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The Exchange Rate and Economic Growth

An Empirical Assessment on Bangladesh

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Abstract

This paper aims to understand the effects of exchange rate changes on economic growth in Bangladesh. It makes use of a Keynesian analytical framework to derive an empirical specification, carefully constructs a real exchange rate series, and employs cointegration techniques to determine the output response to taka depreciations. The results show that in the long run a 10 per cent depreciation of the real exchange rate is associated with a 3.2 per cent rise in aggregate output. A contractionary effect is however observed in the short-run so that the same magnitude of real depreciation would result in about half a per cent decline in GDP. While the long-run expansionary effect of real depreciations may appeal for considering the role of exchange rate policy as a development strategy, the evidence of very high degree of exchange rate pass-through to consumer prices would severely constrain such an option. For Bangladesh the need for maintaining external competitiveness and promoting growth remains a delicate task for policymakers as it involves managing an exchange rate regime accompanied by other consistent macroeconomic policies. Notwithstanding, rather than pursuing an effectively fixed nominal exchange rate until external imbalances become unsustainable, a more pragmatic approach will be to tolerate creeping depreciations that would avoid any significant contractionary effect in the short run while allowing for improved competitiveness and output growth in the long run.

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I. Introduction

During the two-year period ending in July 2012, Bangladesh experienced an almost 20 per cent fall in the value of its currency with respect to the US dollar. Unlike many other developing countries, Bangladesh has been successful in avoiding large and abrupt rise in nominal exchange rates. Therefore, although a 15-taka rise in dollar price within the period mentioned above may not appear to be inordinate in comparison with other country experiences, such a magnitude of depreciation over a comparable period has been unprecedented in Bangladesh.[1] At the same time, the country's macroeconomic situations had also come under some strain because of rising prices on the one hand, and growing deficit in the trade account of the balance of payments on the other. This put the policymakers in a dilemma: while the depreciation of taka to boost export competitiveness, attract remittances, and discourage imports should be a natural response to correct the balance of payments problems, such an option was likely to feed into the inflationary pressure.

The management of the exchange rate is considered to be a major policy objective in Bangladesh to achieve a set of diverse objectives of economic growth, containment of inflation and maintenance of external competiveness. Policy discussions regularly emphasise on it as the academic literature provides compelling evidence to suggest that a wrongly managed exchange rate regime can be a major impediment for improved economic performance. Reform of the exchange rate management was an important component of trade liberalization measures that Bangladesh undertook, eventually replacing the earlier 'fixed rate' system with a 'freely-floating' regime.

Despite putting in place a supposed-to-be market based mechanism in 2003, Bangladesh's exchange rate regime continues to be characterised widespread interventions (Hossain and Ahmed, 2009). One important reason for not allowing market forces to work is the concern

[1] The rate of appreciation reported here is based on the official nominal exchange rates, published by the Bangladesh Bank.

about the inflationary pressure impinged through currency depreciations. Indeed, attempts have apparently been made to maintain the nominal exchange rate effectively fixed for as long as possible before external imbalances become unsustainable. This in the process has disregarded the policy emphasis on maintaining the country's external competitiveness.

The apprehension about currency depreciation is well-understood. For a small country devaluations will lead to increase in import prices and subsequently prices of other products and services through feedback effects. However, the growth implications of currency adjustments are generally overlooked. According to the 'orthodox' economic theory, devaluation of a country's currency triggers an "expenditure switching" mechanism, which leads to domestic demand away from imports to locally produced import-competing goods. It also improves international competitiveness thereby boosting exports. These two effects together exert an expansionary effect on overall economic activity. In contrast, there are several reasons why devaluation can decrease an economy's output growth. Amongst them, the rise of prices of imports affecting the domestic production is the most dominant one. The empirical evidence on the effects of currency adjustments on output is mixed and consequently country-specific case studies are the best possible option to guide policy directions.

While the issue of nominal devaluations attracts so much attention in Bangladesh, discussions surrounding it mainly focus on inflation and are generally without any reference to overall economic activity. This generates a lack of attention to the role of exchange rate management for promoting economic growth and maintaining the external competitiveness. The present paper therefore aims to contribute to the relevant macro policy discourse more effectively by carrying out an empirical investigation of the effects of exchange rate changes on economic growth.

This paper is organized as follows: after this introduction, Section II provides a brief review of the literature; Section III presents a theoretical framework based on which the empirical exercise will be undertaken; Section IV describes the data used in the paper and the construction of the real exchange rate for Bangladesh; Section VI outlines the methodology used for empirical assessment and presents the results; finally Section VII provides a summary of main findings along with their policy implications.

II. A Brief Review of the Literature

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 in import competing sectors. Devaluations are also believed to contribute to the enhancement of external competitiveness stimulating production in the export sector. 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, and many developing countries have relied upon devaluations 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 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.[3] 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-Alezandro (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

[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.

[3] A detailed review of the channels through devaluations can exert a contractionary effect, see Acar (2000).

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 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.

There are numerous studies that have assessed the relationship between exchange rate movements and economic growth. In general, earlier studies such as Connolly (1983), Gylfason and Schmid (1983), Krueger (1978), Taylor and Rosensweig (1984), and Kamin (1988) tended to provide support for expansionary effects of devaluations. Subsequently, the contractionary effects became prominent in a large number of studies such as Gylfason and Radetzki (1985), Atkins (2000), Kamin and Roger (2000), Odusola and Akinlo (2001), Berument and Pasaogullari (2003), El-Ramly and Abdel-Haleim (2008). Mixed results have also been reported in a number of studies. For example, Edwards (1986) and Rhodd (1993) found negative (contractionary) short-run effects but in the long-run the output response to devaluation appeared to be positive. El-Ramly and Abdel-Haleim (2008) reveals negative response for several years before expansionary effects can show up. There are also studies that do not find any significant effect of exchange rate movements (e.g., Bahmani-Oskooee, 1998 and Upadhyaya and Upadhyay, 1999). Studies considering multiple countries have often reported differing findings. For instance, Bahmani-Oskooee and Miteza (2006) using a panel of 42 countries find that in the long-run devaluations are contractionary in non-OECD countries, while for OECD economies the results are mixed.

There is another branch of work devoted to examine if the exchange rate regimes of developing countries have been overvalued. Popularised by Edwards (1986), this line of research hypothesises that a country's equilibrium exchange rate is determined by a set of 'fundamentals', but various factors such as distortionary monetary and fiscal policies can cause the actual rate to deviate from the equilibrium rate. In comparison with the equilibrium values, one can then assess if the real exchange rate is overvalued. Overwhelming evidence exists and there seems to be a consensus that overvaluation has adverse implications for growth (Dollar, 1992; Easterly, 2005).

Finally, of late a new literature has sprung up showing that undervalued exchange rates foster economic growth. In one of the most widely cited papers of recent times, Rodrik (2008), having constructed an index of undervaluaton based on a purchasing power parity real exchange rate for countries, demonstrates robust evidence of growth-enhancing effect of undervalued currencies. He argues that tradable sectors are more severely affected by bad institutions and market failures, resulting in their size being smaller than optimal. Undervaluation of national currencies helps overcome these problems. The successful experiences of East Asian countries in general, and China in particular, seem to extend support to this findng. This is in contrast to the so-called Washington Consensus view that undervaluation will lead to overheating and excessive inflation to harm the economy (Berg and Miao, 2010).[4]

Rodrik's assessments have come under close scrutiny with subsequent studies (e.g. Gulzmann et al. 2012; Mario et al., 2011; Mbaye, 2012; and Rapetti et al., 2011) however generally confirming the positive association between higher growth and undervalued exchange rate. However, the mechanism through which undervaluation works remains a subjectmatter of debate (Magud and Sosa, 2010). Gulzmann et al. (2012) provides results to argue that depreciated exchange rate does not influence the tradable sectors as suggested by Rodrik, but it is through increased saving and investment that growth is facilitated. In contrast, Mbaye (2012) finds the evidence of total factor productivity growth as a result of

[4] Berg and Miao (2010) however try to reconcile these differing outcomes. According to them, undervaluation may be supportive of higher growth in the short to medium term after which adjustments take place to return to the normal output path.

undervaluation. Mario et al. (2011) unearths further evidence that higher RER helps diversify exports and raise technological intensity of exports.

However, it needs to be pointed out that the notion of undervaluation is essentially based on a ‘norm’ or equilibrium rate: Rodrik estimates undervaluation comparing with the PPP adjusted rates. Devaluations and currency depreciations may not be associated with undervaluation. Nevertheless, since undervaluations are essentially related to depreciated nomial exchange rates, it has an appealing factor in thinking about a role of the exchange rate policy in development strategies. However, it has been argued that developing countries may not have the capacity of maintaining a depreciated exchange rate. This policy choice is likely to cause tension with other countries and would invite costly and financially disruptive inflation (Eichengreen, 2008).

III. Analytical Framework

Most recent analyses looking into the relationship between the movements in the exchange rate and economic activity do not follow any particular theoretical framework to derive a specific model for empirical estimation. This is facilitated by the recently popular econometric method of the Vector Autoregression (VAR) analysis. This approach is essentially a non-structural framework in which no particular relationships are imposed on the variables based on economic theory. As long as there is some hint of certain variables being related, a VAR model can be employed to investigate the empirical relationship. Recent advances in econometrics (e.g. unit roots and cointegration methods) and wider recognition that many macro variables are jointly determined (i.e. endogenous) have made widespread use of VAR models. This is particularly true for the recent studies on the impact of exchange rate changes on output (e.g. Kamin and Rogers, 2000; Odusola and Akinlo, 2001, Vinh and Fujita, 2007; El-Ramly and Abdel-Haleim 2008). Nevertheless it remains extremely important to use a sound theoretical framework to explicitly understand the underlying mechanism through which the variables of interest influence any particular outcome. It can also help determine the most appropriate relationships amongst a set of possible interactions often reported by atheoretical VAR analyses.

Amongst others, Edwards (1986) while analysing the effects of stabilization programmes on aggregate production in developing countries develops a useful model linking exchange rate and overall output. This became one of the most influential analytical tools in guiding the empirical analysis on the effects of devaluation on output. Based on a simple three-market Keynesian model, another useful framework is due to Rhodd (1993), which this paper makes use of to derive a reduced form equation for empirical estimation.

In Rhodd's model, the goods market is represented by:[5]

$$Y = C + I + G + X - M \tag{3.1}$$

Or,

$$Y - C - G = I + X - M \tag{3.2}$$

$$S = I_d + I_f \tag{3.3}$$

$$S = S(Y, r);\ \frac{\partial S}{\partial Y} > 0,\ \frac{\partial S}{\partial r} > 0 \tag{3.4}$$

$$I_d = I_d(Y, r);\ \frac{\partial I_d}{\partial Y} > 0,\ \frac{\partial I_d}{\partial Y} < 0 \tag{3.5}$$

$$I_f = I_f(Y, e);\ \frac{\partial I_f}{\partial Y} < 0,\ \frac{\partial I_f}{\partial e} \qquad I_{fY} < 0,\ I_{fe} > 0 \tag{3.6}$$

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 (3.3) shows the equilibrium between aggregate demand and aggregate supply. Equations (3.4), (3.5) and (3.6) show how *S*, I_d and I_f are determined in the model. Foreign investmemt (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.

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.

$$M_{s0} = M_d \tag{3.7}$$

[5] This section draws on Rhodd (1993).

$$M_d = L(Y,r); \quad L_Y = \frac{\partial M_d}{\partial Y} > 0, \qquad L_i = \frac{\partial M_d}{\partial r} < 0 \tag{3.8}$$

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) (3.9)

$$\frac{\partial B}{\partial Y} < 0, \qquad \frac{\partial B}{\partial r} = ? \tag{3.10}$$

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

$$S_0 + S_1Y + S_2r - I_{d0} - I_{d1}Y - I_{d2}r - I_{f0} - I_{f1}Y - I_{f2}R = 0 \tag{3.11}$$

$$L_0 + L_1Y + L_2r = M_s \tag{3.12}$$

$$T_0 + T_1Y + T_2e + F_0 + F_1Y + F_2r - B = 0 \tag{3.13}$$

Equations (3.11) – (3.13) can be written in matrix form to give:

$$\begin{bmatrix} (S_1 - I_{d1} - I_{f1}) & (S_2 - I_{f2}) & 0 \\ L_1 & L_2 & 0 \\ (T_1 + F_1) & F_2 & -1 \end{bmatrix} \begin{bmatrix} Y \\ r \\ B \end{bmatrix} = \begin{bmatrix} -S_0 + I_{d0} + I_{f0} + I_{f2}e \\ M_s - L_0 \\ T_0 - T_2e - F_0 \end{bmatrix} \tag{3.14}$$

From (2.14) Y can be determined, which is given by:[6]

$$Y = \frac{(L_2S_0 - L_2I_{do} - L_2I_{f0} - L_2I_{f2}e + M_sS_2 + I_{f2}M_s) - (S_2L_0 + I_{f2}M_s)}{D} \tag{3.15}$$

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

$$\frac{\partial Y}{\partial e} = -\frac{L_2I_{f2}}{D} > 0 \tag{3.16}$$

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

[6] One can use the Cramer's rule to determine Y.

Therefore, the empirical model will be based on equation (3.15) that shows the relationship between real output, a measure of monetary policy as indicated by M_s, the exchange rate, and public 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.

IV. Empirical model and data

Our empirical specification will closely follow the analytical framework presented above. The theoretical model represents a long-run relationship between aggregate output and a number of other variables comprising RER, a measure of fiscal policy, and an indicator of monetary policy. However, a large number of empirical studies e.g. Atkins (2000), Edwards (1986), Rhodd (1993), and Upadhyaya and Upadhyay (1999) also include respective countries' external terms of trade (TT). For a small open economy TT is truly exogenous and when not controlled for explicitly in the experiment, some of its unaccounted for influence could be transmitted through the indicator of external competitiveness (RER). The real exchange rate is often considered to be the terms of trade of the country. However, for many countries, the movements in their TT and RER are quite different. Therefore, the distinct effect of TT cannot be captured by the real exchange rate variable. For the present paper, we include TT in our main empirical model.

There seems to be a consensus in the literature with regards to the use of government expenditure as a measure of fiscal policy stance, but different indicators have been used to represent monetary policy. For example, Edwards (1986) and Upadhyaya and Upadhyay (1999) construct a 'money surprise' or unexpected money growth term, Atkins (2000) and Rhodd (1993) consider total domestic credit. The estimation of the money surprise function for Bangladesh, as specified by Edwards (1986), was not found to be satisfactory.[7] On the

[7] Edwards defines money surprise as the actual rate of growth of nominal money (Δlog M) less the expected rate of growth of nominal money (Δlog 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 Δlog M are subtracted from the actual money growth to arrive at the surprise money growth series.Rhodd (1993) reports that money surprise functions also does not work out satisfactorily in his

other hand, the use total domestic credit could 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. One problem however is that a part of domestic credit is also included in government expenditure. Therefore, we decided to use credit to the private sector as an indicator of monetary policy. Currently, close to 80 per cent of domestic credit is due to the private sector and the limited autonomy of the central bank would imply that the impact of monetary policy is mainly driven through the private sector credit flows.[8] Using the logarithmic transformation of the variables the empirical specification of the model thus can be written as:[9]

$$\ln Y_t = \beta_0 + \beta_1 \ln(GE_t) + \beta_2 \ln(PCR_t) + \beta_3 \ln(TT_t) + \beta_4 \ln(RER_t) + \upsilon_t$$

where, *ln* stands for natural logarithm, time is denoted by subscript *t*, *Y*, *GE*, *PCR, TT,* and *RER* stand respectively for real GDP, real government expenditure, credit to the private sector, terms of trade, and real exchange rate. υ represents the error term. It is expected that β_1 and β_2 are positive while the sign of β_3 cannot be determined a priori. The coefficient β_4 capturing the effect of real devaluation on real output growth is the primary interest of this study and its sign cannot also be predetermined.

Data required for empirical estimation have been gathered from different sources. We use the GDP data from the annual publications of the Bangladesh Bureau of Statistics. BBS undertook a major revision of GDP calculation in the late 1990s and then revised various components of economic activities retrospectively, but going back up to 1980 only. Compared to the previous estimates, this revision resulted in upward adjustments of national output by 26-43 per cent for different years. As no estimates for pre-1980 period was attempted, and given that there was no general statistical relationship was used in revising the national income data, it is not possible to extend the series in a satisfactory manner to include information for the 1970s. Therefore, the sample period for this paper

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empirical investigation. Furthermore, in most regressions of Upadhyaya and Upadhyay (1999) money surprise terms were not significant.

[8] Kamin and Rogers (2000) and El-Ramly and Abdel -Haleim (2008) have used broad money supply in their empirical specification.

[9] 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.

has been confined to 1980-2012. Bangladesh also does not have a quarterly series of GDP. Consequently, the empirical exercise uses annual data.

Government expenditure comprises government consumption expenditure (recurrent expenditure) as well as public investment expenditure allocated via the annual development plans. These data have been gathered from various government documents including the *Bangladesh Economic Survey* published by the Ministry of Finance. Data on credit to the private sector come from Bangladesh Bank publications. The data on government expenditure and domestic credit are initially gathered in current prices. Using the implicit GDP deflator these data are converted into real prices.

The TT index has been estimated from the reported unit value indices for exports and imports in UNCTAD (2012), which provides historical time series data for most developing countries including Bangladesh. The TT is defined here as the unit value index for exports divided the unit value index for imports.

Finally, the variable, RER, has been carefully constructed. This has been a relatively involved task and given the importance, its definition and construction procedures and data sources are explained below.

Construction of the Real Exchange Rate

Discussions surrounding the effects of currency appreciation/depreciation on economic activity usually focus on the changes in the 'real' rather than the 'nominal' exchange rate. The real exchange rate (RER) has two main strands following from the purchasing power parity (*PPP*) and trade theory definitions. According to the *PPP* theory, exchange rate movements are determined by the difference between the domestic and foreign rates of inflation. When domestic inflation relative to changes in foreign prices rises, 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*.[10] On the other

[10] Often the nominal exchange rate used to illustrate the *PPP* theory, 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,

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.

Recent empirical studies, however, have most frequently used the trade theory definition of *RER*. 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. Following Edwards (1989), it has now become a standard practice to use the world price of tradables (P^f_T) as a proxy for P_T, and domestic price of non-tradables (P^d_{NT}) for P_{NT}. Therefore, the operational definition of *RER* can be represented as:

$$RER = E\left[\frac{P_T^f}{P_{NT}^d}\right] \qquad (1)$$

The computation of RER based on (1) would require using proxies for the two price indices. Most empirical analysts, following Harberger (1986) and Edwards (1989), use the foreign whole sale price index (WPI^f) as a proxy for P^f_T and the domestic consumers' price index (CPI^d) for P^d_{NT}.[11] The foreign price index is generally constructed on the basis of a weighted

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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: $lnR = \theta + lnP_f - lnP_d$,and under the absolute version θ will be equal to zero. However, if we consider a change in the exchange rate then $\Delta lnR = \Delta lnP^f - \Delta lnP^d$. 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: $lnR = \theta + \beta_1 lnP_f + \beta_2 lnP_d + v_1$, 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.

[11] 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.

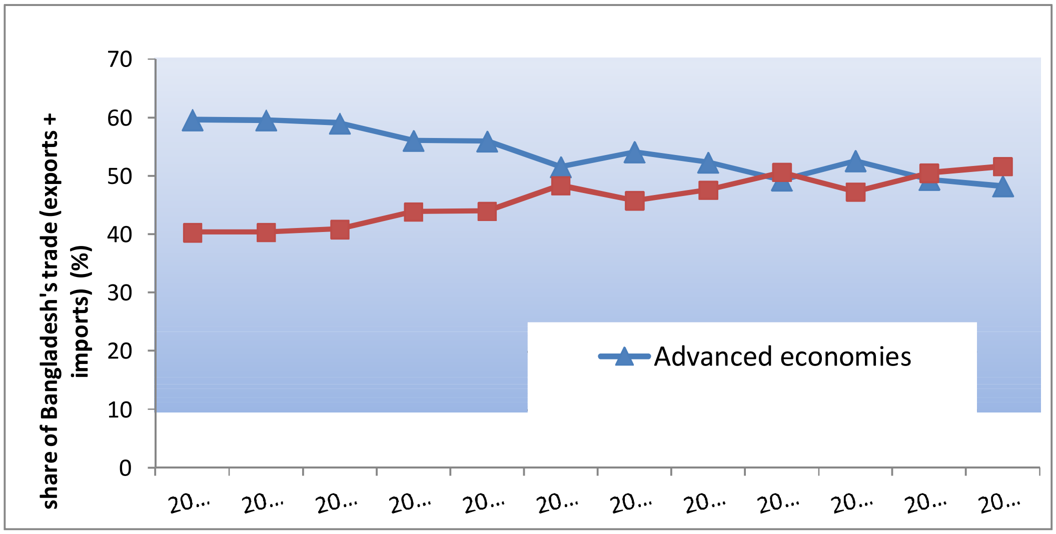
average of WPIs of several important trade partners with weights (α) corresponding to the share of each partner in the country's trade transactions (i.e. either exports or imports or both). Equation (2) thus provides the formulation of *RER*, where subscript *t* denotes time.

$$RER_t = \left[\sum_{i=1}^{k} \alpha_{it} \frac{E_{it} \; WPI_t^{\,f}}{CPI_t^{\,d}} \right] \text{ with } \sum_{i=1}^{k} \alpha_i = 1 \qquad (2)$$

For constructing the RER series, initially Bangladesh's top 25 most important trade partners were chosen in terms of total value of trade transactions (*i.e.* exports plus imports) during 2007-11.[12] These countries account for almost 80 per cent of Bangladesh's trade (see Appendix Table A1). However, we decided to drop Brazil, Indonesia and Turkey as they had experienced sudden and massive shocks in their nominal exchange rates and consequently their inclusion would result significant movements in Bangladesh's RER despite their being relatively less important trade partners. Finally, the UAE was excluded because of lack of any price data (either WPI or CPI) for the full sample period. The remaining 21 countries contributed to just about three-quarters of Bangladesh's trade. The weights worked out are: China (0.141), India (0.127), the USA (0.123), Germany (0.086), the UK (0.056), Singapore (0.052), Kuwait (0.041), Japan (0.039), France (0.036), Hong Kong (0.033), Malaysia (0.032), Korea (0.032), Canada (0.030), the Netherlands (0.029), Italy (0.028), Thailand (0.023), Australia (0.021), Spain (0.020), Saudi Arabia (0.019), Belgium (0.018), and Pakistan (0.014).

[12] Data on Bangladesh's exports to and imports from different countries have been obtained from the Direction of Trade Statistics of the IMF.

Figure 1: Changing relative significance of trade partners' of Bangladesh: Advanced vis-à-vis developing economies



Note: Author's estimate from IMF data

It may be interesting to note that during the most recent past decade noticeable changes in terms of Bangladesh's most significant trade partners have taken place. Advanced economies (such as Europe and the USA) have traditionally been the dominant trade partners but since 2000 their share has declined from 60 per cent to less than 50 per cent. Largely because of being import sources a number of developing countries including China and India have emerged to be important trading partners of Bangladesh. Until now, Bangladesh's exports remain overwhelmingly dependent on developed country markets.

The data on partners' *WPI* and Bangladesh's *CPI* were gathered from IMF (2012).[13] We modified the P^f_T formula as in equation (3) so that each partners' exchange rate with the US dollar (i.e. currency units per dollar) could be used in constructing the RER. It is important to mention here that since several European countries accepted the euro as the single currency, bilateral dollar exchange rates with respect to their previous respective currencies ceased to exist from 1998. If one is to use the euro, the construction of RER would be possible only from 1999 thereby giving a very short sample to do any useful econometric analysis. To deal with this problem, we computed their implicit bilateral exchange rates with

[13] Note that the data on WPI are not available for China, France, and Hong Kong. For these countries consumers' prices (CPI) have been used. As correlations between WPI and CPI are usually high, the choice of price series should not make major differences.

the dollar for the period 1998-2012, using the information on the relationship between the dollar and special drawing rights (SDR) exchange rates of the defunct currencies and the relationship between the euro and SDR since 1999. Finally, equation (3) was divided by Bangladesh's *CPI* and then multiplied by the end period exchange rate of Bangladesh taka vis-à-vis the US dollar.[14] The way it is constructed a higher value of RER indicates real depreciation while a lower value would imply depreciation.

$$P_T^f = \sum_{i=1}^{k} \alpha_i \left(\frac{WPI_i}{e_i} \right) \tag{3}$$

Figure 2 exhibits the graphical plots of RER (indexed to 1995-96=100) and the nominal exchange rate (NER) (tk/dollar). It shows that while the taka value of one dollar has shown a steady rise throughout the sample period, the RER has hardly changed since the 1980s. Therefore, numerous dosages of devaluations administered during the period of pegged exchange rate regime (until 2003) and frequent adjustments taken place since the introduction of a more flexible regime have mainly resulted in cancelling out the inflation differentials between Bangladesh and its trading partners. This is reflected in Figure 3. Since 1996, Bangladesh's domestic price level, measured by the consumers' price index, has increased by about 175 per cent as against of trade partners' rise in prices by only 50 per cent. Therefore, more than a 100 per cent nominal devaluation of taka since 1996 has barely been enough to prevent deterioration in competitiveness.

[14] The exchanges rate between dollar and other partners' currencies are the period average rate as reported in the International Financial Statistics of the IMF. 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 calculations showed that there would not have been any significant difference had we used the period average nominal exchange rate of taka.

Figure 2: Movements in Bangladesh's exchange rates: Real vis-à-vis nominal

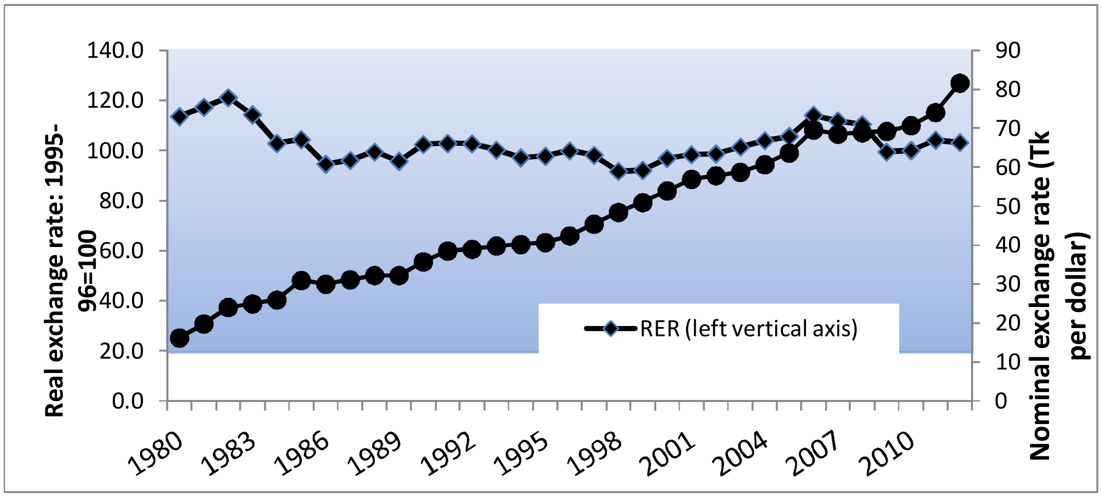


Figure 3: Comparison of prices – Bangladesh and trade partners

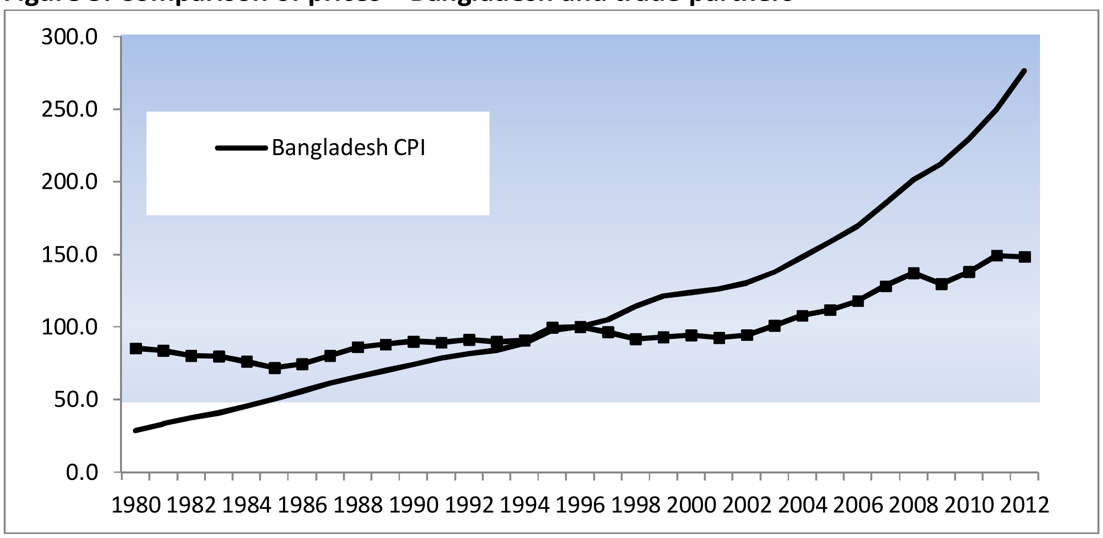


Figure 4 provides a simple relationship between nominal devaluations and changes in the RER. During 1980-2012, only on two instances (1988 and 2003) were the changes in RER greater than those of NER. The scatter plot in Figure 5 depicts a positive but rather a weak relationship between the changes in NER and RER. The regression equation shows that a one percentage point change in the nominal rate is associated with just 0.32 percentage point rise in the RER index. However, this relationship must be considered with some caution, as Figure 5 shows the relationship has a large variance.[15]

Figure 4: Changes in nominal and real exchange rates

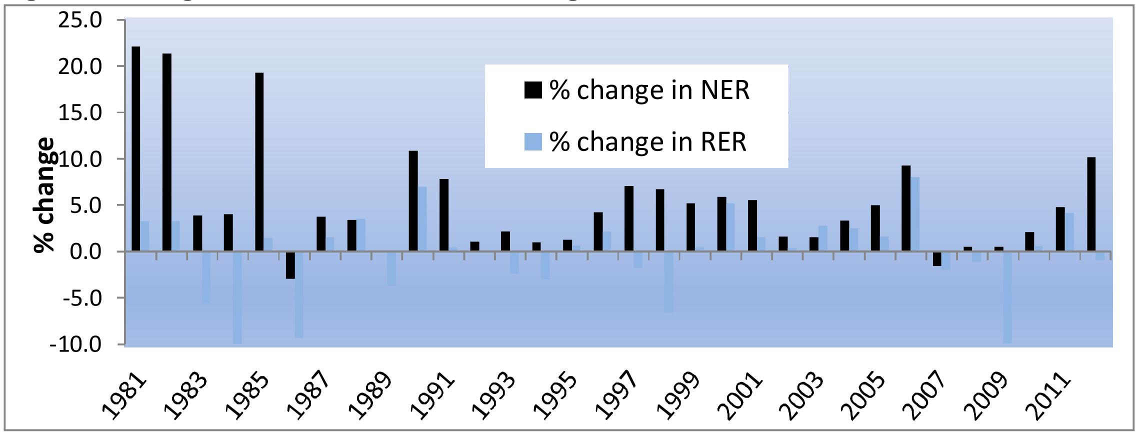
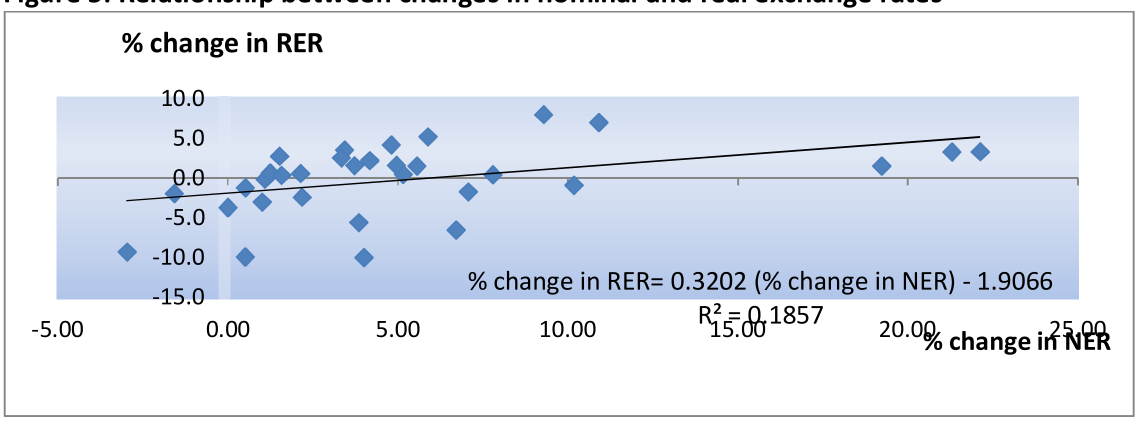


Figure 5: Relationship between changes in nominal and real exchange rates



V. Estimation Strategy and Results

The empirical exercise involves estimation using time series data. Modern econometrics attaches a lot of importance to the characteristics of these 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. Many macroeconomic time series are non-stationary in nature and consequently the ordinary least squares (OLS) regressions using these data might produce

[15] Note that this regressions does not suffer from the problems associated with the non-stationarity of the time series data, as the two first differenced variables are stationary. The estimated coefficient on % change in RER is also statistically significant at the five per cent level.

inconsistent, inefficient and often spurious estimates. In order to avoid such problems the integrating properties of the variables are now routinely examined by testing for the existence of unit roots in variables under consideration.

There are several methods for testing unit roots. Having applied different tests as part of this empirical exercise, it is found that one single test would be enough for summarizing the main statistical results and deciding about the outcomes. This test is known as the Augmented Dickey Fuller (ADF) test, which remains the most popular method for assessing time series properties. This 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 ADF test for the presence of a unit root.[16] However, 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.[17]

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

It is most 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 *~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 quite challenging. Not only that the results emanating from different unit root testing

[16] Note that the 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 ADF regressions can be applied without it. In the case of 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 other high frequency data.

[17] 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 ADF tests are rejected. Computed t-ratios and the corresponding critical values are compared on their absolute levels. These days many econometric software provide simulated critical values based on the model specifications, e.g. if the interest and/or trend term are included or not, and the number of observations.

procedures can be inconclusive but also that the tests like ADF often suffer from 'low power' (Engle and Granger, 1987), which 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. In the case of small sample Hall (1986), therefore, 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.[18] When the estimated autocorrlation coefficients at different lags are plotted against k, population correlogram is obtained.[19] 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. In this paper, we have also made use of autocorrelation coefficients and correlograms to consider the integrating properties of the variables.

Cointegration and Error Correction Modelling

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 procedures. The basic idea behind it 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 $\sim I(0)$.[20] This is what is known as the first step of Engle-Granger procedure.

[18] The autocorrelation coefficient like any ordinary correlation coefficient lies between –1 and +1.

[19] 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: $\frac{\sum (Y_t - \bar{Y})(Y_{t+k} - \bar{Y})}{\sum (Y_t - \bar{Y})^2}$

[20] One can employ the ADF test to test the stationarity of the residuals. However, in contrast to the regular ADF regressions, the test for residuals does not include an intercept term as the mean of the residual should

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 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:

$$\Delta Z_t = \pi_0 + \sum_{i=0}^{m} \pi_{1i} \Delta X_t + \sum_{i=1}^{n} \pi_{2i} \Delta Z_t + \pi_3 \hat{\upsilon}_{t-1} + \varepsilon_t \tag{5.2}$$

Equation (5.2) gives a very general representation of the ECM. Since all variables in (5.2) 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 $\hat{\upsilon}_{t-1}$, 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.

Even when the Engle-Granger first stage regression represents a valid long-run relationship, the standard errors obtained 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, one cannot be certain whether any particular

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be zero. Therefore, the OLS regression equation becomes: $\Delta \hat{v}_t = \rho \hat{v}_{t-1} + \kappa \Delta \hat{v}_{t-1} + \tau$. The null hypothesis for the test is that $\rho=1$ (non-cointegration, i.e. residuals are non-stationary) against the alternative of $\rho<1$ (cointegration or residuals are stationary). Like the regular ADF test statistics the estimated standard errors 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). Most econometric software now routinely provide these critical values.

explanatory variable is significant or not.[21] This problem may be overcome by using the Phillips-Hansen Fully Modified OLS (PHFMOLS) technique (Phillips and Hansen, 1990). This 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, correct for endogeneity in the regressors, and provides standard errors that follow standard normal distribution asymptotically and thus are valid for drawing inferences.

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. Johansen's cointegration test is used as a starting point in the vector autoregression (VAR) model. The vector autoregression model of order *p* (VAR (*p*)) is constructed as the following equation:

$$\Delta yt = \Phi 0 + \Sigma\, \Gamma i\, \Delta yt\text{-}i + \Pi yt\text{-}1 + \varepsilon t \tag{5.3}$$

where yt is a vector of variables in the model. $\Phi 0$ is the intercept vector and εt is a vector white noise process. The matrix of coefficients and information regarding the short-run relationships among the variables is denoted by Γi. On the other hand, the long-run information is contained in the matrix Π. If the rank of Π is r, where $r \leq n - 1$, then Π can be decomposed into two nxr matrices, α and β, such that $\Pi = \alpha\beta'$ where β is the matrix of cointegrating vectors; the elements of α are known as the adjustment parameters in the vector error correction model. The Johansen-Juselius procedure is based on the maximum likelihood estimation in a VAR model, and calculates two statistics – the trace statistic and the maximum Eigenvalue in order to test for the presence of r cointegrating vectors. Both these testing procedures consider the possibility of k-1 cointegrating vectors, where k is the number of endogenous variables in the specification.

[21] 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.

Apart from the advantages of finding multivariate vectors, the VAR system employed under the Johansen-Juselius procedure allows considering all variables as jointly determined (endogenous) and/or modeling certain variables as exogenous. However, there are important problems associated with this method. First of all, the results from the Johansen procedure can be very sensitive to the choice of lag-length. Although there are statistical tests for choosing the appropriate lag-lengths, in a small sample such tests may not be feasible. Use of lags involving all the the variable in the model could significantly reduce the degrees of freedom. Moreover, severe problem of collinearity among the regressors may also arise when a considerable size of VAR is used. Yet another problem is that multiple cointegrating vectors could indicate contradictory and often not very plausible economic interpretation and thus the use of economic theory may still be important despite starting with an atheoretic approach. These issues are given serious consideration as undertaken the actual econometric estimation of our model.

Results

Time Series Properties of Variables

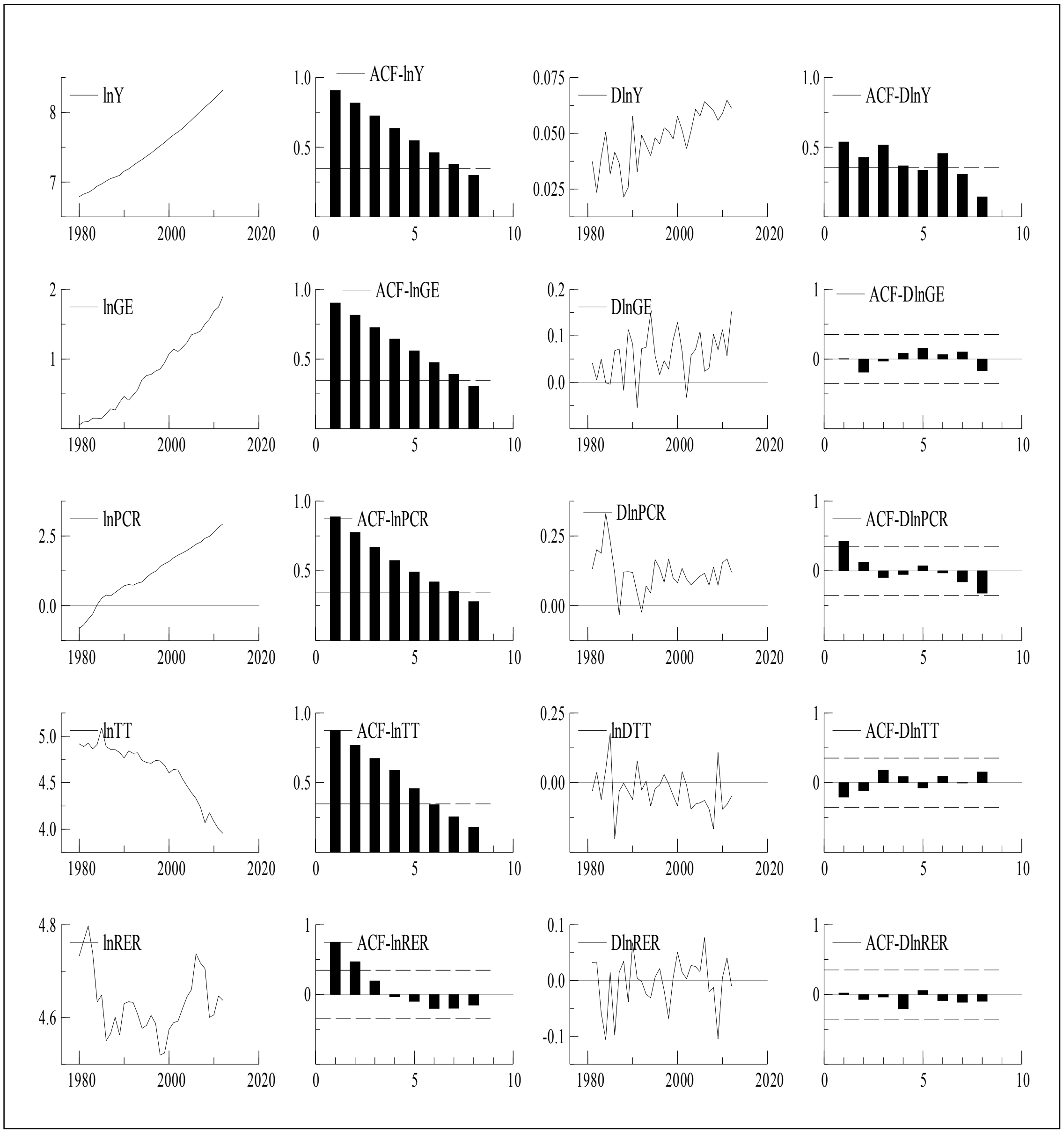
First we present the test results to determine whether the variables in the model can be represented as stationary or non-stationary processes. Table 1 provides ADF test statistics on level and first difference of the variables both with and without the trend term in the regressions, while Figure 6 presents the graphical plot of variables along with their correlograms.

Table 1: Unit Root tests of the variables

Variables	ADF test statistics		Variables	ADF test statistics	
	Without trend	With trend		Without trend	With trend
lnY	4.7402	-0.3431	ΔlnY	-2.0462	-4.7820
lnGE	1.8998	-2.0993	ΔlnGE	-4.3622	-5.3675
lnPCR	-1.0345	-3.9072	ΔlnPCR	-3.0858	-3.0018
lnTT	1.5084	-1.2030	ΔlnTT	-4.8318	-5.9685
lnRER	-2.7689	-2.5895	ΔlnRER	-4.0316	-4.3012

Note: The 95 per cent asymptotic critical value using the corresponding sample size and 1,000 replications for the level variables without trend is -2.9558 and for models with trend is -3.5815. For differenced variables, the comparable critical values are slightly different: -2.8916 and -3.5959 for models without and with the trend, respectively.

Figure 6: Graphical plots of variables and their correlograms



In the case of aggregate output (ln Y), the ADF regression with the trend term included returns a test statistic -0.343, which is smaller (absolutely) than the 95 per cent critical value of -3.5815. Consequently the null hypothesis of non-stationarity cannot be rejected. However, the ADF test without the trend term provides contradictory results as the computed test statistics of 4.74 turns out to be greater than the critical value. Since lnY is strongly trended (see Figure 6), the ADF test should be undertaken with the trend term included, in which case lnY comes out to be non-stationary on its level. ADF tests on ΔlnY also provide inconclusive results: the test with the trend term rejects the unit root hypothesis while the one without it cannot. Figure 6 seems to suggest ΔlnY to be trended and thus one needs to consider the unit root regressions with a trend. On this basis, ΔlnY is to be regarded as stationary. In other words, lnY is ~ I(1). The correlograms of lnY do not show any tendency of damping down while those on ΔlnY tail off on the first lag just like any stationary variable.

For lnGE, lnTT, and lnRER the ADF tests provide conclusive evidence irrespective of whether the trend is included or not. The computed test statistics on their levels do not reject the null hypothesis of unit roots, but on their corresponding first differences do. Therefore, lnGE, lnTT, and lnRER are to be regarded as I(1). Finally, lnPCR is strongly trended but the ADF test statistics with the trend term turn out to be higher than the critical values.

However, its correlogram strongly suggests its being non-stationary. On the other hand, ΔlnPCR does not appear to be trended, and the ADF test without a trend rejects the null hypothesis of stationarity. Therefore, we also conclude that lnPCR is I(1). With the consideration that all our variables possess unit roots on their levels but not on their first difference, we proceed to investigate the long-run relationship.

6.2. Estimating the Long-run Relationship

The estimating equation in (4.1) postulates a static long-run relationship among the variables. However, given the existence of unit roots in the level variables, a valid long-run relationship can only be found if the variables in the model cointegrate. Before testing the cointegration, Table 1 presents the Ordinary Least Squares (OLS) regression results. It shows that all variables are plausibly signed and the estimated t-ratios on the coefficients are high enough to be regarded as statistically significant using the classical inferential properties. However, in the presence of non-stationary variables, the estimated standard errors do not follow standard distribution and consequently one should not draw inferences. Similarly, although the model diagnostic tests with the null hypotheses of absence of serial correlation, no functional form problem, normality of residuals, and no heteroscedasticity problem cannot be rejected, these tests are not considered to be valid in when non-stationary variables are used in the regression.

The estimated regression in Table 2, known as the first step Engle-Granger procedure, would only represent a valid long-run relationship if there is evidence of cointegration. When residuals from the regression were tested for the existence of a unit root, the ADF test statistic turned out to be -5.2557 against the 95 per cent critical value of -4.8806. Therefore, the null hypothesis of nonstationarity (no cointegration) is rejected, providing the evidence for a valid long-run relationship. The graphical plot of the residuals and their correlogram (given in Appendix Figure A1) also seem to be supportive of this conclusion.

Table 2: OLS estimates of the static long-run relationship

lnY =	6.3056 +	0.4527 lnGE +	0.1774 lnPCR –	0.1102 lnTT +	0.2366 lnRER
(s.e.)	(0.5493)	(0.0499)	(0.0231)	(0.0494)	(0.0837)
t-ratio	11.4780	9.0678	7.6929	-2.2280	2.8248

Diagnostic Tests

Adjusted R^2 = 0.65

Serial Correlation: $\chi^2(1)$ = 2.9348	Functional Form: $\chi^2(1)$ = 1.5624
Normality: $\chi^2(2)$ = 1.3469	Heteroscedasticity: $\chi^2(1)$ = 0.0371

Given the evidence of cointegration, the long-run equation is estimated using the Phillips-Hansen Fully Modified OLS technique. This procedure makes a semi-parametric correction to OLS estimates, tackles the problems associated with endogeneity of regressors, and provides standard errors that follow standard normal distribution asymptotically, allowing for valid inferences to be drawn. Table 3 reports the relevant results.

Table 3: Philips-Hansen Fully Modified OLS estimates of the static long-run relationship

lnY =	5.993*** +	0.4622*** lnGE +	0.1833*** lnPCR -	0.0862** lnTT +	0.2771*** lnRER
(s.e.)	(0.4685)	(0.0452)	(0.0223)	(0.0419)	(0.0715)
t-ratio	12.791	102107	8.20	-2.0553	3.8752
Adjusted R^2 = 0.996					

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

It becomes clear that there is not much difference between the coefficient estimates presented in Tables 2 and 3. The PHFMOLS confirms the statistical significance of lnGE, lnPCR, and lnRER at the one percent level, and the same for the coefficient on lnTT at the five per cent level. While a 10 per cent increase in government expenditure is found to be associated with a 4.6 per cent increase in aggregate output, a comparable rise in credit to the private sector results in a 1.8 per cent expansion in GDP. The terms of trade elasticity is estimated at -0.086. Although a positive sign of this variable is generally expected, a negative coefficient is not implausible. Perhaps terms of trade shocks lead to more resource allocation in non-tradable and import competing sectors, more than compensating the loss of output resulting from subdued export activity. When a country's terms of trade experience sustained deterioration, as in the case of Bangladesh (see Figure 6 above), such an effect on aggregate output can be plausible.

The parameter of our interest in Table 2, *i.e.*, the coefficient on lnRER, is positvely signed and is statistically significant at the one per cent level. It indicates that as the RER index rises, (or nominal currency adjusts causing real devaluations), aggregate output rises. The estimated coefficient on lnRER suggests that a one per cent rise in RER index will result in a 0.28 per cent rise in GDP.

The OLS and PHFMOLS do not consider the possibility of multiple cointegrating vectors or more than one long-run relationship. This can be examined with the help of the Johansen-Juselius (JJ) cointegration procedure. One advantage of JJ approach is that it treats either all or a set of variables in the model as jointly determined. Macroeconomic time series described as dependent variables in theoretical specifications can exert influence on certain explanatory variables. If not tackled, such reverse feedback could cause biased and inefficient coefficient estimates in the model.

Although in theoretical and empirical specifications, the variation in the dependent variable (real GDP) is being explained by the variation in a number of other variables - regarded as explanatory variables - including the real exchange rate, it has been discussed in the literature that the growth in aggregate output can also lead to RER appreciation. This is popularly known as the Balassa-Samuelson effect. The underlying reasoning is based on the premise that during the development process tradable sectors experience faster productivity growth than their non-tradable counterparts. As tradable goods prices are set by international competition, an increase in productivity in this sector leads to an increase in wages, which is not detrimental to competitiveness. The mobility of labour would however imply that the wage increase in one sector spreads to the economy as a whole, raising the prices in the non-tradable goods sector without any commensurate productivity improvements. Given that the price index is an average of these two sectors, there is an increase in the prices of domestic goods relative to those from abroad. This results in an appreciation of the real exchange rate.[22] Similar reverse causality can also be argued for other variables. By undertaking a VAR framework, the JJ cointegration technique allows controlling for this jointly determined nature of relationship.

The JJ procedure starts with selecting a suitable order of VAR. The results can be quite sensitive to the chosen lag lengths and consequently a great caution is to be exercised. This can be problematic when the sample size is small. Ideally one can test for VAR orders using statistical methods, such as Akaike Information Criterion (AIC) and Schwarz Bayesian Criterion (SBC). However, in small samples, it is difficult to initially set for large VAR orders based on which testing down can be carried out. With 32 observations and 5 variables, we initially set the maximum lag length as 3. At this the AIC suggested a lag order of 2 as optimal, while SBC picked 1. When allowed a maximum order of 2, both AIC and SBC indicated the optimal lag length being. Given these mixed results, both VAR orders (1 and 2) were used to test for cointegration. In the JJ procedure, we consider lnY, lnGE, lnPCR and lnRER as jointly determined while lnTT as exogenous. This is in line with the modeling framework for a small economy for which prices of exports and imports are given for Bangladesh.

Table 4 reports the results of the JJ cointegration tests using both the Maximal Eigenvalue and Trace tests. Considering the Maximal Eigenvalue results for VAR order 1, it is found that the null hypothesis of no cointegrating vector (CV), i.e. no long-run relationship, is strongly rejected against the alternative of 1 cointegrating relationship as the computed statistics, 126.1, turns out to be greater than the 95% critical value of 31.48. However, the number of CVs less than or equal to 1 as against of 2 cannot be rejected. The null hypothesis of 2 or more CVs is also being rejected by the test. Therefore, the test favours just one cointegrating vector.

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[22] The Balassa-Samuelson effect is mainly studied for the countries that are growing faster than their trade partners or global economies. The idea is that in the catching up process, faster growing developing countries tend to face appreciated real exchange rate. There has been some estimate to suggest that for every one percentage point higher growth, faster growing countries would face RER appreciation by 0.4 per cent (Rogoff, 1996 and Frankel, 2006).

Turning to the Trace test (with order of VAR=1), the null hypothesis of no cointegrating vector is strongly rejected. Similarly, the null hypothesis of CV≤1 is also rejected against the alternative of at least 2 CVs. The computed test statistics in this case is 51.97 in comparison with the 95% critical value of 42.40. Furthermore, the test statistics also provide support for 3 CVs as the computed test statistic (29.01) with the null hypothesis of CV ≤2 (against an alternative of CV=3) turns out to be greater than the corresponding critical value (12.45). Therefore, on the whole there is strong evidence for cointegration amongst the variables in our model. But it is the number of CVs about which the two tests differ with the Maximal Eigenvalue suggesting only 1 and the trace test favouring 3 cointegrating vectors.

Table 4: Testing for cointegration in Johansen-Juselius procedure

Null hypothesis	Alternative hypothesis	Computed statistics: Order of VAR = 1	Computed statistics: Order of VAR = 2	95% Critical values
Maximal Eigenvalue test				
CV =0	CV =1	126.10	41.46	31.48
CV ≤1	CV =2	22.96	27.26	25.54
CV ≤2	CV =3	18.74	21.12	18.88
CV ≤3	CV =4	10.27	8.46	12.45
Trace test				
CV =0	CV =1	178.08	98.31	62.75
CV ≤1	CV =2	51.97	56.85	42.40
CV ≤2	CV =3	29.01	29.59	25.23
CV ≤3	CV =4	10.27	8.46	12.45

The fourth column in Table 4 reports the test results when VAR order 2 is considered. Now it is obvious that both the Maximal Eigenvalue and Trace tests support 3 cointegrating vectors. In light of the inconclusiveness of the test results regarding the number of cointegrating vectors, we allowed the maximum number of them. Table 5 therefore reports three long run relationships under each of VAR orders.

Table 5: Cointegrating vectors with VAR orders of 1 and 2

	VAR Order = 1			**VAR order of 2**		
	CV1	CV2	CV3	CV1	CV2	CV3
lnY	-1.00	-1.00	-1.00	-1.00	-1.00	-1.00
lnGE	0.60	0.87	0.51	0.44	0.44	0.92
lnPCR	0.18	1.61	0.17	0.42	0.21	-0.07
lnTT	-0.05	0.66	-0.03	-0.33	-0.04	-0.09
lnRER	0.20	2.91	0.41	1.24	0.37	0.42
Intercept	2.99	-9.85	5.08	-0.19	5.31	5.43

Note: lnY, lnGE, lnPCR, and lnRER have been used as jointly determined, while lnTT as exogenous variable. All CVs have been normalized on lnY. All the estimated coefficients are found to be statistically significant at the five per cent level.

When multiple cointegrating vectors are found, one needs to use either theory or other available information (or a combination of both), to select the most appropriate long-run relationship(s). First, taking the three CVs obtained with VAR order 1, we can perhaps eliminate CV2 as a possible long-run relationship in our case. This is because the effects of credit to private sectors and real exchange rate appear to be much larger than found elsewhere. The sign on the terms of trade variable also contradict the findings from other CVs. The remaining two CVs, 1 and 3, are comparable although the size of the coefficient on lnRER in CV3 appears to be more than double than that of CV1. Turning to the cointegrating vectors obtained with VAR order 2, CV3 can be discarded as a plausible long-run relationship as the coefficient on lnPCR is perversely signed. The real exchange rate effect in CV1 appears to be excessive, so is the size of the coefficient on TT. Hence, we are more inclined to consider CV2 as a possible long-run relationship.

When the 3 most plausible cointegrating relationships (CVs 1 and 3 under VAR order 1, and CV2 under VAR order 2) in Table 5 are compared, the RER elasticity varies between 0.20 and 0.42. They are comparable to OLS and PHFMOLS estimates of 0.23 and 0.27 respectively. All CVs under the JJ procedure produce a positive sign on the estimated coefficient of lnRER. The results therefore show irrespective of modeling techniques employed there are valid long-run relationships involving the variables in the equations. Although the JJ results are found to be sensitive to the choice of VAR orders, nevertheless the positive effect of RER is strongly maintained. If one has to be concerned about the application of a VAR model using a short sample like ours, the results provided by the Engle-Granger procedure and PHFMOLS can be used as a more reasonable estimate of the long-run relationship.

Short-run Dynamics

The estimated long-run relationships allow modelling the corresponding short-run dynamic adjustments using the error-correction mechanism. With a valid long-run relationship, there exists one corresponding error-correction model (ECM). Given the OLS, PHFMOLS, and three other long-run relationships, it is possible to obtain five short-run ECMs. Our estimates showed that they were not very different from one another and hence we report here just one ECM, using the long-run estimates derived from the first step Engel-Granger technique. These results are given in Table 6.

The error-correction model involves an OLS regression on the first difference of the variables incorporating the estimated long-run information (ecm_{t-1}), 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 as provided in Table 6.

Table 6: Short-run Error-Correction Model

$$\Delta \ln Y = 0.0187^{***} + 0.088^{**}\,\Delta \ln GE - 0.058^{**}\,\Delta \ln TT - 0.059^{*}\,\Delta \ln RER + 0.009^{**}\,D95SH + 0.33^{**}\,\Delta \ln Y_{t-1} - 0.252^{***}\,ecm_{t-1}$$

(s.e.)	(0.006)	(0.034)	(0.022)	(0.03)	(0.003)	(0.155)	(0.08)
t-ratio	2.983	2.309	-2.57	-1.80	-2.56	(2.17)	(-2.85)

Diagnostic Tests

Adjusted R^2 = 0.65

Serial Correlation: $\chi^2(1) = 0.08$	Functional Form: $\chi^2(1) = 0.81$
Normality: $\chi^2(2) = 1.88$	Heteroscedasticity: $\chi^2(1) = 2.38$

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.

In the short-run, credit to the private sector fails to register any significant effect and hence it has been dropped from the estimated equation. The magnitude of the short-run government expenditure elasticity is rather low – a one percent increase in government expenditure is associated with only 0.08 per cent increase in output. The effect however is highly significant at the one per cent level. The terms of trade effect is negative and the estimated short-run coefficient, which is significant at the five per cent level, is half the long-run estimate found in the Engle-Granger long-run equation. It is a standard practice to include the lagged dependent variable in the short run regression. The coefficient on $\Delta \ln Y_{t-1}$ turns out to be statistically significant and is retained in the parsimonious representation. Since the graphical plot of $\Delta \ln Y$ appears to be trended, a careful examination showed that incorporation of the shift dummy from 1995 helped explain significant variation in the data.

This implies that there has been a one-time intercept shift in the growth of output from 1995. Bangladesh initiated wide ranging macroeconomic reforms and trade liberalization programmes from the early 1990s, and it could be that significantly higher growth following those measures started to take place from 1995. The dummy D95SH turns out to be significant at the five per cent level.

Turning to our variable of interest, ΔlnRER, it is observed that in contrast to its long-run counterpart, the estimated short-run effect is negative. That is, the results suggest a contractionary impact of taka depreciation in the short-run. The effect is small - a 10 per cent real devaluation is associated with just above half a per cent decline in GDP – and significant only at the 10 per cent level. The ECM technique permits the possibility of finding differing short and long-run effects and in this case the negative coefficient on the real exchange rate is plausible. It could be that currency depreciations led to certain adjustments in terms of resource allocation and inflationary pressure (including expectation about price changes), resulting in declining economic activity in the short-run. However, with the improvement in the incentive structure for resource allocation in favour of the more productive tradable sector, and given its enhanced competitiveness, the expansionary effect sets in the long-run.

Finally, ecm_{t-1}, is correctly signed and highly significant at the one per cent level, indicating a valid representation of the error-correction model. This once again validates the existence of a long-run relationship as postulated in our model. The coefficient suggests that it takes just under four years to correct all short-run disequilibrium errors.

The explanatory power of the short-run model is quite reasonable as 65 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's (1980) test for heteroscedasticity and Jarque-Bera's (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. Therefore, all the model diagnostics performed satisfactorily.

Inflation and exchange rate movements

Although not a primary objective, given its relevance we decided to understand the impact of exchange rate movements on the domestic price level. No theoretical structure is being utilized to derive a full-blown empirical model for inflation. Rather, the objective is to shed some light on how the domestic price level is influenced by the exchange rate and global prices. The Johansen-Juselius method is used for studying this relationship.

The domestic price is represented by the consumers' price index (CPI), exchange rate is denoted by the nominal exchange rates of dollar in taka (R), and world price is proxied by the trade weighted partners' wholesale price indices (PT) as constructed earlier. Logarithmic transformation (ln) of these variables is undertaken so that the relationship involving lnPD, lnR and lnPT is examined.

As usual, the analysis begins with testing for unit roots in the variables. The ADF test results (as presented in Appendix Table A2) show that lnR and lnPT can be modelled as I(1) as the two variables were found to be non-stationary on their levels but their first difference transformations were stationary. However, the data generating process for lnPD turned out to be quite complicated: both the level and first differenced transformation (ΔlnDP) fail to reject non-stationarity. Only its second difference could reject the unit root hypothesis. Under normal circumstances, one should not expect a valid long-run relationship between one I(2) and two I(1) variables. Use of the GDP deflator, as a measure of domestic prices, did not make any difference as it too failed to be stationary on its first difference. One alternative is to use all variables on their stationary forms, but that would imply information on the long-run relationship being lost. Also, in a regression involving second difference of a variable along with first difference of others would make the interpretation of results less relevant for drawing policy implications. While the complicated data generating processes associated with Bangladeh's price variables are noted, it is also important to acknowledge the problem of low power of unit root testing procedures.

Despite not being able to confirm the same order of integration for lnP formally, we still decided to proceed with the JJ procedure. First, an unrestricted VAR model is specified to determine the optimal lag length. Initially, a generous VAR order of 3 to is specified to test for an optimal lag length. At this, the AIC suggests a lag length of 2, in contrast to just 1 depicted by the SBC. Given the ambiguity, both VAR orders of 2 and 1 are used.

For testing the long-run relationship, PT has been included as an exogenous variable, while R and PD are treated as jointly determined. With the VAR order of 2, the Maximal Eigenvalue test did not support for cointegration amongst the variable, but the Trace test indicated one cointegrating vector. On the other hand, with just one lag length, both the tests provided evidence for one long-run relationship (see Appendix Table A3). Given all this evidence, it would perhaps be sensible to consider one cointegrating relationship, which is estimated (having normalized on lnPD) as:

$$lnPD = -1.365 + 0.6246\ lnR + 0.762\ lnPT$$

where the estimated coefficients on lnPT and lnR turned out to be statistically significant at the one percent level, while the intercept was significant at the five per cent level. These results would indicate an exchange rate pass-through of 0.6, i.e. a 10 per cent nominal devaluation would result in a 6 per cent rise in the domestic price level. On the other hand, the international price elasticity (0.76) is also quite high. The corresponding short-run model (Table 7) appears to be a well-behaved one in the sense that the error correction term is correctly signed and highly significant. The speed on adjustment is rather low: only about 15 per cent of the disequilibrium errors are corrected within a year once a deviation from the long-run steady state relationship is caused. No model diagnostic problems are found. The error-correction model provides some justification for our proceeding with the long-run relationship despite challenged by the data generating process of lnP. According to the estimates presented in Table 7, in the short run nominal devaluation do not appear to have any significant effect on domestic prices. However, the world prices and lagged domestic prices exert significant positive influences.

Table 7: Short-run error-correction model for domestic prices

$\Delta lnPD =$	-0.0159^{***}	$+ 0.356^{**} \Delta lnPD_{t-1}$	$+ 0.1401^{**} \Delta lnPT$	$- 0.158^{***} ecm_{t-1}$	
(s.e.)	(0.0101)	(0.131))	(0.068)	(0.035)	
t-ratio	-1.56	2.72	2.07	(-4.23)	
Diagnostic Tests					
Adjusted $R^2 = 0.68$					
Serial Correlation: $\chi^2(1) = 0.19$			Functional Form: $\chi^2(1) = 0.17$		
Normality: $\chi^2(2) = 0.45$			Heteroscedasticity: $\chi^2(1) = 1.27$		

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.

When the long-run relationship is estimated using the PHFMOLS procedure, the coefficient on lnR comes out to be 0.998 (Appedix Table A4). This would imply a one-to-one correspondence between the movements in the nominal exchange rate and the domestic price level. Having imposed this restriction a priori, we construct lnP1, i.e., lnP1 = lnDP - lnR. The ADF tests unambiguously suggests lnP1 to be ~I(1) (Appendix Table A5). Finally, we now test if lnP1 and lnPT - both of them being ~ I(1), cointegrate. The tests are carried out with VAR orders of 3, 2, and 1. In each case, there is no evidence of a long-run relationship (Appendix Table A6). This suggest that there are certain domestic factors involved in determining domestic prices and as such changes in nominal exchange rate alone may not help protect and promote Bangladesh's competitiveness in the long-run.

VII. Summary of Findings and Policy Implications

How movements in the exchange rate affect overall economic activity has been a subject of longstanding controversy in macroeconomics and empirical scrutiny in applied policy analysis. In Bangladesh the news of taka depreciation is mostly greeted with skepticism about macroeconomic soundness of the economy, triggering public and policy debates. The issue is certainly controversial as exporters often demand for downward adjustments of the domestic currency in order to become more competitive in international markets in sharp contrast to protests by consumers and others who rely on imported goods for consumption and production and the increased taka value of dollar (or other foreign currencies for that matter) get translated into their higher prices. In recent times when the country has witnessed a general rise in prices, currency depreciation is widely regarded as a wrong policy choice contributing to the inflationary pressure. Bangladesh officially maintains a flexible exchange rate policy which would imply that in the face of a deteriorating external balance downward adjustment of taka becomes inevitable.

While the discussions on devaluations focus mainly on inflation, growth implications are rarely given adequate attention. In theory, currency depreciation is associated with possibilities for both contractionary and expansionary effects on outputs of different sectors. This would imply that for any country the net impact has to be determined empirically. Experiences of different countries and regions - as demonstrated in different studies - differ, further illustrating the need for undertaking country-specific empirical analysis.

This paper has made an attempt to examine the effects of exchange rate changes on Bangladesh's aggregate output, measured by GDP. Using a three-market Keynesian model, this paper has made use of an empirical specification that posits a long-run relationship between the real GDP and a vector of variables including terms of trade, government expenditure, credit to the private sector and the real exchange rate. For empirical investigation, this study has constructed the real exchange rate of Bangladesh taking into consideration of movements in prices and exchange rates of Bangladesh's most important trade partners.

The estimation strategy has duly examined the time series properties of the variables in regression analyses to avoid the problem of finding out spurious relationships and invalid inferences. In light of the non-stationarity of the model variables, cointegration techniques have been used to present and validate the long-run estimates. The short-run dynamics have also been captured applying the popular error-correction modeling techniques. The empirical analyses provide several major findings:

(i) Despite a steady rise in the nominal exchange rate, Bangladesh's real external competitiveness as measured by the real exchange rate has hardly changed since the 1980s. Nominal devaluations have largely accounted for cancelling out the differential inflation rates of Bangladesh vis-à-vis its trading partners. Indeed, a 10 per cent rise in the taka value of a dollar raises the RER (i.e. improvement in competitiveness) by only 3.2 per cent.

(ii) The movements in the real exchange rate do affect the overall output, and the effect is borne out even after controlling for other variables.

(iii) The long-run effects of real devaluations are found to be positive, i.e. downward adjustments of taka leading to the depreciation of the real exchange rate has an overall expansionary effect. The estimated real exchange rate elasticities lie in the range of 0.24 – 0.42 with our preferred estimates being 0.24 – 0.28. That is, a 10 per cent real depreciation of taka would lead to 2.4% to 2.8% increase in GDP.

(iv) However, in the short run, the impact of devaluations is likely to be contractionary. The effect is small: a 10 per cent real devaluation is associated with just above half a per cent decline in GDP.

(v) There is significant exchange rate pass-through to domestic prices with the long-run pass-through coefficients being estimated from 0.6 to as high as 1.0 (complete pass-through).

(vi) Despite the significant influence of world prices, as measured by trade partners weighted price indices for tradables, and the nominal exchange rate, local prices (measured by CPI) are likely to be affected by other domestic factors. As a result, movements in domestic prices differ from world prices and the nominal exchange rate adjustment has not been enough to compensate for this deviation.

Several important policy implications are to be derived from these findings. First of all, there are important growth implications of exchange rate management. Real devaluations turn out to be expansionary, which can operate through two possible channels. Firstly, by enhancing the external competitiveness, they can help expand output of the export sector. Although the country's export basket is dominated by textile and apparel items, which are known to be dependent on imported raw materials and capital goods, overtime their domestic value-added content has increased considerably thereby generating larger expansionary effects. Secondly, real exchange rate deprecations can also improve the competitiveness of the import-competing sector, supporting its growth.

Similarly, the appreciation of the real exchange rate will have adverse consequences for overall economic output. From that perspective taka depreciations have mitigated some of the adverse consequences. The analysis presented in this paper shows Bangladesh's domestic prices relative to those of tradable sector of its major trading partners have risen substantially. Nominal devaluations have just helped maintain Bangladesh's international competitiveness at more of less the same level since the 1980s.

The finding of the expansionary effect of real depreciations points towards a major policy issue concerning the role of the exchange rate as a development strategy. This has been a subject matter of recent debates and discussions in the literature following empirical results presented by Rodrik (2008), but also upheld in subsequent empirical exercises, showing that the undervaluation of the currency stimulates economic growth. It has been argued that since bad institutions tax tradables more heavily, and market failures are more prevalent in the tradable sector, devaluations alleviate the consequences of these distortions. Given that the findings in the present paper can be interpreted as consistent with Rodrik (2008), one relevant question is what policy lesson can be drawn for Bangladesh?

It needs to be pointed out that the ability to maintain a depreciated real exchange rate regime is going to be far more challenging than staging nominal devaluations. As there is the evidence of strong and significant effect of nominal devaluations on domestic prices, downward adjustments of the currency will result in mounting inflationary pressure in the long run. On the other hand, in the short-run, perhaps due to adjustments taking place in terms of resource allocation and certain price shocks, currency depreciation can be contractionary. Therefore, while one episode of nominal devaluations can help promote competitiveness and expand output over the long-run, rising price levels will erode at least some significant portion of these gains. It has therefore been argued that countries seeking to use a competitive exchange rate to aid growth need to develop an exit strategy to avoid getting locked into a strategy that has outlived its usefulness (Eichengreen, 2008).

Furthermore, it has also been pointed out that nominal devaluations may not always translate into real exchange rate depreciations. Fiscal and monetary policies, in particular, must be consistent with the exchange rate regime to tackle the problem of rising prices. That is why using the real exchange rate as a policy variable may be a very difficult option. In the long-run the movements in the real exchange rate is to be determined by fundamentals that influence the resource allocations between tradable and non-tradable sectors. Policies that can help more spending on tradables and/or avoid generating undue pressure on the prices of non-traded goods may prove to be more effective to promote international competitiveness.

This paper has highlighted the changing composition of Bangladesh's trade partners. Since the late 1990s, a set of fast-growing and emerging developing countries have gained more prominence, pushing down the relative significance of advanced and developed economies. Prices in developing countries are generally higher and faster growing countries are likely to have appreciated real exchange rates, as pointed out earlier. Therefore, the shifts in the significance of trading partners might provide opportunities for Bangladesh maintaining a competitive real exchange rate regime, provided that domestic inflationary pressure is contained.

Even after adjusted for nominal devaluations, Bangladesh's domestic prices seem not to have a long-run relationship with trading partners' prices. This indicates there are certain internal factors that cause different movements in domestic prices. In this context, it is important to better understand the allocation of resources in non-traded and import-competing sectors and their implications for productivity and output growth, and prices. In the face of a sustained deterioration in terms of trade, the estimated positive effect on output could imply resources being allocated to import-competing sectors or high value added services sector. In the aftermath of financial crisis, there has been a lot of discussion on whether developing countries can enhance their reliance on their domestic sectors to promote economic growth. Bangladesh has a relatively large domestic economy, compared to many other countries with small population, and as such domestic demand management will always have an important influence on overall economic growth. While the need for maintaining competitiveness cannot be overemphasized a better understanding of inter-sectoral resource allocation and productivity growth will make informed exchange rate management that would promote competitiveness of exports as well as expansion of import-competing sectors.

In fine, the management of the exchange rate would remain a delicate task for policymakers. While the effects of nominal devaluations are likely to have positive effects on output growth, tackling inflation would be critically important. In the long-run Bangladesh needs to maintain its international competiveness for which nominal exchange rate adjustment will be an important instrument. Its effectiveness will however be depend on the nature of accompanying monetary and fiscal policies and other factors influencing the relative prices of tradables. There is evidence that despite embarking on an official policy of freely-floating system, the management of exchange rate has made it work like a fixed rate regime. This has resulted in occasional large nominal devaluations, particularly when the disequilibrium in the external balance becomes unsustainable. A more pragmatic approach would be to allow creeping depreciations that would avoid any significant contractionary effect in the short run while allowing for improved competitiveness and output growth in the long-run.

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Appendix

Appendix Table A1: Bangladesh's important trade partners 2000-2011 (% of trade share)

	2001	2003	2005	2007	2008	2009	2010	2011	Avg 2007-2011
World	100	100	100	100	100	100	100	100	100
China	5.27	6.80	8.58	9.17	9.63	10.02	11.48	12.29	10.52
India	8.52	9.55	9.26	9.24	10.19	8.34	9.85	9.64	9.45
United States	13.30	10.60	10.43	10.91	8.68	9.34	8.91	8.25	9.22
Germany	5.30	6.31	6.53	6.62	6.03	6.25	6.14	7.01	6.41
United Kingdom	4.91	4.65	5.00	4.56	4.03	4.35	3.77	4.09	4.16
Singapore	5.85	6.64	4.05	3.81	4.66	4.69	3.67	2.69	3.90
Kuwait	0.21	1.09	5.30	4.96	4.82	1.05	2.05	2.41	3.06
Japan	5.31	3.81	2.84	2.46	2.83	3.11	3.13	3.08	2.92
France	2.62	2.78	2.69	2.56	2.63	2.92	2.58	2.76	2.69
Hong Kong, China	3.82	3.23	3.01	2.83	2.67	2.48	2.43	1.73	2.43
Malaysia	1.03	1.35	1.39	1.12	1.73	2.79	3.23	3.23	2.42
Korea, Rep	2.83	2.36	2.10	2.16	2.25	2.39	2.51	2.60	2.38
Canada	1.18	1.60	1.51	2.14	1.91	2.52	2.50	2.31	2.27
Netherlands	2.30	1.87	1.36	1.78	2.09	2.65	2.44	1.73	2.14
Italy	2.31	2.07	2.16	2.19	1.99	2.04	1.89	2.18	2.06
Thailand	1.33	1.15	1.48	1.46	1.41	1.71	1.82	2.10	1.70
Indonesia	1.25	1.34	1.31	1.28	1.38	1.67	1.87	1.93	1.63
Australia	1.55	1.34	1.33	1.17	1.14	1.84	1.96	1.59	1.54
Spain	0.74	1.18	1.51	1.60	1.50	1.52	1.30	1.67	1.52
UAE	1.11	1.30	1.68	1.19	1.80	1.20	1.53	1.60	1.47
Saudi Arabia	1.69	1.30	1.43	1.21	1.58	1.26	1.58	1.42	1.41
Belgium	1.40	1.55	1.31	1.86	1.39	1.17	1.05	1.33	1.36
Turkey	0.23	0.40	0.62	0.79	0.86	1.08	1.45	1.33	1.10
Pakistan	0.75	0.85	0.88	0.87	1.00	1.00	1.15	1.39	1.08
Brazil	0.30	0.31	0.43	0.73	0.53	1.22	1.12	1.79	1.08

Appendix Table A2: ADF test results for PD, PT and R

Variables	ADF test statistics		Variables	ADF test statistics	
	Without trend	With trend		Without trend	With trend
lnPD	0.6099	-1.896	ΔlnPD	-2.9123	-2.4980
lnPT	0.1617	-2.3514	ΔlnPT	-4.009	-4.2533
lnR	-1.2365	-2.1536	ΔlnR	-4.2312	-4.2533

Note: The 95 per cent asymptotic critical values using the corresponding sample size and 1,000 replications for the level variables without trend is -2.9558 and for models with trend is -3.5815. For differenced variables, the same 95 per cent critical values are slightly different: -2.8916 for models without trend and -3.5959.

Appendix Table A3: Maximal Eigenvalue and Trace tests for cointegraton in the price model

Null hypothesis	Alternative hypothesis	Computed statistics: Order of VAR = 1	Computed statistics: Order of VAR = 2	95% Critical values
Maximal Eigenvalue test				
CV =0	CV =1	91.7351	16.5989	18.8800
CV ≤1	CV =2	7.2194	10.9845	12.4500
Trace test				
CV =0	CV =1	98.9545	27.5834	25.2300
CV ≤1	CV =2	7.2194	10.9845	12.4500

Appendix Table A4: PHFMOLS estimates of the price model

$$\begin{array}{lllll} \ln PD & = & -1.1428^{***} & + \; 0.9987^{***} \ln R & + \; 0.4414^{***} \ln PT \\ \text{(s.e.)} & & (0.3882) & (0.0622) & (0.1271) \\ \text{t-ratio} & & -2.9437 & 16.0344 & 3.471 \end{array}$$

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

Appendix Table A5: Unit root testing for lnP1

Variables	ADF test statistics		Variables	ADF test statistics	
	Without trend	With trend		Without trend	With trend
lnP1	-1.3710	-2.9566	∆lnP1	-5.2334	-4.9745

Note: The 95 per cent asymptotic critical values using the corresponding sample size and 1,000 replications for the level variables without trend is -2.9558 and for models with trend is -3.5815. For differenced variables, the same 95 per cent critical values are slightly different: -2.8916 for models without trend and -3.5959.

Appendix Table A6: Cointegration test for lnP1 and lnPT

Null hypothesis	Alternative hypothesis	Computed statistics: Order of VAR = 1	Computed statistics: Order of VAR = 2	95% Critical values
Maximal Eigenvalue test				
CV =0	CV =1	126.10	41.46	31.48
CV ≤1	CV =2	22.96	27.26	25.54
CV ≤2	CV =3	18.74	21.12	18.88
CV ≤3	CV =4	10.27	8.46	12.45
Trace test				
CV =0	CV =1	178.08	98.31	62.75
CV ≤1	CV =2	51.97	56.85	42.40
CV ≤2	CV =3	29.01	29.59	25.23
CV ≤3	CV =4	10.27	8.46	12.45

Figure A1: The cointegrated long-run relationship involving lnY, lnGE, lnPCR, lnTT and lnPER and its partial autocorrelation function

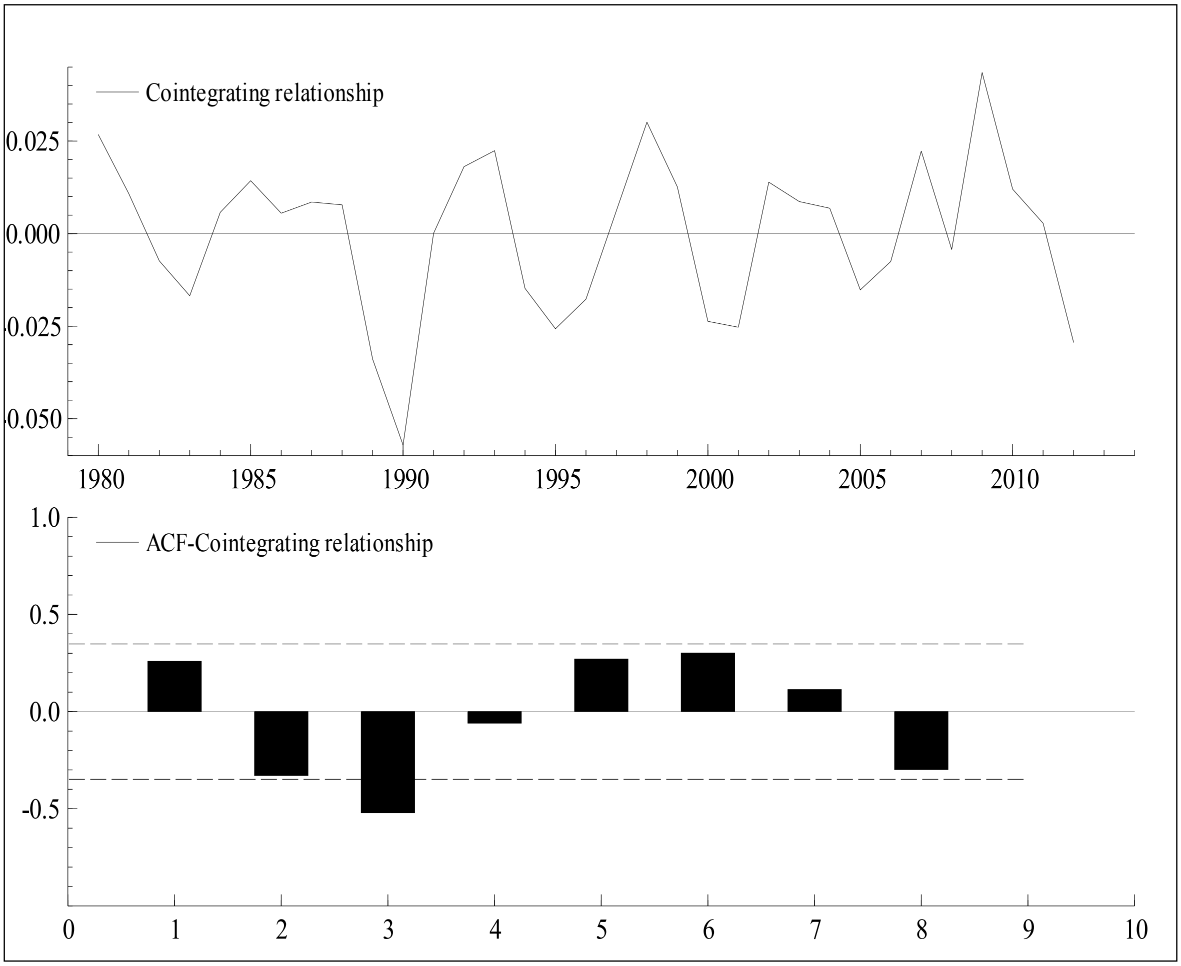
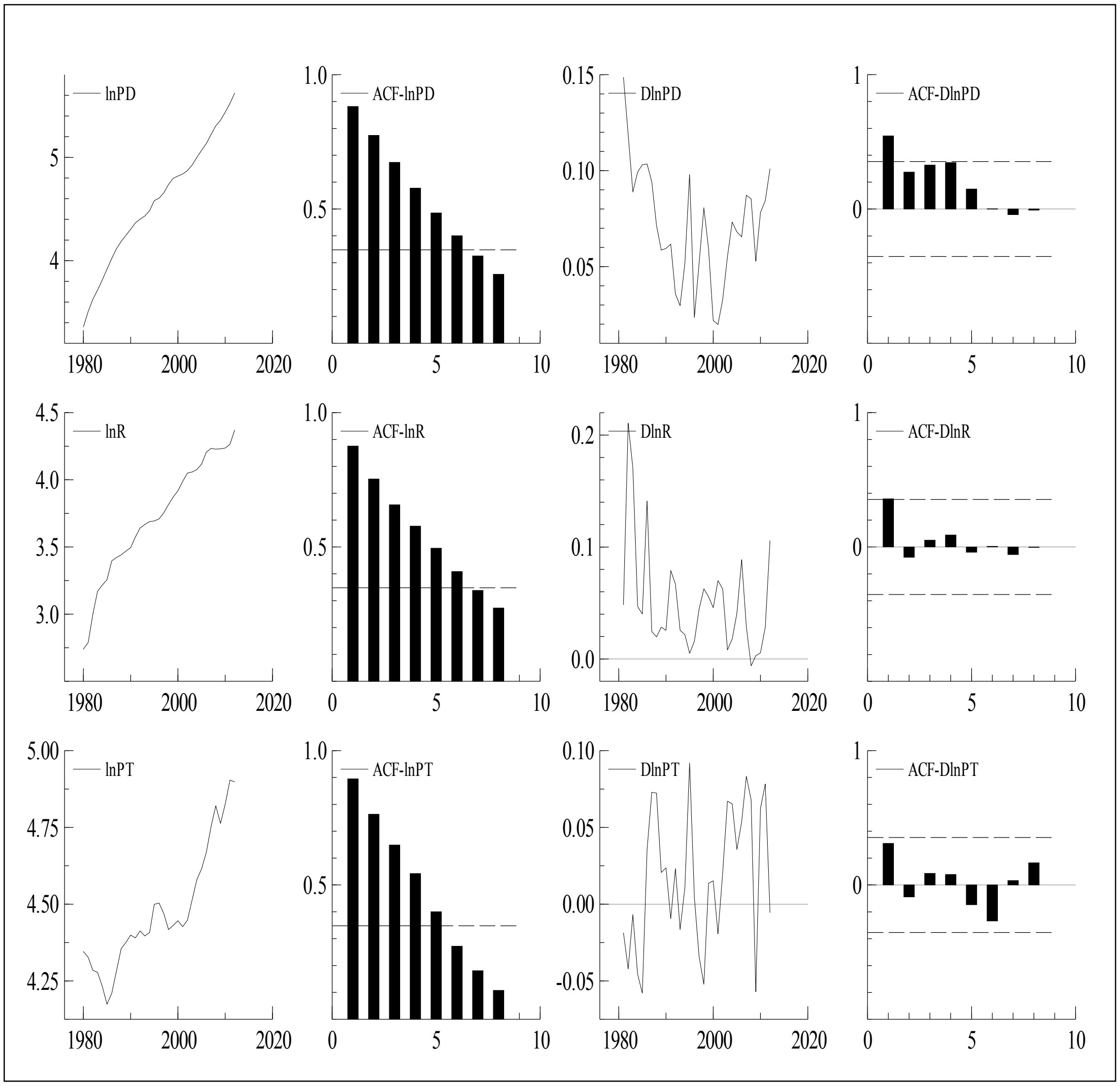


Figure A2: Graphical plot of price variables and exchange rate their partial autocorrelation functions



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