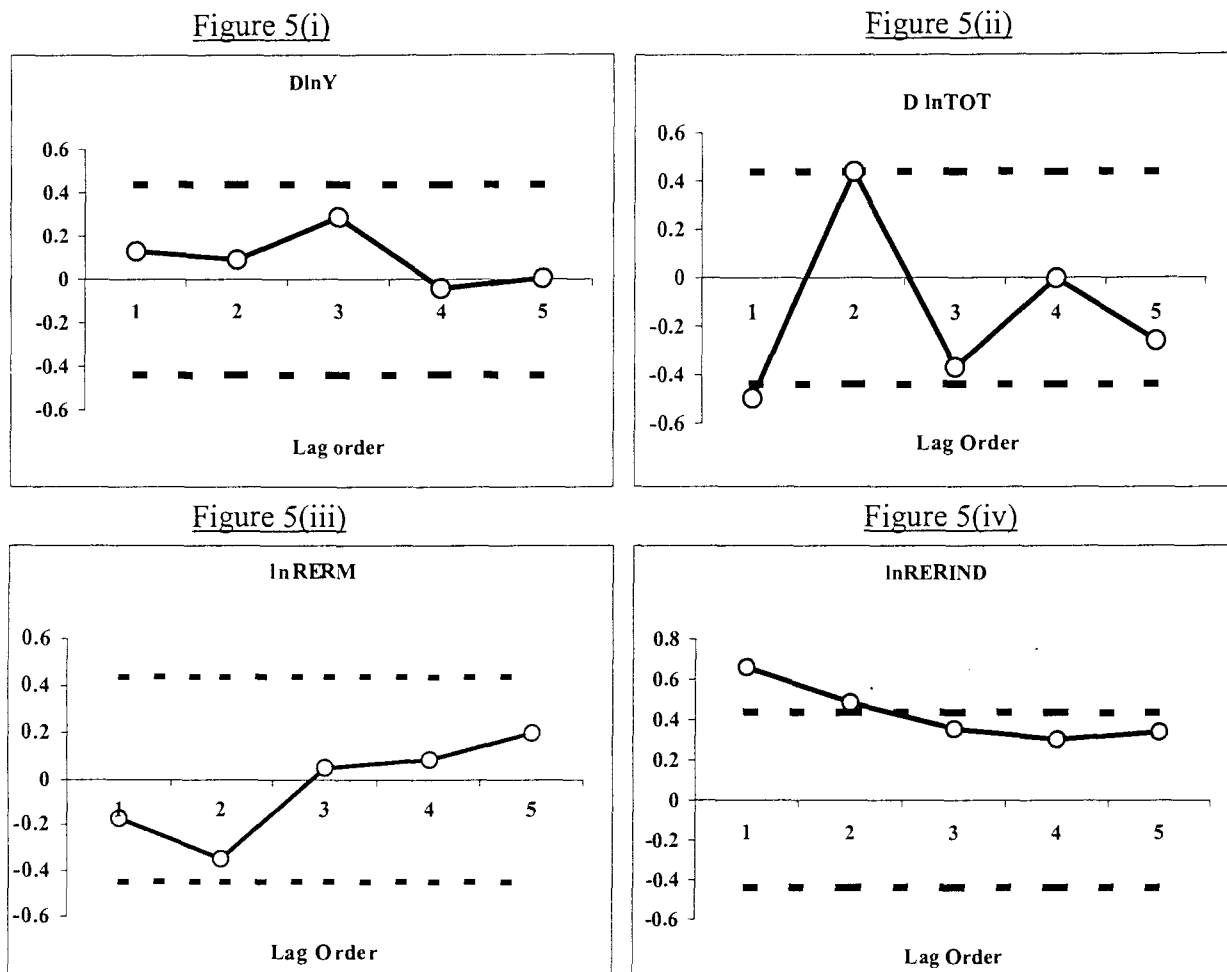
DF-ADF statistics fall below the corresponding 95 per cent critical values but ADF statistics on $\Delta \ln DC$ exceed such critical values. The correlograms also clearly provide support to the unit root test results.

Finally, we consider the unit root properties of the two RER series, RERM and RERIND. All the DF and ADF test statistics very strongly reject the null hypothesis of non-stationarity on both the level and first difference of RERM suggesting that $\ln RERM$ is $\sim I(0)$. The graphical plots of the $\ln RERM$ and $\Delta \ln RERM$ along with the corresponding correlograms also validate the unit root test results. Figure 5(iii) plots the sample autocorrelation coefficients of $\ln RERM$ and the error-bars, which also confirms the stationarity of the variable.

In contrast, the other multilateral RER, RERIND, appears to be an $\sim I(1)$ variable. Apart from the DF test with the trend term, all tests in Table 2 cannot reject the unit root for $\ln RERIND$, while the hypothesis of non-stationarity $\Delta \ln RERIND$ is decisively rejected by every test. Examination of correlograms and autocorrelation coefficients also support the non-stationarity of $\ln RERIND$ and stationarity of $\Delta \ln RERIND$.

That the integrating properties of $\ln RERM$ and $\ln RERIND$ are different is obvious from Figure 4. First differences of the two variables and their corresponding correlograms behave similarly and according to a stationary variable but the graphical plot of the two level variables and their correlograms are completely opposite: correlograms for $\ln RERIND$ do not damp down while those of $\ln RERM$ represent small autocorrelation coefficients and depict random movements. Figure 5(iii) and 5(iv) exhibit the sample ACFs and error bars for both $\ln RERM$ and $\ln RERIND$ where all autocorrelation coefficients for $\ln RERM$ fall within the 95 per cent confidence interval but in the case of $\ln RERIND$ such coefficients at the first two lags fall outside of the error bands. The most important implication of this contrast is that $\ln RERM$ and $\ln RERIND$ do not move uniformly overtime and it will not be possible to find a long-run relationship between these two variables. In other words, RER obtained by considering the existence of informal border trade results in a completely different series than the one obtained by considering the formal trade only.

Figure 5: Sample ACFs and Error-bars for $\Delta \ln Y$, $\Delta \ln TOT$, $\ln RERM$ and $\ln RERIND$



6.2. Estimating the Long-run Relationship

The estimating equation in (4.1) postulates a static long-run relationship among the variables. However, a valid long-run relationship can only be found if the variables in the model cointegrate. If $\ln RERM$ is used as a measure of RER then equation (4.1) will have a mixture of $\sim I(0)$ and $\sim I(1)$ variables requiring us to use Pesaran *et al.* (2001) test for cointegration.⁴² Therefore, in order to determine the long-run relationship (5.14) was run with $p=1$ and $g=0$.⁴³ The F statistic after deleting the lagged dependent variables was computed at 4.47, which needs

⁴² This is because in section 6.1 it is found that $\ln Y$, $\ln GE$, $\ln DC$, and $\ln TOT$ are $I(1)$ while $\ln RERM$ is $I(0)$.

to be compared with Pesaran *et al.*'s critical values. The critical values for a model with four regressors at the 95 percent level are $F_L = 2.86$ and $F_U = 4.01$.⁴⁴ Since the computed F statistic exceeds F_U , the null hypothesis of no long-run relationship between the variables in the model can be rejected. With this finding of cointegration we now proceed to estimate the long-run equation using the PHFMOLS.

Table 3: PHFMOLS Estimates of Long-run Relationship

LnY =	6.147 ^{***}	+ 0.081 ^{***}	lnTOT + 0.331 ^{***}	lnGE + 0.32 ^{***}	lnDC - 0.093 ^{**}	lnRERM
(s.e.)	(0.222)	(0.026)	(0.028)	(0.022)	(0.045)	
t-ratio	28.36	3.12	11.72	14.50	-2.04	
Adjusted R ²	= 0.996					
LnY =	6.227 ^{***}	+ 0.065 ^{**}	lnTOT + 0.327 ^{**}	lnGE + 0.317 ^{**}	lnDC - 0.077	lnRERIND
(s.e.)	(0.316)	(0.0323)	(0.029)	(0.023)	(0.05)	
t-ratio	19.69	2.01	10.95	13.32	-1.53	
Adjusted R ²	= 0.996					

Note: Statistical significance at one, five, and 10 per cent level are given respectively by ^{***}, ^{**}, and ^{*}.

The top row in Table 3 gives the Phillips-Hansen estimate of the long-run relationship when the real exchange rate is represented by lnRERM. It is found that the terms of trade, government expenditure and domestic credit turn out to be correctly signed. The terms of trade elasticity is estimated at 0.08. While a 10 per cent rise in government expenditure is found to be associated with a 3.3 per cent increase in aggregate output, almost similar magnitude of elasticity is also estimated for domestic credit. TOT, GE and DC all come out to be statistically significant at the one per cent level.

The parameter of our interest in Table 3, *i.e.*, the coefficient on lnRERM, is negatively signed and is statistically significant at the five per cent level. The negative sign on the coefficient indicates that as the RER index rises, (or nominal currency adjustments causing in real devaluations), aggregate output falls. The estimated coefficient on lnRERM suggests that a one per cent rise in RER index will result in a 0.09 per cent decline in the GDP.

⁴³ Since we have a small sample, over parameterisation of the model can be very problematic in terms of having fewer degrees of freedom. Such choice of lag order can be rationalised by the fact that we are using annual data in this exercise.

⁴⁴ The critical values correspond to unrestricted intercept and no trend in the regression equation.

Although Pesaran *et al.* (2001) test was carried out to infer about the cointegrating relationship of the estimated equation, autocorrelation coefficients of the residuals from the long-run relationship were examined to verify the cointegration results. Figure 6 gives the graphical plot of the cointegrating relationship and its autocorrelation coefficients upto the 5th lag order along with the 95 per cent level error bar. The graph seems to portray a stationary nature of the long-run relationship and all the autocorrelation coefficients are also found to lie within the confidence interval.

The bottom row in Table 3 estimates the long-run relationship using lnRERIND as a measure of RER. Recall that in contrast to lnRERM, lnRERIND is an $\sim I(1)$ variable. Therefore, usual Engle-Granger procedure can be undertaken to infer about the long-run relationship. It is observed that the use of lnRERIND does not change the sign and significance of TOT, GE and DC. Moreover, the size of the estimated parameters is very much comparable to the previous set of estimates. The coefficient on lnRERIND is again negatively signed and its magnitude is comparable to lnRERM but it fails to be significant at the conventional level. In fact, the coefficient becomes significant only at the 20 per cent level. The ADF test statistic of the estimated residuals from this equation was computed at -4.12 against its critical value of -5.19 thus failing to reject the null hypothesis of non-cointegration. However, given the low power of the ADF test and the demanding critical values in small samples, examination of the residual plot and autocorrelation coefficients should be considered before discarding the long-run validity of the model. Figure 7, however, shows that the residuals behave like a stationary variable and particularly based on the evidence provided by the sample autocorrelation coefficients, the cointegrating relationship represented by the estimated equation in the bottom row of Table 5 should not be rejected.

Figure 6: Estimated Cointegrating Relationship with $\ln RERM$ and the Sample ACF

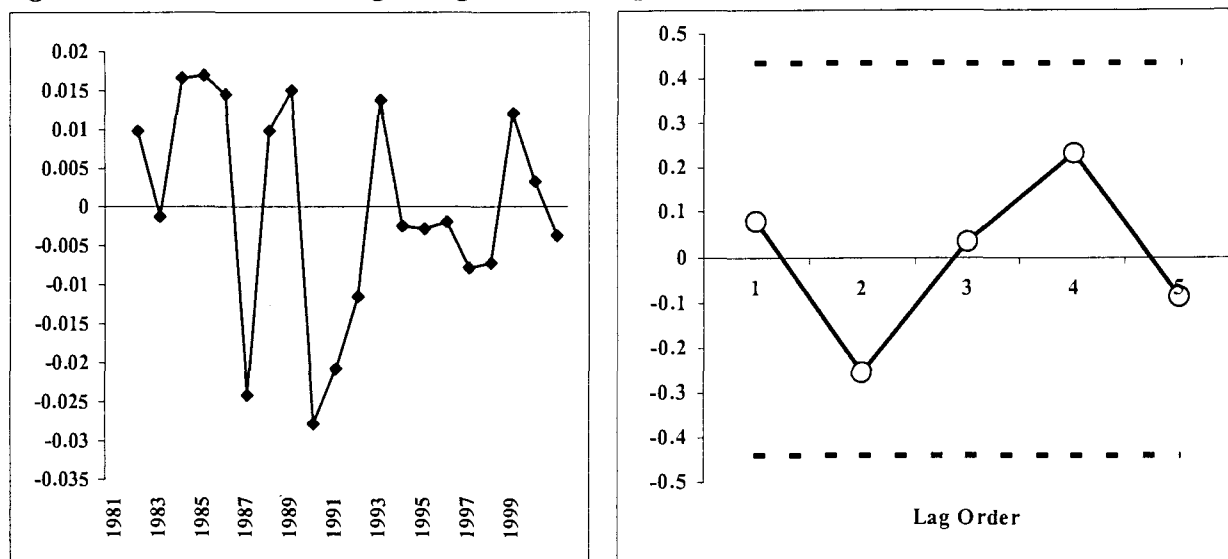
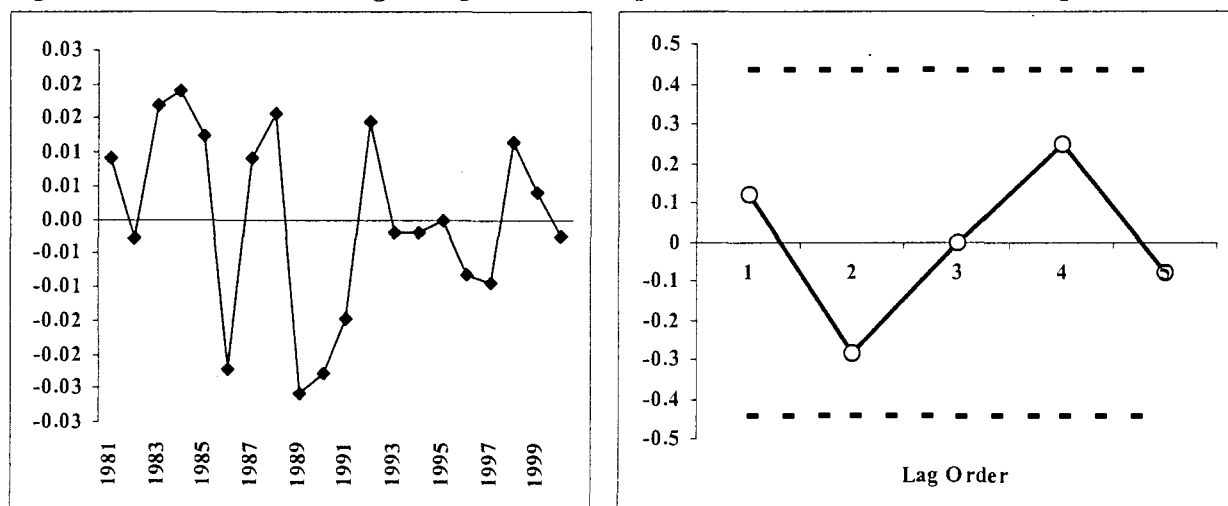


Figure 7: Estimated Cointegrating Relationship with $\ln RERIND$ and the Sample ACF



In Section 4 it was mentioned that there is some debate over the use of TOT in the regression model. Therefore, it might be of interest to know how the results would change if $\ln TOT$ was dropped from the equations in Table 3. Table 4 now shows that the exclusion of $\ln TOT$ does not change the size and significance of $\ln GE$ and $\ln DC$ much. While the coefficient on $\ln RERM$ is only marginally increased (absolutely), its negative sign and the level of statistical significance remain unchanged. The biggest change, however, is associated with $\ln RERIND$. The estimated parameter is now almost twice as big as the one in Table 3 eventually making it significant at the five per cent level while retaining the negative sign on it. These results strongly suggest that the contractionary effect of changes in the exchange rate on real output over the long-run is not

subject to the choice of the inclusion of the terms of trade variable in the model. The residuals obtained from the two estimated equations in Table 4 behaved almost in a similar fashion as the ones in Figure 6 and 7.⁴⁵

Table 4: PHFMOLS Estimates of the Long-run Relationship Excluding lnTOT

lnY =	6.039 ^{***}	+ 0.353 ^{***}	lnGE + 0.344 ^{***}	lnDC - 0.112 ^{**}	lnRERM
(s.e.)	(0.262)	(0.032)	(0.0257)	(0.05)	
t-ratio	28.36	10.85	13.36	-2.07	
Adjusted R ² = 0.996					
lnY =	6.477 ^{***}	+ 0.342 ^{***}	lnGE + 0.3286 ^{***}	lnDC - 0.136 ^{**}	lnRERIND
(s.e.)	(0.330)	(0.031)	(0.0265)	(0.047)	
t-ratio	19.60	11.01	12.37	-2.889	
Adjusted R ² = 0.996					

Note: Statistical significance at one, five, and 10 per cent level are given respectively by ^{***}, ^{**}, and ^{*}.

6.3. Short-run Dynamics

The estimated long-run relationships allow us to model the corresponding short-run dynamic adjustments using the error-correction mechanism. Following the Engle-Granger procedure the error-correction model incorporates the lagged residuals from the estimated long-run equation, which captures the time required to converge to the steady state relationship from any short-run deviations. With the error-correction models the common practice is to adopt the ‘general to specific’ modeling strategy of building a very general model by including the first difference of the variables along with their first or higher order lags and subsequently deleting the insignificant variables to arrive at the most parsimonious representation.

Table 5: Short-run Error-Correction Model with lnRERM

$\Delta \ln Y = 0.241^{***} + 0.178^{***} \Delta \ln DC + 0.094^{**} \Delta \ln GE - 0.058^{**} \Delta \ln RERM - 0.57^{***} \text{RESM}_{t-1}$
(s.e.) (0.004) (0.046) (0.036) (0.029) (0.155)
t-ratio 5.06 3.81 2.56 -2.03 -3.65
Diagnostic Tests
Adjusted R ² = 0.46
Serial Correlation: $\chi^2(1) = 2.68$
Normality: $\chi^2(2) = 1.00$
Functional Form: $\chi^2(1) = 1.68$
Heteroscedasticity: $\chi^2(1) = 0.57$

Note: ^{***}, ^{**}, and ^{*} are for statistical significance at the one and five per cent levels, respectively. The serial correlation test is based on Godfrey’s (1978) LM test for serial correlation; Functional Form on Ramsey’s (1969) RESET test; Heteroscedasticity on White’s (1980) test; and Normality of residuals on Jarque-Bera (1987) test. The computed test statistics for serial correlation, functional form and heteroscedasticity follow a chi-square distribution with one degree of freedom while the statistic for normality follows a chi-square distribution with 2 degrees of freedom.

⁴⁵ These are available from the authors on request.

Table 5 gives the resultant short-run model corresponding to the top row in Table 3.⁴⁶ It is observed that the short-run domestic credit and government expenditure elasticities are respectively 0.18 and 0.09. The terms of trade variable failed to register statistical significance and hence was dropped. Just like its long-run counterpart, the coefficient on $\Delta \ln \text{RERM}$ is negative and statistically significant. In the short-run, therefore, a one per cent rise in RER index results in about 0.06 per cent fall in aggregate output. The error-correction term, RESM_{t-1} , is correctly signed and significant at the one per cent level indicating a valid representation of the error-correction model. The coefficient suggests that it takes about two years to correct all short-run disequilibrium errors. The explanatory power of the short-run model is, however, low as only about 46 per cent variation in the growth of the real GDP can be explained by the right-hand side explanatory variables. For diagnostics Godfrey's (1978) LM test for serial correlation, Ramsey's (1969) RESET test for functional form, White (1980) test for heteroscedasticity and Jarque-Bera (1987) test for normality of errors are performed. The computed test statistics for serial correlation, functional form and heteroscedasticity follow a chi-square distribution with 1 degree of freedom, while the normality test statistic has a chi-square distribution with 2 degrees of freedom. Since the 95 per cent critical values for $\chi^2(1)$ and $\chi^2(2)$ are 3.84 and 5.99 respectively, on the basis of the computed diagnostic statistics we cannot reject the null hypotheses of no problem of serial correlation, no wrong functional form problem, normality of residuals and homoscedastic distribution of errors.

Table 6: Short-run Error-Correction Model with $\ln \text{RERIND}$

$\Delta \ln Y = 0.255^{***} - 0.024 \Delta \ln \text{TOT} + 0.169^{***} \Delta \ln \text{DC} + 0.086^{**} \Delta \ln \text{GE} - 0.049^{**} \Delta \ln \text{RERI} - 0.50^{***} \text{RESI}_{t-1}$						
(s.e.)	(0.005)	(0.027)	(0.049)	(0.042)	(0.03)	(0.161)
t-ratio	4.76	-0.89	3.42	2.03	-1.47	-3.13
Diagnostic Tests						
Adjusted $R^2 = 0.39$						
Serial Correlation: $\chi^2(1) = 1.87$			Functional Form: $\chi^2(1) = 1.63$			
Normality: $\chi^2(1) = 0.75$			Heteroscedasticity: $\chi^2(1) = 1.11$			

Note: ^{***}, ^{**}, and ^{*} are for statistical significance at the one and five per cent levels, respectively. The serial correlation test is based on Godfrey's (1978) LM test for serial correlation; Functional Form on Ramsey's (1969) RESET test; Heteroscedasticity on White's (1980) test; and Normality of residuals on Jarque-Bera (1987) test. The computed test statistics for serial correlation, functional form and heteroscedasticity are follow a chi-square distribution with one degree of freedom while normality test statistic follows a chi-square distribution with 2 degrees of freedom.

⁴⁶ Due to small sample size only the first lag of the first differenced variables were tried in the general model. None

The estimated parameters of the short-run model in Table 6 are, in general, comparable to those of Table 5. The coefficient on $\Delta \ln \text{RERIND}$ comes out to be negative but falls short of becoming statistically significant at the conventional levels.⁴⁷ Therefore, although the point estimate indicates a contractionary effect of devaluation, the confidence interval of the coefficient also contains a zero value for it. The error-correction term RESI_{t-1} , which is the lagged residual from the long-run relationship, is correctly signed and shows that it takes just about 2 years to converge to the long-run relationship from any short-run disequilibrium situation. The adjusted R^2 for the model is only 0.39 and but the diagnostic tests do not report any problem.

Table 7: Estimation of Short-run Models with additional dummy variables

Variables	Dependent Variable is: $\Delta \ln Y$			
	Coefficients (s.e.)	Coefficients (s.e.)	Coefficients (s.e.)	Coefficients (s.e.)
C	0.008 (0.008)	-0.027** (0.013)	0.0065 (0.01)	-0.0292 (0.036)
$\Delta \ln \text{DC}$	0.17*** (0.042)	0.199*** (0.033)	0.158*** (0.044)	0.179*** (0.038)
$\Delta \ln \text{GE}$	0.112*** (0.034)	0.176*** (0.033)	0.121*** (0.039)	0.188*** (0.04)
$\Delta \ln \text{RERM}$	-0.068** (0.026)	-0.11*** (0.024)	-	-
$\Delta \ln \text{REEIND}$	-	-	-0.057* (0.029)	-0.099*** (0.029)
RESM_{t-1}	-0.505*** (0.143)	-0.48*** (0.11)	-	-
RESI_{t-1}	-	-	-0.453*** (0.147)	-0.42*** (0.12)
D89	0.016** (0.007)	0.025*** (0.006)	0.018*** (0.008)	0.0276*** (0.008)
D86	-	0.024*** (0.007)	-	0.022** (0.008)
		Diagnostic Tests		
Adjusted R^2	0.57	0.75	0.50	0.66
Serial Correlation: $\chi^2(1)$	0.21	0.12	0.038	0.55
Functional Form $\chi^2(1)$	4.28**	2.64	3.38	3.24
Normality: $\chi^2(1)$	0.89	0.53	0.918	0.61
Heteroscedasticity: $\chi^2(1)$	0.35	0.0015	0.07	0.009

Note: Statistical significance at the one, five and ten per cent levels are indicated by ***, **, and *. D89 is the dummy variable – 0 for 1989 and 1 for otherwise and similarly D86 is another dummy variable with 0 for 1986 and 1 for otherwise.

of these variables were found to be statistically significant justifying their deletion.

⁴⁷ The coefficient is significant only at the 20 per cent level.

As mentioned earlier, both the short-run models have somewhat low explanatory powers in contrast to a very high R^2 associated with their counterpart long-run models. It is possible that in the short-run many other factors influence aggregate output growth about which the theory is silent. This should not point toward any estimation or misspecification problem especially when the diagnostic tests do not detect any such problem. Nevertheless, an attempt was made to see whether such low explanatory power could be attributable to any particular atypical year. For both models in Tables 5 and 6, the largest error was associated with the year 1989. When a dummy for this year was incorporated into the model, the explanatory power was increased to 0.57 for the equation with RERM and to 0.50 with RERIND (see Table 7). Insertion of just another dummy for 1986 increased the adjusted R^2 to 0.75 and 0.66 respectively. Interestingly, Table 7 shows that the regressions with the dummy variables (either only for 1989 or for both 1989 and 1986) make the coefficient on $\Delta \ln RERIND$ statistically significant, while the significance of $\ln RERM$ becomes even more prominent.⁴⁸ These results seem to suggest that significant contractionary effect of downward adjustment of the RERIND on the aggregate output in the short-run is overshadowed by influential observations like 1989 and 1986.

7. Summary of Findings and Implications

The effect of the exchange rate on aggregate output has been a longstanding controversy in applied macroeconomics. The issue is very important for a country like Bangladesh where exporters often demand for downward adjustment of the domestic currency in order to become more competitive in international markets in sharp contrast to counter-productive arguments of devaluation put forward by the consumers and the industrial circle relying on the imported goods for consumption and production. Apart from these, the government has a principal objective of maintaining sustainable trade balance for which often currency adjustment becomes essential. Theoretical possibilities for having both contractionary and expansionary effects of devaluations on aggregate output would imply that for any country the net impact has to be determined empirically. This paper has made such an attempt to study the output effects of exchange rate changes in Bangladesh by using the new national income accounting data for 1980-2000.

⁴⁸ Since $\Delta \ln TOT$ was not significant in various regressions, it has been deleted from the results reported in Table 3.8.

The empirical specification used in this paper is derived from a three-market Keynesian model that posits a long-run relationship between the real GDP and a vector of right hand side variables including terms of trade, government expenditure, domestic credit and real exchange rate. For empirical investigation this study constructs two series of multilateral real exchange rate: one is based on the weights associated with Bangladesh's bilateral formal trade with top 20 countries (RERM) while the other (RERIND) still considering the same partners but assigns a greater weight to India due to the existence of a large volume of informal border trade.

The estimation strategy considered examination of the time series properties of the variables in the regression to avoid the problem of estimating a spurious relationship. In light of the non-stationarity of the model variables, cointegration techniques were used to validate the long-run relationship. The results show that in the long-run the real exchange rate variables are associated with a negative sign, which would imply contractionary effects of devaluation for Bangladesh. However, the contractionary effect appears to be statistically significant only in equations that use either the multilateral RER purely based on formal trade or excludes the terms of trade variable. In order to capture the short-run dynamics, error-correction models were constructed and again the results were in line with the long-run equations. That is, significant contractionary effect was found for the equation with RERM but when RERIND was used a negative but insignificant impact was observed. The paper, however, reveals that in the short-run the negative coefficient on RERIND becomes significant if just one outlier is forced to fall on the fitted line.

Therefore, the main finding of the study can be summarized as follows: There is no evidence of any expansionary effect of devaluation on output. Point estimates of RERs are always negatively signed suggesting that downward adjustments of the taka adversely affect overall production.

Several reasons can be held responsible for the above result. First of all, it is not very difficult to perceive that the scope of enhancing external competitiveness through nominal devaluations is actually limited for Bangladesh. The country's export basket is dominated by ready-made garments, which are overwhelmingly dependent on imported raw materials and equipments. For a small open economy import prices in foreign currency are fixed and consequently when the home currency depreciates import prices will rise by the full extent. This severely reduces the

capacity of exporters in a heavily-import dependent industry to benefit from devaluation. Moreover, when the backward integration of any particular export industry is weak, whatever price incentive given to it fails to generate adequate supply stimulus that would eventually outweigh the negative effect of rising costs of production in all other sectors following devaluation. For the same reason, the argument that other countries improve competitiveness and expand output by devaluing their currencies may not justify devaluation in Bangladesh. Considering the case of textiles and clothing there is no denying that Bangladesh's competitors, such as China, India, Korea and Pakistan, enjoy a greater amount of domestic value added and, as a result, devaluation of their currencies might not increase the cost of production by as much as would be in the case of Bangladesh's producers. Thus, while devaluation might constitute a competitiveness argument for exporters leading to expanded economic activities in those countries, for Bangladesh such a scope would be severely limited.

For all other exports with relatively large domestic content, such as, jute and jute goods, tea, frozen fish, and leather and leather products can devaluation be argued to be an effective supply stimulus? The problem of jute and jute goods is well-known as they compete with cheap synthetic substitutes in world market and it is doubtful whether devaluation alone can protect its competitiveness in the long-run. On the other hand, for tea, frozen fish, and leather products Bangladesh is a very small supplier of these commodities and the relative significance of these sectors in overall GDP is not prominent enough to generate adequate supply response.⁴⁹ In fact it is the supply capacity that is most important for increasing exports and providing only price incentives through devaluation is not sufficient for achieving an expanded capacity.

Bangladesh is a country with a high ratio of import to GDP (currently about 18 per cent) and a significant proportion of these imports is in the form of intermediate inputs, raw materials, and plants and machinery, which cannot be produced domestically but are essential to the country's production processes in the non-export and non-traded sectors. By making these imports costlier devaluation might have resulted in reduced profits and funds to be reinvested in these sectors thereby adversely affecting output and economic growth. Very little is known about how

⁴⁹ Total receipts for Bangladesh from combined exports of leather, tea and frozen food stood at US\$ 0.8 billion in 2000.

consumers substitute between imported goods and domestically produced goods (or, between tradables and non-tradables). If imports are price inelastic, it could be possible that following devaluation consumers cut back spending on home goods to offset the price rise of imported items. This can result in contractionary effect in the import-competing and non-traded sectors.

It is to be acknowledged that maintaining a sustainable external balance under a fixed or managed system of exchange rate is a challenging task. In the face of an unsustainable trade balance it often requires downward adjustment of the nominal exchange rate, which is undoubtedly among the most unpopular policy decisions. On the other hand, under a flexible exchange rate system the adjustment is automatic and the policy makers do not have to decide whether to devalue or not although the government can exert some influence in a market friendly way. In recent times, Bangladesh has been suggested to adopt a freely floating exchange rate regime and the government is seriously contemplating a move to that direction. Free-floating will certainly ensure long-run equilibrium in the balance of payments but in the absence of a sound management there might be short-run fluctuations which are destabilizing in nature, hostile towards inflow of foreign capital, and impede domestic investment decision both in the traded and non-trade sector. There is also an apprehension that free floating will result in considerable depreciation of the taka. The results obtained from our study imply that such a situation might have serious consequences for Bangladesh as a 10 per cent devaluation is found to be associated with as high as 1.3 per cent decline in aggregate output. Therefore, it is important to strike a delicate balance between maintaining a sustainable trade balance and ensuring growth in overall output. A smooth transition toward flexible system along with the central bank's capacity of absorbing shocks and preventing rapid depreciation would have important bearings on the growth performance of the economy.

References

- Atkins, F. (2000), "Revisiting the Exchange Rate Debate: The Jamaican Experience", *Journal of International Trade and Economic Development*, vol.12, pp.121-131.
- Bahmani-Oskooee, M. (1998), "Are Devaluations Contractionary?", *Journal of Economic Development*, vol.23, No. 1, pp.131-43.
- Bangladesh Bank (2002), *Economic Trends*, Statistics Department, Bangladesh Bank, August.
- BBS (2000), *National Accounts Statistics of Bangladesh (Revised Estimates, 1989-90 to 1998-99)*, BBS, December.
- BBS (2001), *National Accounts Statistics (Gross Domestic Product, 2000-2001)*, BBS, June.
- Buffie, E. (1986), "Devaluation, investment, and growth in LDCs", *Journal of Development Economics*, pp.361-79.
- Chowdhury, O. H. (1995), "Fiscal Performance of Bangladesh", Chapter 5 in Sobhan, R. (ed) *Experiences With Economic Reforms: A Review of Bangladesh's Development 1995*, University Press Limited.
- Connolly, M. (1983), "Exchange Rates, Real Economic Activity and the Balance of Payments: Evidence from the 1960s", in E. Classensn, and P. Salin (eds.) *Recent Issues in the Theory of Flexible Exchange Rates*, Amsterdam: North Holland.
- Connolly, M. (1983), "Exchange rates, real economic activity and balance of payments: Evidence from the 1960s" in E. Classensn and P. Salin (eds.) *Recent Issues in the Theory of the Flexible Exchange Rates*, North Holland.
- Diaz-Alejandro, C. (1963), "A note on the impact of devaluation and the redistributive effect", *Journal of Political Economy*, vol71, pp.577-80.
- Diaz-Alejandro, C. F. (1965), *Exchange Rate Devaluation in a Semi-Industrialised Economy: The Experience of Argentina 1955-61*, MIT Press.
- Dornbusch, R. (1980), *Open Economy Macroeconomics*, New York.
- Edwards, S. (1986), "Are Devaluations Contractionary?", *Review of Economics and Statistics*, pp.501-508.
- Edwards, S. (1986), "Are devaluations contractionary", *Review of Economics and Statistics*, pp.501-508.
- Edwards, S. (1989), *Real Exchange Rates, Devaluation, and Adjustment: Exchange Rate Policies in Developing Countries*, MIT Press.

Gylfason, T. and Radetzki, M. (1985), "Does Devaluation Make Sense in the Least Developed Countries?" IIES Seminar Paper, University of Stockholm.

Gylfason, T., and Schmid, M. (1983), "Does Devaluation Cause Stagflation?" *Canadian Journal of Economics*, pp.641-54.

Gylfason, T. and Schmidt, M. (1983), "Does devaluation cause stagflation?" *Canadian Journal of Economics*, vol.16, pp.641-54.

Hanson, J. (1983), "Contractionary Devaluation, Substitution in Production and Consumption, and the Role of Labour Market", *Journal of International Economics*, vol.14., pp.179-189.

Krugman, P. and Taylor, L. (1978), "Contractionary Effects of Devaluation", *Journal of International Economics*, pp.445-456.

Lizondo, J.S. and Montiel, P. (1989), "Contractionary devaluation in developing countries: an analytical overview", *IMF Staff Papers*, vol.36.

Rhodd, R. (1993), "The effects of real exchange rate changes on output: Jamaica's devaluation experience", *Journal of International Development*, vol.5, pp.291-303.

Upadhyaya, K. and Upadhyay, M. (1999), "Output effects of devaluation: Evidence from Asia", *Journal of Development Studies*, vol.35, pp.89-103.