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Exchange Rate Responses to Inflation in Bangladesh

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Abstract

This paper investigates the exchange rate responses to inflation in Bangladesh during the period from 1972–73 to 1999. Both annual and monthly data are used in the investigation. The results suggest that past consumer price inflation generally led to currency devaluation, measured as a decline in the value of the currency in terms of the trade-weighted nominal effective exchange rate. The effect of inflation on devaluation, however, became weaker following the financial reforms undertaken in the early 1980s. The effect of devaluation on inflation was not significant, and this result remained robust throughout the sample period.

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I. Introduction

As an instrument of macroeconomic management, currency devaluation remains controversial, especially in developing countries. The perception is that devaluation causes inflation, if not stagflation (Cooper, 1971; Kamin, 1988; Krugman and Taylor, 1978). This perception is not innocuous but influences macroeconomic policymaking in general and exchange rate policy in particular. For example, structural adjustment programs supported by the IMF and the World Bank are subjected to persistent criticism, irrespective of whether these programs promote macroeconomic stability and economic growth.[2]

The issues relating to the possibility of a causal linkage between devaluation and inflation are debated at both the theoretical and the empirical level. At the \theoretical level, two broad but opposite strands of thought can be traced to the debate over structuralism versus monetarism in Latin America (Corbo, 1974). Emphasizing the presence of myriad bottlenecks in developing economies, the structuralists identify various channels through which devaluation, often considered an exogenous policy instrument, causes inflation or stagflation (Agénor and Montiel, 1999; Sachs and Larrain, 1993; Taylor, 1979, 1983). For example, it is argued that devaluation raises the price of tradables directly,[3] and, depending on conditions in the labor markets and on institutional arrangements such as wage indexation, may raise the price of nontradables as well. Even if the price of nontradables does not rise in response to wage increases but remains sticky downward, devaluation may raise the general price level or its growth rate.

In contrast, the monetarist literature in general does not consider devaluation a primary source of inflation. The classic monetarist formulation is that "inflation is always and everywhere a monetary phenomenon,"[4] implying that an excess money supply is the primary source of inflation. Excess money, in turn, may originate on the supply or the demand side of money markets, or both.[5]

[2] Devaluation may help to bring about macroeconomic stability when a country experiences external imbalances caused by, for example, expansionary monetary and fiscal policies. It is accepted that macroeconomic stability is necessary, if not sufficient, for sustained economic growth. However, because devaluation is believed to have effects on inflation and income distribution, these effects complicate the design of stabilization policies. See Aghevli, Khan, and Montiel (1991), Agénor and Montiel (1999), and Edwards (1989a,b) for detailed discussion of the economic effects of devaluation in developing countries.

[3] This follows from the purchasing power parity proposition. At the microeconomic level, the literature on the exchange rate pass-through deals with the effect of devaluation on the price of tradables. The exchange rate pass-through, which is usually less than unity for manufactured products, differs a great deal across industries, depending on industrial organization factors including market concentration, product substitutability, and the market shares of domestic and foreign firms (Dornbusch, 1987; Feenstra, 1995; Kenen and Pack, 1980; Krugman, 1987).

[4] This statement is attributed to Milton Friedman (Mishkin, 1998, p. 659).

[5] Ceteris paribus, the expectation of a devaluation can increase an excess of money supply and thereby inflation by lowering the demand for domestic money. In a market economy operating under a flexible exchange rate system, this effect is captured by the nominal interest rate via the uncovered interest rate

Devaluation is then considered not an exogenous policy instrument but rather a passive response of the monetary authorities to, say, inflation differentials between the home country and foreign countries. In the early literature, such as Cassel (1918), this idea was formalized in the relative purchasing power parity proposition. The recent literature is somewhat eclectic and suggests a feedback relationship between devaluation and inflation. However, the nature of this relationship depends on the time lag, as well as the inflation history and exchange rate arrangements of the country concerned. Nonetheless, the core idea remains that devaluation is at most a propagator, rather than an initiator, of inflation (Agenor and Montiel, 1999).

The debate on the linkage between devaluation and inflation has some relevance to policymaking in developing countries, including Bangladesh. For example, Bangladesh was under various IMF- and World Bank-supported structural adjustment programs throughout the 1980s and 1990s. Devaluation was and remains a key component of Bangladesh's macroeconomic policy package, which aims at achieving and maintaining macroeconomic stability in general and external stability, meaning sustainable current account deficits, in particular. In fact, following the decline of annual inflation from about 11 percent in the 1980s to 5 percent in the 1990s, external stability has gained priority as an objective of macroeconomic policies. This is reflected in the relatively greater exchange rate flexibility observed in the 1990s under an adjustable pegged exchange rate system.[6] In a broader context, considerable foreign capital inflows to Bangladesh have created a policy impetus to prevent the currency from appreciating in real terms (Hossain, 1996, 2000a).

Although the rise in exchange rate flexibility is generally viewed as a sensible strategy for achieving and maintaining external stability, critics hold that devaluation remains inflationary and self-defeating as far as improving the country's external balance position is concerned. Although this may not be the case in reality, policymakers remain on the defensive and find it difficult to justify devaluation even when it could be the response of monetary authorities to past inflation that might have caused the real exchange rate to appreciate and created unsustainable external imbalances.

An investigation of the linkage between inflation and devaluation is therefore timely, especially as this issue has yet to be investigated for Bangladesh. Using both annual and monthly data for the

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parity proposition. However, if the domestic nominal interest rate is determined administratively or remains unresponsive to any expected change in the exchange rate, the exchange rate may affect money demand directly (provided that asset holders can substitute alternative assets, including foreign currency) and thereby affect inflation. Under a fixed or a pegged exchange rate system, changes in exchange rates may also affect the money supply through developments in the balance of payments. This paper, given its limited scope, does not explore such possibilities of devaluation affecting inflation through various monetary channels.

[6] This is reflected in the coefficient of variation of the change in the nominal effective exchange rate, which, in the case of monthly data, increased from 2.7 during the 1980s to 17.7 during the 1990s. For details see Appendix I.

period from 1972–73 to 1999,[7] this paper investigates the monetary authorities' exchange rate responses to inflation, given that Bangladesh has been operating under a pegged exchange rate system since it gained independence in 1971. Because the paper applies the error-correction modeling approach, it also allows an investigation of the feedback effect of devaluation on inflation.[8]

The rest of the paper is structured as follows. Section II develops a model of inflation within the modeling approach based on the classification of goods and services as tradable or nontradable and explains why devaluation may not necessarily be an independent source of inflation. Section III reviews exchange rate policy regimes in Bangladesh since the early 1970s with a view to establishing any plausible linkage between the exchange rate regime and inflation. Section IV estimates cointegral and error-correction models of inflation and conducts a Granger causality test between devaluation and inflation. Section V offers concluding remarks. Appendix I describes the construction of the trade-weighted nominal effective exchange rate for the sample period of this study. Appendix II discusses data issues, such as variable definition, data sources, and the time-series properties of variables in the regression analysis.

II. A Model of Inflation

This section develops a model of inflation within the tradable-nontradable modeling framework and establishes a linkage between devaluation and inflation.

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[7] I received conflicting suggestions from reviewers of this paper on the use of both annual and monthly data. Although the use of only monthly data would have been economical, one reviewer suggested that I use both annual and monthly data to find out whether the results are robust to both. This has become somewhat necessary as, in the case of monthly data, I use industrial output as a proxy for real GDP. Also, because there is no clear-cut choice between the narrow and the broad definition of money, I have used both definitions in my investigation.

[8] In general, the causal linkage between inflation and devaluation can be examined by applying Granger causality tests within an error correction modeling framework, provided that both these series are stochastic in nature or generated through market mechanisms within a general-equilibrium framework. Apparently this cannot be said for Bangladesh, whose exchange rate has been pegged to a single currency or a basket of foreign currencies since the early 1970s but where the authorities have adjusted the official exchange rate at different intervals under different exchange rate regimes. Nevertheless, the Granger causality framework remains valid in the present context. Note that although the exchange rate is probably set through an implicit policy rule, it is a function of random variables such as the inflation rate and hence can be considered a random variable. In any case, even if the exchange rate only moves in discrete steps, what is needed for the approach to remain valid is that the coefficient of the time trend in the equations for changes in the exchange rate not be zero. I owe these classifications to Professor Adrian Pagan.

A. The Model

The domestic price level

Assume that goods transacted in a small, open economy can be divided into tradables and nontradables. Let the domestic price of transacted goods (P_t) be defined as the geometric average of the prices of tradables (PT_t) and nontradables (PNT_t). The price level can then be specified in the following natural logarithmic form:

$$\ln P_t = \phi \ln PT_t + (1 - \phi) \ln PNT_t, \qquad (2.1)$$

where ϕ is the share of tradables in total expenditure. For simplicity, ϕ is assumed constant.[9]

The price of tradables

For a small economy the price of tradables in foreign currency is determined in the international market. Following the purchasing power parity proposition, the price of tradables in domestic currency can then be expressed as

$$\ln PT_t = \ln ER_t + \ln PT^f_t, \qquad (2.2)$$

where ER is the exchange rate (defined as units of domestic currency for each unit of foreign currency) and PT^f is the price of tradables in foreign currency.

Equation (2.2) suggests that the price of tradables in domestic currency may change in response to a change either in the exchange rate or in the price of tradables in foreign currency or both. When the exchange rate remains fixed, the price of tradables in domestic currency changes with the change in the price of tradables in foreign currency. However, under a flexible exchange rate system, the price of tradables in domestic currency may change when the exchange rate changes with or without a change in the price of tradables in foreign currency.

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[9] This is a restrictive assumption. In general, openness and economic growth have different effects on the share of tradables in total expenditure. For example, the share of tradables in total expenditure rises as the economy opens, provided that opening lowers trade restrictions and transactions costs, both implicit and explicit. However, a rise in income per capita, with or without an increase in openness, increases the relative demand for nontradables and thereby causes a structural transformation of the economy in favor of nontradables. Therefore, analytically, ϕ can be expressed as a function of the real exchange rate, which, in turn, depends on a set of real and nominal factors, including economic growth, capital flows, the terms of trade, and the stance of monetary and fiscal policies (Edwards, 1989a; Hossain, 2000a).

For simplicity, PT^f is normalized to unity. Equation (2.2) can then be written as

$$\ln PT_t = \ln ER_t. \tag{2.2'}$$

This shows a one-to-one relationship between the price of tradables and the exchange rate. If the exchange rate is considered an exogenous policy instrument, ceteris paribus, devaluation may lead to a rise in the domestic price level (Prachowny, 1975). However, the exchange rate is not a purely exogenous policy instrument. It can be largely endogenous in the sense of being either determined by market forces under a flexible exchange rate system, or changed by the monetary authorities through a reaction function in response to pressure built up on the exchange rate under a fixed or an adjustable pegged exchange rate system. In general, it is suggested that a feedback relationship exists between the price level and the exchange rate. Such a relationship usually becomes pronounced during periods of high inflation or hyperinflation (Dornbusch, 1988).

The price of nontradables

Assume that the price of nontradables changes in response to disequilibrium in the money market. Within a flow disequilibrium in the money market, the price of nontradables may then change in response to a discrepancy between the log difference of actual real balances at the beginning of the period ($\Delta\ln m_{t-1}$) and the log difference of real balances that individuals desire to hold at the end of the period ($\Delta\ln m^d{}_t$), such that

$$\ln PNT_t - \ln PNT_{t-1} = \gamma\,(\Delta\ln m_{t-1} - \Delta\ln m^d{}_t) + u_t, \tag{2.3}$$

where $m = M/P$ (M being the nominal money stock and P the price level); γ is the coefficient of adjustment, whose value is expected to lie between zero and unity; and u is a random error term with zero mean and a constant variance. Equation (2.3) shows that only a proportion (γ) of disequilibrium in the money market is eliminated between periods $t - 1$ and t.

Inflation

Take the first-order logarithmic differences of equations (2.1) and (2.2'), such that

$$\ln (P_t/P_{t-1}) = \phi \ln (PT_t/PT_{t-1}) + (1 - \phi) \ln (PNT_t/PNT_{t-1}) \tag{2.4}$$
$$\ln (PT_t/PT_{t-1}) = \ln (ER_t/ER_{t-1}). \tag{2.5}$$

Substitution of equations (2.3) and (2.5) into equation (2.4), following rearrangement of terms, yields

$$\ln (P_t/P_{t-1}) = \phi \ln (ER_t/ER_{t-1}) + (1 - \phi)\gamma\,(\Delta\ln m_{t-1} - \Delta\ln m^d{}_t) + (1 - \phi)\, u_t. \tag{2.6}$$

Because the term representing flow disequilibrium in the money market cannot be measured directly, one way of simplifying this equation is to specify an equation for changes in desired real balances,[10] such that

$$\Delta \ln m^d_t = \beta 0 + \beta 1\ \Delta \ln y_t, \tag{2.7}$$

where y is real income and the βs are structural parameters. The intercept term $\beta 0$ captures any trend element in the level of money demand.

Substitution of equation (2.7) into equation (2.6), after manipulations, yields

$$\begin{aligned}\ln (P_t/P_{t\text{-}1}) = \quad & - (1 - \phi)\,\gamma\beta 0 + \phi \ln (ER_t/ER_{t\text{-}1}) + (1 - \phi)\gamma\ \Delta\ln M_{t\text{-}1} - (1 - \phi)\gamma\ \Delta \ln P_{t\text{-}1} \\ & - (1\text{-}\phi)\gamma\beta 1\ \Delta\ln y_t + (1 - \phi)\, u_t. \end{aligned} \tag{2.8}$$

This is an estimable model of inflation for a developing country like Bangladesh, which operates an adjustable pegged exchange rate system (Hossain and Chowdhury, 1996). It shows that inflation depends on the rate of devaluation, the rate of economic growth, and the rate of growth of the real money stock with a one-period lag.[11] The variable of interest is the rate of devaluation. Its impact on the inflation rate is given by ϕ, which represents the average share of tradables in total expenditure.[12]

B. Exchange Rate Reponses to Inflation

The inflation equation (2.8) is in quasi-reduced form because devaluation, although it appears in the right-hand side of the equation, cannot be considered an exogenous policy instrument. As indicated earlier, under an adjustable pegged exchange rate system, devaluation is essentially a response of the monetary authorities to high inflation in the home country relative to inflation in its trading partners. In the event that the monetary authorities follow a real target approach to exchange rate policy (Corden, 1991, 1993), they are likely to adjust the nominal exchange rate to

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[10] For simplicity, the rate of expected inflation or the nominal interest rate (a proxy for the opportunity cost of holding money) is assumed constant.

[11] In the simplified equation (2.8), the growth of real money is expressed as the differential between the growth of the money stock and the rate of inflation.

[12] If the exchange rate remains fixed under a fixed exchange rate system or does not change much under a pegged exchange rate arrangement, the coefficient of $\ln (ER_t/ER_{t\text{-}1})$ would not be different from zero. Imposing a zero restriction, equation (2.8) can then be written in the following form:

$$\ln (P_t/P_{t\text{-}1}) = - \gamma\beta 0 + \gamma\ \Delta\ln M_{t\text{-}1} - \gamma\ \Delta \ln P_{t\text{-}1} - \gamma\beta 1\ \Delta\ln y_t\ + u_t. \tag{2.8'}$$

This is a variant of the conventional monetary model of inflation, such as Harberger's (1963) model of inflation for Chile. This restricted model can be derived from the flow equilibrium condition of the money market, such that $\Delta\ln (M/P) = \Delta\ln m^d(y)$.

inflation differentials between the home country and its trading partners with a view to preventing the real exchange rate from appreciating. Accordingly, a policy reaction function that is compatible with the relative purchasing power parity proposition can be specified as follows:

$$\ln (ER_t/ER_{t-1}) = \delta\,(\Delta \ln P_{t-1} - \pi^e_f), \tag{2.9}$$

where π^e_f is foreign inflation (assumed constant) and δ is the reaction coefficient, whose value lies between zero and one. It suggests that devaluation in the current period is in response to an inflation differential between the home country and its trading partners in the last period.

Substitution of equation (2.9) into equation (2.8) yields

$$\begin{aligned}\ln (P_t/P_{t-1}) = \; & -[(1-\phi)\gamma\beta0 + \delta\pi^e_f] + (1-\phi)\gamma\,\Delta\ln M_{t-1} + [\phi\delta - (1-\phi)\gamma]\,\Delta\ln P_{t-1} \\ & -(1-\phi)\gamma\beta1\,\Delta\ln y_t + (1-\phi)\,u_t.\end{aligned} \tag{2.10}$$

This is a restricted model of inflation that has eliminated devaluation from the specification. However, it is consistent with the classical-monetarist tradition, which considers devaluation more an effect than a primary source of inflation. Whether this restricted model is a valid representation of the inflationary process in Bangladesh can be considered a testable hypothesis. This hypothesis could be accepted if devaluation is found superfluous in a monetary model of inflation.

After a brief review of the history of exchange rate regimes in Bangladesh, the rest of the paper proceeds with specifying and estimating a monetary model of inflation within the cointegration and error correction modeling framework. A Granger causality test (Granger, 1969; Roca, 2000) is then conducted to determine whether devaluation has independent explanatory power in a generalized model of inflation. Similarly, in a generalized model of devaluation, specified within the Granger causality modeling framework, the distributed lag effect of inflation on devaluation would be estimated and interpreted as the government's adjustment of exchange rates to inflation to prevent the real exchange rate from appreciating.

III. Exchange Rate Regimes and Inflation

One can identify four discrete exchange rate regimes in Bangladesh since the 1960s. Using data in Table 1, I briefly review these regimes with a view to highlighting any association between the exchange rate regime and inflation.

A. Before 1971

From 1947 until independence in 1971, the present territory of Bangladesh was part of Pakistan. The exchange rate of the Pakistani rupee was fixed at the rate of 4.793 to the dollar in the early 1960s and remained unchanged until April 1972. Annual inflation in Pakistan during the 1960s was less than 5 percent (Hossain, 2000a). However, this rate was high compared with those of Pakistan's major trading partners. Consequently, the rupee became overvalued in real effective terms and remained so throughout the 1960s. As Islam (1981) pointed out, this was part of Pakistan's strategy of development through forced industrialization.

Table 1. Devaluation and Inflation Under Successive Exchange Rate Regimes
(In percent a year)

Exchange Rate Regime	Change in Nominal Effective Exchange Rate[1]	CPI Inflation Rate[1]
Peg to pound sterling		
1972–74	-9.3[2]	48.3
1975–79	-9.2	9.8
Peg to currency basket, with pound sterling as intervention currency		
1980–82	-7.3	14.0
Peg to currency basket, with dollar as intervention currency		
1983–90	-7.4	9.8
1991–99	-1.9	5.6

Sources: See Appendixes I and II.
[1]Data are annual averages; a negative value for the change in the exchange rate indicates a depreciation.
[2]Data are for 1973–74.

B. 1972–75

In January 1972 the exchange rate of the taka, which replaced the rupee as the currency of the newly independent Bangladesh, was fixed against the pound sterling at 18.9677 to the pound. This new exchange rate represented a one-time devaluation of the taka against sterling of about 33 percent. However, because sterling floated against the dollar after January 23, 1972, following the breakdown of the Bretton Woods exchange rate agreement, the taka also quasi-floated against the dollar through its link to sterling.

Bangladesh experienced relatively high inflation during the early 1970s, caused primarily by excessive monetary expansion. Annual inflation averaged 48 percent during 1972–74, the highest rate the country had experienced since the 1950s. However, the nominal effective exchange rate during this period depreciated by only 9 percent during the same period. The resulting overvaluation caused a major balance of payments crisis in 1974 (Hossain, 2000a). It was not until May 1975 that the taka was devalued against sterling by about 40 percent. Thus the 1975 devaluation was essentially a lagged adjustment of the exchange rate of the taka against sterling. The exchange rate situation improved somewhat during the rest of the decade. Average annual inflation declined sharply, to about 10 percent during 1975–79. The nominal effective exchange rate also depreciated by about 9 percent a year during this period.

C. 1979–82

In 1979 the government pegged the taka to a basket of hard currencies and used the pound sterling as the intervention currency. During the next three years the nominal effective exchange rate depreciated by 7 percent a year, and annual inflation averaged about 14 percent. Unlike in the late 1970s, there was thus a considerable appreciation in terms of the real exchange rate, which later created some balance of payments problems.

D. 1983 to the Present

Since 1983 the taka has remained pegged to a basket of hard currencies, but the dollar has replaced the pound sterling as the intervention currency. Bangladesh started to deregulate its financial sector in the mid-1980s. Apparently, as part of the authorities' real exchange rate approach to exchange rate policy,[13] the frequency of adjustment of the exchange rate has increased since then, but the adjustments have not necessarily been complete. For example, the nominal effective exchange rate depreciated by about 7 percent a year during 1983–90, but annual average inflation was about 10 percent. This kept the real exchange rate roughly steady, but any policy of full adjustment was not strictly followed through the 1990s, plausibly because of the political sensitivity of devaluation in the midst of democratic, albeit *hartal*, politics (Hossain, 2000b). For example, annual inflation during 1991–99 averaged about 6 percent, yet the nominal effective exchange rate depreciated by only 2 percent a year. An appreciation of the real exchange rate resulted, which led to balance of payments problems.

This brief discussion reveals that the evolution of exchange rate policy in Bangladesh since the early 1970s has essentially been a move from a single-currency peg to a multiple-currency peg. There are two points of contention. First, do Bangladesh's exchange rate responses to inflation differentials between the home and foreign countries reflect a formal or informal real exchange rate targeting approach (Corden, 1991)? Second, was there any variation in the exchange rate response to inflation under the different exchange rate regimes? As pointed out earlier, little information is available on the conduct of exchange rate policy in Bangladesh. Therefore these questions cannot be answered adequately with statistical information. This paper takes the view that the government of Bangladesh adjusted the official exchange rate of the taka with the intervening currency under a loosely defined real exchange rate targeting approach.

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[13] Although the Bangladeshi authorities did not explicitly follow the real exchange rate targeting approach to exchange rate policy, such an approach was somewhat revealed by or could be discerned from the movement of the real exchange rate, especially since the early 1980s. However, this proposition is essentially a working hypothesis as part of establishing the point that external balance, rather than price stability, was the main objective of the government's monetary and exchange rate policy. For a detailed discussion of monetary policy in Bangladesh, see Hossain (2002).

The wide fluctuations of inflation rates under the different exchange rate regimes indicate that the government did not have much control over inflation under the pegged exchange rate system. Because it did not pursue an active monetary policy with the objective of maintaining price stability, the lagged adjustment of exchange rates to inflation was essentially an attempt to prevent the real exchange rate from appreciating to a level that would create unsustainable current account deficits. The policy paradigm and the adjustment mechanism outlined above are compatible with the fact that Bangladesh was under various IMF- and World Bank-supported structural adjustment programs throughout the 1980s and early 1990s. However, the data reported in Table 1 show that the exchange rate responses to inflation varied under different exchange rate policy regimes.

IV. Estimation Results

The procedure adopted for examining whether devaluation is an independent source of inflation in Bangladesh has three main parts, arranged in a sequential manner. The first part involves testing for a cointegral relationship among the consumer price index (CPI), the narrow (M1) or the broad (M2) money stock, and real output (RGDP) or industrial output (the industrial production index, IPI; the latter is used in the analyses using monthly data.[14] The presence of a cointegral relationship among these variables is interpreted as a long-term price-level relationship that can be derived from an equilibrium condition in the money market.[15] For example, equation (2.10) above can be considered a dynamic variant of a long-term price-level relationship. The second part of the procedure investigates the short-term effect of devaluation on inflation. This is done by estimating an error correction model of inflation in which devaluation, a stationary variable, is included with the contention that it may have a short-term effect on inflation. The third part investigates the plausible feedback effects between devaluation and inflation by conducting a Granger causality test within a generalized error correction modeling framework.

A. Cointegral Relationship Among Consumer Prices, the Money Stock, and Output

Using the conventional cointegration testing procedure of Engle and Granger (1987), this section investigates the presence of a long-term relationship among consumer prices, the money stock, and real output (or industrial output).[16]

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[14] See Appendix II for the time-series properties of all variables used in the empirical investigation.

[15] The equilibrium condition is $\ln (M/P) = \ln m^d(y, i)$, where M is the money stock, P is the price level, and m^d is demand for real balances, which depends positively on real income (y) and negatively on the nominal interest rate (i). The interest rate can be ignored in the specification if it is determined administratively or does not change much over time. This equilibrium condition gives a price-level equation that depends positively on the money supply and negatively on real income.

[16] Further tests are conducted by the Johansen procedure (Johansen, 1988, 1991; Johansen and Juselius, 1990). The results, which are consistent with those obtained by the Engle-Granger procedure, are not reported here but will be available from the author upon request.

The Engle-Granger two-stage procedure

The Engle-Granger approach involves a two-stage procedure in which a long-term relationship among variables that are I(1) variables, is established by estimating a regression equation containing such variables and then testing for unit roots on the residuals of the regression. In general, the ordinary least squares (OLS) method is used for the estimation, and any misspecification bias that emerges in the cointegration regression as a result of omitted dynamics being forced into the error term is removed by adding the first-difference terms of the explanatory variables, with or without lags.[17] The resulting estimate of the cointegration vector is considered equivalent to what may be obtained using the maximum likelihood estimation technique.

For estimation purposes, ln CPI is chosen as the dependent variable. In the results, the numbers in parentheses below the coefficients are absolute t-ratios, R^2 is the coefficient of determination adjusted for degrees of freedom, CRDW is the cointegration regression Durbin-Watson statistic, and DF and ADF are, respectively, the unaugmented and the augmented Dickey-Fuller statistics for unit roots in the residuals of the cointegrating regression. AIC is the Akaike information criterion and SBC the Schwarz Bayesian criterion. In results with monthly data, the estimated coefficients of the first-difference terms of explanatory variables and seasonal dummies are not reported. In general, the first-difference terms are unimportant and are included only to lower the finite sample bias in estimates of coefficients of variables in the level term.

I. Results with annual data

When annual data were used, the long-term price-level equation with narrow money (M1) gave the following results:

$$\begin{aligned} \ln CPI = \; & \underset{(2.13)}{4.99} + \underset{(5.81)}{0.83 \ln M1} - \underset{(1.36)}{0.60 \ln RGDP} \\ & - \underset{(3.26)}{0.65\, \Delta\ln M1(-1)} + \underset{(1.27)}{1.01\, \Delta\ln RGDP(-1)} \end{aligned}$$

Sample: 1974-99; R^2 = 0.98; CRDW = 0.71; ADF(3){AIC} = -2.93; DF{SBC} = -2.31.

[17] One reviewer questioned the appropriateness of using the first-difference terms of cointegrating variables while estimating a cointegral relationship, as it might lead to a misspecification error. The use of such first-difference terms is a common practice in the applied literature, however. As pointed out in the text, these first-difference terms are considered unimportant but are included to lower the finite sample bias in estimates of long-run coefficients of variables in level form (Phillips, 1988; Reserve Bank of New Zealand, 1991). Insofar as the present paper indicates, the first-difference terms are I(0) and therefore can be used in a regression in which long-run relationships are formed among variables that are I(1).

The same equation with broad money (M2) gave the following results:

$$\begin{array}{llll}
\ln CPI = & 5.10 & + 0.66 \ln M2 & - 0.62 \ln RGDP \\
 & (4.49) & (12.58) & (2.92) \\
 & \multicolumn{2}{l}{- 0.33\ \Delta\ln M2(-1) \quad + 1.01\Delta \ln RGDP(-1)} & \\
 & \multicolumn{2}{l}{(2.47) \qquad\qquad\quad (1.42)} &
\end{array}$$

Sample: 1974–99; $R^2 = 0.99$; CRDW = 1.40; ADF(3){AIC and SBC} = -4.61.

II. Results with monthly data

When monthly data were used, the long-term price-level equation with narrow money (M1) gave the following results:

$$\begin{array}{llll}
\ln CPI = & -0.49 & + 0.69 \ln M1 & - 0.14 \ln IPI \\
 & (4.54) & (55.29) & (3.80)
\end{array}$$

+ 12 lagged money growth terms + 12 lagged industrial output growth terms
+ seasonal dummies

Sample: 1973M2-99M12; $R^2 = 0.98$; DW = 0.16; DF{SBC} = -3.89; ADF(12){AIC} = -1.66.

The same equation with broad money (M2) gave the following results:

$$\begin{array}{llll}
\ln CPI = & 0.51 & + 0.51 \ln M2 & - 0.06 \ln IPI \\
 & (6.82) & (86.92) & (2.86)
\end{array}$$

+ 12 lagged money growth terms + 12 lagged industrial output growth terms
+ seasonal dummies

Sample: 1973M2-99M12; $R^2 = 0.99$; CRDW = 0.21; DF{SBC} = -4.58; ADF (12){AIC} = -2.74.

These results are consistent with theoretical expectations. They show that, in the long run, consumer prices rise with an expansion of the money supply but fall with an increase in real GDP (or its proxy, industrial output). The *t*-ratios for the coefficients of both money and real output are high (especially in the results with broad money) and significant at the conventional 5 percent level. The R^2 suggests that the regression equation possesses a high degree of explanatory power irrespective of the frequency of data and of which definition of the money variable is used. The DF and ADF test statistics, chosen on the basis of either the AIC or the SBC, show that the residuals of the cointegrating regression satisfy the stationarity condition.[18]

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[18] The value of the CRDW statistic is relatively low, however, conflicting with the ADF test results.

To summarize, the cointegration test results suggest that there is a long-term relationship among consumer prices, the money stock, and real GDP (or industrial output). This relationship is consistent with theoretical predictions that the price level rises with an expansion of the money supply and decreases with an increase in real output. Underlying such a price-level equation is a stable money demand function, which is confirmed for Bangladesh in various studies, including Hossain (2002).[19]

B. An Error-Correction Model of Inflation

Given that there is a cointegral relationship among consumer prices, the money stock, and real GDP (or industrial output), it is possible to specify and estimate an error correction model of inflation in the following generic form:

$$\Delta \ln CPI = \alpha 0 + \alpha 1\, EC(-1) + \Sigma \alpha i\, \mathbf{Z}_i\; (i = 2, \ldots, j) + \upsilon,$$

where *EC* is the residual of the cointegrating regression, **Z** is a vector of stationary variables that explain the short-term behavior of inflation, and υ is an error term with zero mean and a constant variance. In this specification the error correction term plays the critical role. It measures the speed of adjustment to the cointegrating relationship if the actual relationship deviates from the long-term relationship due to disturbances or shocks (Engle and Granger, 1987). The usual stationary regression theory applies to the error correction model. For example, because all variables in the above specification are required to satisfy the stationarity property, OLS remains the preferred method of estimation.

However, in order to arrive at a parsimonious model, the generalized form of the error correction model is subjected to systematic experimentation following, say, a general-to-specific modeling strategy that uses a range of diagnostic tests for model selection. When the error correction model fails the diagnostic tests, it may indicate problems with specification of the model. The error correction model may then be respecified until the diagnostic tests indicate that the model is theoretically consistent and statistically adequate. Sometimes theoretical considerations are given precedence over minor statistical inadequacies.

In the present case, the **Z** vector includes the lagged dependent variable Δln CPI(-1), the lagged first-difference terms of explanatory variables in the cointegrating equation, that is, Δln M1 or Δln M2, and Δln RGDP or Δln IPI, seasonal dummies, and any other variables of interest

[19] Given that a stable money demand function is a key component of the inflation model, one reviewer suggested that I estimate a money demand function, examine its stability, and test for causality between money supply growth and inflation. In a recent paper (Hossain, 2002), I adopted this approach to investigate the inflationary process in Bangladesh. To avoid repetition, therefore, I have not estimated a money demand function for this paper. However, the main finding of the earlier paper on money demand remains consistent with empirical results of the present paper.

provided they satisfy the stationarity condition. For example, even though the nominal effective exchange rate (NEER) does not enter into the long-term price-level relationship, the rate of devaluation (Δln NEER) can be included in the error correction model with the contention that it may have a short-term effect on inflation.

Results with annual data

After experimentation with the general form of the error correction model with two lagged terms for Δln M1 (or Δln M2), Δln RGDP, and Δln NEER, the following parsimonious specifications are found to fit the data best. In the results, the numbers in parentheses below the coefficients are absolute *t*-ratios.

The error correction model of inflation with narrow money (M1) gave the following results:

$$\begin{aligned}\Delta \ln CPI = \; & \underset{(8.26)}{0.07} - \underset{(3.19)}{0.17}\, EC(-1) + \underset{(8.31)}{0.53}\, \Delta \ln CPI(-1) \\ & - \underset{(4.13)}{0.74}\, \Delta \ln y(-2) + \underset{(2.53)}{0.12}\, \Delta \ln NEER(-1)\end{aligned}$$

Sample: 1975–99; $R^2 = 0.75$; Dh = -0.67. Additional statistics: serial correlation = $F(1, 19) = 0.17$; functional form: $F(1, 19) = 0.01$; heteroskedasticity: $F(1, 23) = 0.52$.

The same model with broad money (M2) gave the following results:

$$\begin{aligned}\Delta \ln CPI = \; & \underset{(7.82)}{0.07} - \underset{(3.07)}{0.35}\, EC(-1) + \underset{(1.08)}{0.50}\, \Delta \ln CPI(-1) \\ & - \underset{(4.01)}{0.72}\, \Delta \ln y(-2) + \underset{(1.08)}{0.06}\, \Delta \ln NEER(-1)\end{aligned}$$

Sample: 1975–99; $R^2 = 0.75$; Dh = -0.67. Additional statistics: serial correlation = $F(1, 19) = 0.17$; functional form: $F(1, 19) = 0.01$; heteroskedasticity: $F(1, 23) = 0.52$.

Results with monthly data

As earlier, after experimentation with the general form of the error correction model with 12 lagged terms for Δln M1 (or Δln M2), Δln IPI, and Δln NEER, the following parsimonious specifications are found to fit the data best. Again, the numbers in parentheses below the coefficients are absolute *t*-ratios.

The error correction model of inflation with narrow money (M1) gave the following results:

$$
\begin{array}{llll}
\Delta\ln \mathrm{CPI} = & -0.01 \quad -0.03\, EC(-1) & +\, 0.40\, \Delta\ln \mathrm{CPI}(-1) & \\
 & (0.56) \quad (3.72) & (7.99) & \\
 & +0.08\, \Delta\ln \mathrm{M1}(-9) & +\, 0.05\, \Delta\ln \mathrm{M1}(-11) & \\
 & (3.24) & (2.01) & \\
 & -\, 0.01\, \Delta\ln \mathrm{IPI}(-1) & -\, 0.01\, \Delta\ln \mathrm{IPI}(-2) & \\
 & (1.10) & (0.96) & \\
 & -\, 0.03\, \Delta\ln \mathrm{NEER}(-8) & -\, 0.04\, \Delta\ln \mathrm{NEER}(-12) & +\ \text{seasonal dummies} \\
 & (1.31) & (1.73) &
\end{array}
$$

Sample: 1973M3-99M12; $R^2 = 0.28$; Dh = -2.67. Additional statistics: serial correlation = $F(12, 290) = 2.20$; functional form: $F(1, 301) = 5.65$; heteroskedasticity: $F(1, 320) = 27.62$.

The same model with broad money (M2) gave the following results:

$$
\begin{array}{llll}
\Delta\ln \mathrm{CPI} = & -0.01 \quad -0.08\, EC(-1) & +\, 0.37\, \Delta\ln \mathrm{CPI}(-1) & \\
 & (0.15) \quad (6.03) & (7.75) & \\
 & +0.13\, \Delta\ln \mathrm{M2}(-9) & +\, 0.10\, \Delta\ln \mathrm{M2}(-11) & \\
 & (3.21) & (2.73) & \\
 & -\, 0.01\, \Delta\ln \mathrm{IPI}(-1) & -\, 0.01\, \Delta\ln \mathrm{IPI}(-2) & \\
 & (1.49) & (1.08) & \\
 & -\, 0.03\, \Delta\ln \mathrm{NEER}(-8) & -\, 0.05\, \Delta\ln \mathrm{NEER}(-12) & +\ \text{seasonal dummies} \\
 & (1.39) & (2.39) &
\end{array}
$$

Sample: 1973M3-99M12; $R^2 = 0.34$; Dh = -1.98. Additional statistics: serial correlation = $F(12, 290) = 2.23$; functional form: $F(1, 301) = 4.79$; heteroskedasticity: $F(1, 320) = 34.03$.

Consistent with the priors, the error correction term is significant whether the narrow or the broad definition of money is used in the regression. Some lagged terms of other variables, such as money growth, real output growth (or industrial output growth), and devaluation, are significant but at different levels of significance. In the specification using annual data, additional diagnostic statistics suggest no statistical problems. However, in the specification using monthly data, the relevant diagnostic statistics show some evidence of serial correlation, functional form misspecification, and heteroskedasticity. The long and complex lag structure and the possible presence of superfluous variables (such as devaluation) may be the source of these problems. The question of whether devaluation is superfluous in an inflation model can be examined using the Granger causality testing procedure.

C. Feedback Effects Between Devaluation and Inflation

In this section the plausible feedback effects between devaluation and inflation are examined within a generalized error correction modeling framework. This is done by augmenting Granger's basic causality model (Granger, 1969) in the generic form described below.[20]

The model

Let Δx_t and Δy_t be two stationary time series, with the property that x and y are cointegrated. The presence of this cointegral relationship indicates that either x or y or both have a causal effect on the other. Within an error correction modeling framework, the causal model between Δx_t and Δy_t can be specified as follows (Roca, 2000):

$$\Delta x_t = \alpha 0 + \alpha 1\ \mu_{t-1} + \Sigma\alpha i\ \Delta x_{t-i} + \Sigma\beta i\ \Delta y_{t-i} + u_t$$
$$\Delta y_t = \gamma 0\ + \gamma 1\ \mu_{t-1}\ + \Sigma\gamma i\ \Delta y_{t-i} + \Sigma\delta i\ \Delta x_{t-i} + v_t,$$

where μ_{t-1} are the lagged residuals from the level-form cointegrating regression between x and y, and u and v are uncorrelated white-noise series. In this specification, although i can take any value, it is finite and depends on the sample size.

The application of the concept of unidirectional Granger causality implies that y is causing x (denoted by $y \rightarrow x$), provided $\Sigma\beta i$ are not statistically equal to zero and $\Sigma\delta i$ are equal to zero. Similarly, x is unidirectionally causing y (denoted by $x \rightarrow y$) if $\Sigma\delta i$ are not statistically equal to zero and $\Sigma\beta i$ are equal to zero. A feedback or bidirectional causality exists between x and y (denoted by $x \Leftrightarrow y$) if $\Sigma\beta i$ and $\Sigma\delta i$ are not equal to zero. If $\Sigma\beta i = \Sigma\delta i = 0$, x and y are independent.

Estimation results

In implementing the above test in the present case, the relevant error correction term is taken from the Engle-Granger long-term price-level equation. In the specification using annual data, the number of lagged terms used is one initially but is then increased to two and finally to three. Similarly, in the specification using monthly data, the number of lagged terms used is 6 initially but is then increased to 12 and finally to 18. The reasoning behind a sequential increase in the number of lagged terms is that the causal inference drawn from causality tests is known to be sensitive to the choice of lag length. The maximum lag lengths of 3 years for annual data and of 18 months for

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[20] One reviewer suggested that I use the basic Granger causality framework where the error correction term does not appear. However, I think that because an error correction modeling approach is adopted here, the augmented causality framework is more appropriate. Without an error term in the specification, the results could be biased (Roca, 2000).

monthly data are considered long, in the sense that they allow a considerable amount of time for devaluation to exert its effect on inflation (or vice versa).[21]

Table 2 shows that, for the complete sample period, the null hypothesis that devaluation is not an independent source of inflation cannot generally be rejected by an F-test at the conventional 5 percent level of significance. This result remains robust irrespective of the number of lags and of which definition of the money variable is used for estimation purposes. By contrast, the maintained hypothesis that, under a pegged exchange rate system, it is inflation that leads to devaluation cannot generally be rejected by an F-test at least at the 10 percent level of significance. This result also remains robust irrespective of the number of lagged terms and of which money variable is used.

The results for the subsample 1973–82 are consistent with those for the complete sample period: inflation led to devaluation, whereas any effect of devaluation on inflation was not significant. However, the results for the second subsample, covering the 1980s and 1990s, suggest that the effect of inflation on devaluation became weaker during this period and that there was no significant effect of devaluation on inflation. Note that the latter results were for the period when the economy was opened up and the taka was pegged to a basket of currencies and then adjusted frequently, under various IMF- and World Bank-supported adjustment programs, to avoid real effective exchange rate appreciation. Given that the time lag of adjustment of exchange rates to inflation has been shortened gradually since the mid-1980s, the feedback effects between devaluation and inflation, being contemporaneous in nature, might have contributed to the weak relationship between them in a statistical sense. As Kenen and Pack (1980, p. 3) aptly put it long ago, "It is hard to determine the direction of causation, let alone the size and speed of the response, when exchange rates and price levels are both in motion."

In short, the overall results show that, for Bangladesh, although devaluation may have a short-term effect on inflation, it does not appear to be an independent source of inflation. On the contrary, there is evidence that it is inflation, caused by, say, excess money supply, that leads to devaluation.

D. A Monetary Model of Inflation

As indicated earlier, in a monetary model, devaluation does not appear as an independent source of inflation. The Granger causality test results obtained within a cointegration and error correction modeling framework are consistent with a monetary model of inflation. By imposing the restriction that devaluation does not have an effect on inflation and after eliminating lagged terms for money and real output (or industrial output) that are not significant, the following error correction models of inflation are found to fit the data best. In the results, the numbers in parentheses below the coefficients are absolute t-ratios.

[21] One reviewer suggested that I use the AIC to select the lag length. As indicated in the text, the reasoning behind a sequential increase in the number of lagged terms is that the causal inference drawn from causality tests is sensitive to the choice of lag length. By following this strategy, I have avoided selecting a lag length using a statistical criterion that is sometimes considered unreliable.

Table 2. Feedback Effects Between Devaluation and Inflation

Dependent Variable	No. of Lag Terms	F-Statistic[1] With M1	F-Statistic[1] With M2	Degrees of Freedom
	Annual data: Estimation period 1975–99[2]			
Inflation	1	2.98*	1.16	$F(1, 19)$
	2	1.50	0.45	$F(2, 15)$
	3	0.17	1.45	$F(3, 10)$
Devaluation	1	51.70***	43.21***	$F(1, 19)$
	2	4.00**	4.00**	$F(2, 15)$
	3	3.75**	0.25	$F(3, 10)$
	Monthly data: Estimation period 1973M3–99M12[3]			
Inflation	6	0.93	0.84	$F(6, 285)$
	12	0.98	0.91	$F(12, 261)$
	18	0.79	0.76	$F(18, 232)$
Devaluation	6	4.48***	3.30***	$F(6, 285)$
	12	5.09***	4.28***	$F(12, 261)$
	18	2.76***	2.43***	$F(18, 232)$
	Monthly data: Estimation period 1973M3–82M3[3]			
Inflation	6	2.92**	2.53**	$F(6, 72)$
	12	1.14	0.78	$F(12, 48)$
	18	1.20	0.83	$F(18, 19)$
Devaluation	6	2.22*	1.34	$F(6, 72)$
	12	2.70***	2.62***	$F(12, 48)$
	18	1.60	3.20***	$F(18, 19)$
	Monthly data: Estimation period 1982M5–99M12[3]			
Inflation	6	0.77	1.22	$F(6, 176)$
	12	0.75	0.91	$F(12, 152)$
	18	0.71	0.94	$F(18, 127)$
Devaluation	6	1.07	1.33	$F(6, 176)$
	12	1.29	1.30	$F(12, 152)$
	18	1.23	1.60	$F(18, 127)$

Source: Author's regressions.

Notes: *** indicates significance at the 1 percent level, ** at the 5 percent level, and * at the 10 percent level.

[1]Test of joint significance of $\Sigma\Delta\ln$ CPI($-i$) or $\Sigma\Delta\ln$ NEER($-i$) in the generalized equation.

[2]The estimating equations are of the following form:

$$\Delta \ln CPI = f(\text{Constant}, EC(-1), \Sigma\Delta\ln CPI(-i), \Sigma\Delta\ln Mj(-i), \Sigma\Delta\ln RGDP(-i), \Sigma\Delta\ln NEER(-i)) \ (i = 1, 2; j = 1, 2)$$

$$\Delta \ln NEER = g(\text{Constant}, EC(-1), \Sigma\Delta\ln CPI(-i), \Sigma\Delta\ln Mj(-i), \Sigma\Delta\ln IPI(-i), \Sigma\Delta\ln NEER(-i)) \ (i = 1, 2; j = 1, 2).$$

[3]The estimating equations are of the following form:

$$\Delta \ln CPI = f(\text{Constant}, EC(-1), \Sigma\Delta\ln CPI(-i), \Sigma\Delta\ln Mj(-i), \Sigma\Delta\ln IPI(-i), \Sigma\Delta\ln NEER(-i), \text{seasonal dummies} \ (i = 1, 2; j = 1, 2)$$

$$\Delta \ln NEER = g(\text{Constant}, EC(-1), \Sigma\Delta\ln CPI(-i), \Sigma\Delta\ln Mj(-i), \Sigma\Delta\ln IPI(-i), \Sigma\Delta\ln NEER(-i), \text{seasonal dummies} \ (i = 1, 2; j = 1, 2).$$

Results with annual data

The restricted error correction model of inflation with narrow money (M1), using annual data, gave the following results:

$$\begin{aligned} \Delta\ln \text{CPI} = \; & 0.07 \quad - 0.19\, EC(-1) \quad + 0.44\, \Delta\ln \text{CPI}(-1) \\ & (6.69) \quad (3.22) \qquad\qquad (4.73) \\ & + 0.07\, \Delta\ln \text{M1}(-2) - 0.76\, \Delta\ln y(-2) \\ & (1.12) \qquad\qquad\qquad (3.83) \end{aligned}$$

Sample: 1975-99; $R^2 = 0.69$; Dh = 0.61. Additional statistics: serial correlation: $F(1, 19) = 0.27$; functional form: $F(1, 19) = 0.22$; heteroskedasticity: $F(1, 23) = 0.62$.

The same model using broad money (M2) gave the following results:

$$\begin{aligned} \Delta\ln \text{CPI} = \; & 0.04 \quad - 0.44\, EC(-1) \quad + 0.53\, \Delta\ln \text{CPI}(-1) \\ & (3.26) \quad (5.11) \qquad\qquad (9.04) \\ & + 0.16\, \Delta\ln \text{M2} \qquad\quad -0.77\, \Delta\ln y(-2) \\ & (2.52) \qquad\qquad\qquad\quad (4.76) \end{aligned}$$

Sample: 1975–99; $R^2 = 0.80$; Dh = -0.77. Additional statistics: serial correlation: $F(1, 19) = 0.48$; functional form: $F(1, 19) = 3.19$; heteroskedasticity: $F(1, 23) = 0.60$.

Results with monthly data

The restricted error correction model of inflation with narrow money (M1), using monthly data, gave the following results:

$$\begin{aligned} \Delta\ln \text{CPI} = \; & -0.01 \quad -0.03\, EC(-1) \quad + 0.39\, \Delta\ln \text{CPI}(-1) \\ & (0.69) \quad (3.76) \qquad\qquad (7.93) \\ & +0.08\, \Delta\ln \text{M1}(-9) \qquad + 0.05\, \Delta\ln \text{M1}(-11) \\ & (3.14) \qquad\qquad\qquad\quad (1.88) \\ & + \text{seasonal dummies} \end{aligned}$$

Sample: 1973M3-99M12; $R^2 = 0.28$; Dh = -2.31. Additional statistics: serial correlation = $F(12, 294) = 2.09$; functional form: $F(1, 305) = 6.82$; heteroskedasticity: $F(1,320) = 29.68$.

The same model using broad money (M2) gave the following results:

$$\begin{aligned} \Delta \ln \text{CPI} = \; & -0.01 \quad -0.07\, \text{EC}(-1) \quad + 0.38\, \Delta \ln \text{CPI}(-1) \\ & (0.39) \quad (5.91) \qquad\qquad (7.82) \\ & + 0.12\, \Delta \ln \text{M2}(-9) \qquad + 0.10\, \Delta \ln \text{M2}(-11) \\ & (3.01) \qquad\qquad\qquad\quad (2.49) \\ & + \text{seasonal dummies} \end{aligned}$$

Sample: 1973M3-99M12; $R^2 = 0.33$; Dh = -1.56. Additional statistics: serial correlation = $F(12, 294) = 2.22$; functional form: $F(1, 305) = 5.24$; heteroskedasticity: $F(1, 320) = 30.82$.

These results are in line with the monetary model of inflation. The coefficient of the error correction term shows the tendency of the price level to revert to its long-term equilibrium state if the actual relationship deviates from its long-term state temporarily as a result of disturbances or shocks. The models using monthly data explain about 30 percent of the total variation of inflation, and the models using annual data with broad money explain about 80 percent. The diagnostic statistics indicate some statistical problems for the models with monthly data, but the annual models do not. To examine the stability of the parameters, the models were also estimated by the recursive estimation procedure. This yielded estimates of the parameters that were found to be stable, especially those obtained with broad money.[22]

V. Concluding Remarks

This paper has investigated, within an error-correction modeling framework, the possibility of feedback effects between devaluation and inflation in Bangladesh during 1972–99. Both annual and monthly data have been used. The empirical results suggest that there was no significant effect of devaluation on inflation during the complete sample period. This finding was also valid for the subsample that ranges from the early 1980s to the 1990s. Therefore the contention that devaluation has a significant effect on inflation cannot be sustained, especially in a deregulated global setting in which exchange rates and prices may not necessarily show a causal relationship but could be determined by complex factors within a generalized economic system.[23] A related but equally important finding is that past CPI inflation led to devaluation, especially during 1973–82, when the pegged exchange rate remained largely unchanged for a few years at a time while the country was experiencing high and unstable inflation. This finding was robust, irrespective of the frequency of data and the definition of the money variable used in the tests. Thus the overall results are consistent with the view that, in Bangladesh, it is past inflation that usually leads to devaluation.

These findings have policy implications for developing countries in general and Bangladesh in particular. Devaluation remains a controversial policy instrument in developing countries. However, given its importance, the role of devaluation in macroeconomic management needs to be evaluated in a broader context. In particular, devaluation should not be seen in isolation and is never a first-best policy choice for macroeconomic management. Its role is as part of a policy package that aims to reduce macroeconomic imbalances, manifested in inflation and current

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[22] The detailed results are not presented here but are available from the author upon request.

[23] The overall results should be interpreted with caution. As indicated earlier, I have not taken into account some indirect effects of devaluation on inflation, which would require specifying and estimating a structural model. Moreover, I have adopted a pragmatic, judgmental approach in conducting tests and interpreting results.

account deficits. However, devaluation has other effects, especially on income distribution, that need to be accounted for.

Although this paper has not examined in detail the source of inflation in Bangladesh, the results show that it can be explained adequately by a monetary model where broad money supply growth and real output growth remain two key determinants of inflation. If a monetary interpretation of inflation is accepted, high inflation can be considered the root cause of devaluation in Bangladesh. The ideal way to avoid devaluing the currency is to keep inflation low through restrictive monetary and fiscal policy. This can be done by reorienting monetary policy toward price stability as its core objective (Hossain, 2002). Low inflation, in turn, would create an environment for more rapid economic growth, which itself may reduce downward pressure on the exchange rate and even lead to an appreciation.

As an alternative to the present, adjustable-pegged exchange rate system, Bangladesh may wish to consider, on both economic and political grounds, introducing a floating exchange rate system. On the economic side, such a system, in addition to bringing some policy discipline, is desirable in the face of foreign capital inflows. It may also reduce the political sensitivity of devaluation provided that devaluation takes place on a small scale and on a steady basis. However, a flexible exchange rate system is alleged to have an inflationary bias unless there are institutional checks and balances on fiscal and monetary policy that Bangladesh now lacks.[24]

Given such pros and cons, the choice of exchange rate regime is not clear cut. What matters is a set of sound economic policies that remain consistent with any chosen exchange rate regime. Provided that the economic and political fundamentals remain strong, a high degree of volatility of the exchange rate (and even a sharp depreciation under a floating exchange rate system) will not necessarily lead to inflation. Moreover, a high degree of exchange rate volatility may be a reflection of policy uncertainties rather than the characteristic of a floating exchange rate regime.

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[24] This view is reflected in a widely cited early study by Aghevli, Khan, and Montiel (1991, p. 13), who report that "the inflation performance of the countries that have operated under a fixed exchange rate regime has been, on the whole, superior to that of the group operating under more flexible arrangement." However, a subsequent IMF study (1997, p. 87) raises doubt about such a relationship: "... it is not the case that flexible exchange rates are necessarily associated with higher inflation, as there are a number of countries with flexible exchange rate arrangements that have had relatively low inflation (and robust growth)."

Estimation of the Nominal Effective Exchange Rate

A country's nominal effective exchange rate is an index of the weighted average, with a common base, of bilateral nominal exchange rates between the country's currency and those of its major trading partners. The weight assigned to each bilateral exchange rate reflects the importance of that country as a trading partner. The effective exchange rate is considered a better measure than any single bilateral exchange rate of whether a country's currency has weakened or strengthened on average over a period against the currencies of its major trading partners.

This paper uses the nominal effective exchange rate for empirical analysis. Unfortunately, for Bangladesh this series is available only from the early 1980s onward. Therefore both annual and monthly data series for this variable were estimated for the entire sample period.

Given the availability of basic data, such as trade shares and bilateral exchange rates, the annual series for the effective exchange rate was estimated using data for Bangladesh's 21 major trading partners: Australia, Canada, China, France, Germany, Hong Kong SAR, India, Indonesia, the Islamic Republic of Iran, Italy, Japan, the Republic of Korea, Malaysia, the Netherlands, Pakistan, Saudi Arabia, Singapore, Sri Lanka, Thailand, the United Kingdom, and the United States. For estimation of the monthly effective exchange rate, data for 19 of these countries are used; complete data for either the monthly exchange rate or the trade share are not available for Sri Lanka and the Islamic Republic of Iran.

The monthly data for the bilateral exchange rates are generated as the unweighted average of bilateral daily exchange rates over the month. These exchange rates are available against the dollar. Except for Bangladesh, Pakistan, and Saudi Arabia, the data source for bilateral exchange rates is the website of the U.S. Federal Reserve: www.federalreserve.gov/releases/H10/hist. Data for the bilateral exchange rates for Bangladesh, Pakistan, and Saudi Arabia are taken from various issues of IMF, *International Financial Statistics* (IFS). IFS is also used to fill in random data gaps for other countries. The monthly bilateral exchange rates of the different currencies against the dollar are then used to calculate the exchange rates of Bangladesh's taka against each currency, in foreign currency units to the taka. These bilateral exchange rates are then converted into index form with a common base of 1995 = 1.000 (using the 12-month unweighted average for that year).

The second part of the estimation process is the calculation of trade shares (exports plus imports) for each of the selected countries and the transformation of those shares into normalized weights. The share of each country in Bangladesh's total trade (also exports plus imports) is calculated as a percentage. Country-specific annual data for Bangladesh's trade flows during 1970–99 are used for this purpose and are drawn from various issues of IMF, *Direction of Trade Statistics*. Together the selected countries on average represent the source or the destination of about 70 percent of Bangladesh's total trade flows for the period. However, the trade shares of individual countries with Bangladesh changed significantly over this period. Once the trade shares have been calculated, they are transformed into normalized weights by dividing each country's trade shares by the sum of trade shares of all the countries combined such that the sum of all normalized trade shares equals one. Table I.1 reports the normalized weights, which are calculated for each five-year period beginning with 1970.

Table I.1. Normalized Weights for Currencies of Major Trading Partners

Country	1970-75	1976-80	1981-85	1986-90	1991-95	1996-99
United States	0.43537	0.23532	0.15815	0.19702	0.19688	0.19140
Japan	0.08212	0.13972	0.14591	0.16932	0.09983	0.07625
Germany	0.05749	0.05208	0.05045	0.05104	0.07079	0.06750
France	0.01306	0.01999	0.02082	0.02206	0.03421	0.04347
Italy	0.01627	0.02374	0.01856	0.03516	0.03462	0.03069
Netherlands	0.01648	0.03320	0.04274	0.03962	0.03225	0.03348
United Kingdom	0.07605	0.09827	0.06718	0.06064	0.06661	0.07393
Canada	0.05591	0.06132	0.04716	0.04297	0.02763	0.01871
China	0.00001	0.02051	0.03164	0.04417	0.05647	0.07536
Hong Kong SAR	0.01058	0.02967	0.02304	0.04140	0.07377	0.06227
Australia	0.05092	0.03921	0.03620	0.02624	0.01801	0.02185
India	0.12248	0.04407	0.03398	0.03780	0.08663	0.12725
Indonesia	0.00524	0.00202	0.01769	0.01542	0.01200	0.01842
Korea, Rep. of	0.00360	0.01147	0.01722	0.03298	0.05504	0.04277
Malaysia	0.00258	0.01199	0.01848	0.01365	0.01089	0.01763
Pakistan	0.00584	0.03957	0.03368	0.02539	0.02469	0.01372
Singapore	0.01913	0.04127	0.04127	0.08843	0.08866	0.06002
Thailand	0.00767	0.01007	0.02288	0.01998	0.01049	0.01449
Iran, Isl. Rep. of	0.01355	0.04631	0.02761	0.01350	0.00956	0.00655
Saudi Arabia	0.00560	0.03954	0.09815	0.02299	0.01961	0.01258

Source: Author's calculations.

Whereas annual normalized weights are used for estimation of the annual effective exchange rate, for estimation of the monthly effective exchange rate unweighted averages of normalized weights over a five-year period are used. There are no accepted guidelines for the interval over which to adjust trade weights to accommodate any changes in trade patterns. There are both advantages and disadvantages to using relatively fixed, or frequently varied, weights for estimation purposes. For the present purposes, the use of the two sets of weights for estimation of the annual and monthly series has allowed some flexibility in the choice of weighting patterns. Thus, in the case of the annual effective exchange rate, both the weights and the exchange rates change every year, whereas for the monthly effective exchange rate, the weights change only every five years, but the exchange rates change every month.

Descriptive Statistics of CPI Inflation and Devaluation

Table I.2 reports descriptive statistics of both CPI inflation and devaluation, measured as the first-order logarithmic difference of the nominal effective exchange rate, for the complete period and two subperiods. The table shows that, for the annual data, the coefficient of variation of the change in the nominal effective exchange rate (Δln NEER) increased sharply from 0.92 in the 1980s to 4.21 in the 1990s; for the monthly data the coefficient of variation increased from 2.67 to 17.69. This confirms the contention that the exchange rate exhibits greater volatility in a floating or adjustable pegged exchange rate system than in a fixed exchange rate system. However, there was no significant increase in the volatility of CPI inflation from the 1970s and 1980s to the 1990s, especially when volatility is computed from the monthly data.

Table I.2. Descriptive Statistics on Inflation and Devaluation

	1973-82		1982-91		1992-99	
Statistic	Δln NEER	Δln CPI	Δln NEER	Δln CPI	Δln NEER	Δln CPI
Annual data						
Maximum	0.04	0.43	0.03	0.12	0.06	0.08
Minimum	-0.35	-0.02	-0.21	-0.07	-0.09	0.00
Mean	-0.10	0.17	-0.08	0.09	-0.01	0.05
Standard deviation	0.13	0.14	0.07	0.01	0.05	0.02
Skewness	-0.91	1.01	-0.18	0.02	-0.06	-0.62
Kurtosis –3	-0.25	-0.31	-0.69	-0.25	-0.78	0.22
Coeff. of variation	1.29	0.84	0.92	0.15	4.21	0.51
	1972M2-82M3		1982M4-91M3		1991M4-99M12	
Monthly data						
Maximum	0.16	0.11	0.03	0.045	0.21	0.03
Minimum	-0.48	-0.05	-0.07	-0.03	-0.21	-0.01
Mean	-0.008	0.015	-0.007	0.008	-0.002	0.003
Standard deviation	0.049	0.026	0.019	0.014	0.032	0.007
Skewness	-6.63	0.79	-0.79	0.45	-0.04	0.28
Kurtosis –3	64.90	1.40	1.09	0.42	32.87	0.93
Coeff. of variation	6.51	1.73	2.67	1.82	17.69	1.96

Source: Author's calculations based on data sources listed in Appendix II.

Data Issues

Definitions of Variables and Data Sources

Both annual and monthly data are used for empirical analysis. For the annual data the sample period is 1973–99, and for the monthly data it is 1972M1–99M12. As lags of different lengths are used in the estimation, the sample periods are adjusted and reported accordingly. Variables used in the analysis include the narrow money stock (M1) in millions of taka, the broad money stock (M2) in millions of taka, the consumer price index (CPI), gross domestic product at constant prices (RGDP) in millions of taka, the industrial production index (IPI), and the trade-weighted nominal effective exchange rate. Because monthly data for RGDP are not available, the IPI is used as a proxy for RGDP in regressions with monthly data. In regressions with annual data the definitions of variables remain the same except for RGDP, which remains the preferred measure of output.

All the basic data are obtained from both domestic and international publications, such as the *Bangladesh Bank Bulletin*, *Economic Trends* (Bangladesh Bank), *Economic Indicators of Bangladesh* (Bangladesh Bureau of Statistics), and IFS. I have estimated the data series for the trade-weighted nominal effective exchange rate. Appendix I reports the procedure followed for estimation purposes and the sources of the basic data. In order to fill in random gaps in the monthly data, I have made adjustments or interpolations to derive a consistent data series.

All the monthly data remain seasonally unadjusted. Seasonally unadjusted data are sometimes better than seasonally adjusted data because the latter may generate artificial smoothness and lower the power of unit root and cointegration tests (Ghysels, 1990; Davidson and MacKinnon, 1993). However, seasonal dummy variables are used in all regressions with the monthly data, although, to conserve space, the estimated coefficients are not reported. All the data series used for regression analysis are transformed into natural logarithms. Such transformed data are believed to improve the properties of various test statistics, and the results derived from them are sometimes easier to interpret. For statistical computations and regression analysis, Microfit 4.0 for Windows, developed by Pesaran and Pesaran (1997), is used. Regression

The Time-Series Properties of Variables

Two tests are conducted to determine the presence of unit roots in all the data series: the augmented Dickey-Fuller (ADF) test and the Phillips-Perron test. The findings of these tests are complemented by visual inspection of the plotted data series both as levels and as first differences.[25] In general, an eclectic approach to unit root testing is useful because the power of most unit root tests is not high, especially with high-frequency data.

[25] The detailed results are not presented here but are available from the author upon request.

Tables II.1 and II.2 present the results of the unit root tests. Table II.1 report the ADF test results, and Table II.2 the Phillips-Perron test results. For the ADF test, two sets of results are presented: one with an intercept but not a trend, and the other with both an intercept and a trend. . Two lag terms are used with the annual data and 12 lag terms with the monthly data. These lag terms are found adequate to transform the residuals in the regressions into white noise.[26]

Visual inspection of the data series reveals that all the variables have a trend in level form but that none have a trend in first-difference form. This suggests that, for the series in level form, the results with both an intercept and a trend are relevant and that, for the series in first-difference form, the results with an intercept but not a trend are relevant. Accordingly, the ADF test statistics in Table II.1 suggest that the null of a unit root in all the variables (except for industrial output) in level form cannot be rejected at the conventional 5 percent level. The variables in first-difference form, whether or not a trend is included, are stationary at the 5 percent level.[27]

The Phillips-Perron test results reported in Table II.2 do not reject the null of a unit root in all the variables (except for real output or industrial output) in level form. And, except for ln M2, all the variables are stationary in first-difference form. The variable ln M2 appears to be I(2), which does not seem consistent with the properties of the other series.

In any case, both sets of results suggest that industrial output is trend-stationary. If these results are accepted, they suggest exclusion of industrial output from the price-level specification. However, because industrial output is used as a proxy for RGDP, which, according to the ADF test results, has a unit root, it is prudent from a theoretical standpoint to use industrial output in the price-level regression. Thus, in short, the regression analysis is based on the contention that all the variables have a unit root but are stationary in their first-difference form.

[26] When both the AIC and the SBC were applied, results with optimal lag lengths were also generated. However, these results are not reported, because they are similar to those obtained with common lag lengths.

[27] Conflicting results were found for ln M2. The results obtained following the defined testing procedure show that ln M2 is I(2). However, when the ADF(12) command facility of Pesaran and Persaran (1997) is applied, the series, reported with an asterisk mark, is found to be I(1). The latter result is accepted for the present purposes as it is consistent with the properties of other series.

Table II.1 Results of Augmented Dickey-Fuller Tests for a Unit Root

		ADF Statistic[1]	
Estimation Period	Variable	Regression with intercept but without linear trend[2]	Regression with intercept and linear trend[3]
Annual data[4]			
1975–99	ln M1	-1.90	-1.68
1975–99	ln M2	-3.54	-0.03
1973–99	ln CPI	-2.34	-2.19
1975–99	ln RGDP	0.56	-1.45
1975–99	ln NEER	-2.89	-2.66
1976–99	Δln M1	-3.14	-3.92
1976–99	Δln M2	-2.01	-3.49
1974–99	Δln CPI	-5.31	-7.37
1976–99	Δln RGDP	-3.34	-3.13
1976–99	Δln NEER	-3.90	-4.08
Monthly data[5]			
1973M2–1999M12	ln M1	-0.52	-1.70
1973M2–1999M12	ln M2	-1.09	-0.44
1973M2–1999M12	ln CPI	-2.44	-3.11
1973M2–1999M12	ln IPI	-3.25	-4.34
1973M2–1999M12	ln NEER	-1.85	-1.81
1973M2–1999M12	Δln M1	-4.65	-4.60
1973M2–1999M12	Δln M2	0.25	0.45
	Δln M2*	-4.74	-4.67
1973M2–1999M12	Δln CPI	-2.89	-2.98
1973M2–1998M12	Δln IPI	-6.45	-6.41
1973M2–1999M12	Δln NEER	-5.30	-5.50
1973M2–1999M12	ΔΔln M2	-7.88	-7.88

Source: Author's calculations.

[1] When the sample size is 100, the 95 percent critical value for the ADF test statistic with an intercept and no linear trend is -2.87. The critical value for the same statistic with an intercept and a linear trend is -3.43 (Pesaran and Pesaran, 1997). When the sample size is 25, the 95 percent critical value for the ADF test statistic with an intercept and no linear trend is -3.00. The critical value for the statistic with an intercept and a linear trend is -3.60 (Fuller, 1976, Table 8.5.2)

[2] For a generic trended variable y in first-difference form that follows an AR(1) process: $y_t = \alpha + (1 - \phi)\delta t + \phi y_{t-1}$, the testing model should have an intercept (α) but not a linear trend, such that

$$\Delta\Delta y_t = \rho\delta - \rho\Delta y_{t-1} + \Sigma\eta_i\Delta\Delta y_{t-i}\ (i = 1, 2, \ldots, L) + \text{seasonal dummies}$$

for the monthly data, with the null hypothesis H_0: $\rho = 1 - \phi = 0$ or $\phi = 1$.

[3] For the variable in level form, the testing model should have an intercept and a linear trend (t), such that

$$\Delta y_t = \alpha + \rho\delta t - \rho y_{t-1} + \Sigma\eta_i\Delta y_{t-i}\ (i = 1, 2, \ldots, L) + \text{seasonal dummies},$$

for the monthly data, with the null hypothesis: H_0: $\rho = 1 - \phi = 0$ or $\phi = 1$. If the absolute value of $\phi < 1$, it is trend stationary; if $\phi = 1$, it is difference stationary with a nonzero drift α.

[4] Number of lags is 2.

[5] Number of lags is 12.

Table II.2 Results of Phillips-Perron Tests for a Unit Root

		Newey-West Lag Windows		
Estimation period	Variable	1	2	3
Annual data				
1973–99	ln M1	-2.87	-2.78	-2.88
1973–99	ln M2	-2.44	-2.28	-2.23
1971–99	ln CPI	-1.26	-1.40	-1.54
1973–99	ln RGDP	-7.48	-8.33	-8.44
1973–99	ln NEER	-1.69	-1.68	-1.65
1974–99	Δln M1	-9.43	-10.38	-10.53
1974–99	Δln M2	-4.87	-5.63	-6.29
1972–99	Δln CPI	-2.38	-3.50	-3.87
1974–99	Δln RGDP	-3.92	-4.10	-4.59
1974–99	Δln NEER	-5.02	-4.91	-5.02
		Newey-West Lag Windows		
Monthly data		12	24	36
1973M2–1999M12	ln M1	-2.70	-2.49	-2.47
1973M2–1999M12	ln M2	-0.42	-0.45	-0.46
1973M2–1999M12	ln CPI	-3.11	-2.67	-2.55
1973M2–1999M12	ln IPI	-5.62	-5.59	-5.44
1972M2–1999M12	ln NEER	-1.53	-1.58	-1.69
1973M2–1999M12	Δln M1	-30.75	-37.08	-39.85
1973M2–1999M12	Δln M2	0.23	0.33	0.40
1973M2–1999M12	Δln CPI	-5.14	-4.77	-4.66
1973M2–1998M12	Δln IPI	-27.06	-25.37	-25.26
1972M2–1999M12	Δln NEER	-10.74	-10.92	-10.95

Source: Author's calculations.

Note: The test results for the level form of variable y are derived from the following generic regression that includes an intercept and a deterministic trend: $\Delta \ln y_t = \alpha + \delta\rho t - \rho \ln y_{t-1}$ + seasonal dummies (for the monthly data), with the null hypothesis that $\rho = 1 - \phi = 0$ or $\phi = 1$. The test results are presented for three Newey-West lag windows, which are used to correct the test statistic for heteroskedasticity in the regression's error term. The reported lags are large enough to reduce bias in the asymptotic distribution of the test statistic. The test results for the first-difference form of variable y are derived form the generic equation, $\Delta\Delta \ln y_t = \delta\rho - \rho\Delta\ln y_{t-1}$ + seasonal dummies, for the monthly data. The critical values for the Phillips-Perron test are the same as for the Dickey-Fuller test and depend on whether the Dickey-Fuller regression contains an intercept term or a time trend (Pesaran and Pesaran, 1997).

References

Agénor, Pierre-Richard, and Peter Montiel, 1999, *Development Macroeconomics,* (Second edition) (Princeton, New Jersey: Princeton University Press).

Aghevli, Bijan B., Mohsin S. Khan, and Peter J. Montiel, 1991, *Exchange Rate Policy in Developing Countries: Some Analytical Issues*, IMF Occasional Paper 78 (Washington: International Monetary Fund).

Cassel, Gustav, 1918, "Abnormal Deviations in International Exchanges," *Economic Journal*, Vol. 28 (December), pp. 413–15.

Cooper, Richard, 1971, "Currency Devaluation in Developing Countries," Essays in International Finance 86 (Princeton, New Jersey: International Finance Section, Department of Economics, Princeton University).

Corbo, Vittorio, 1974, *Inflation in Developing Countries: An Econometric Study of Chilean Inflation* (Amsterdam: North-Holland).

Corden, W. Max, 1991, "Exchange Rate Policy in Developing Countries," in *Trade Theory and Economic Reform: North, South and East,* ed. by Jaime de Melo and Andre Sapir (Oxford: Basil Blackwell).

———, 1993, "Exchange Rate Policies for Developing Countries," *Economic Journal*, Vol. 103 (January), pp. 198–207.

Davidson, Russel, and James MacKinnon, 1993, *Estimation and Inference in Econometrics* (New York: Oxford University Press).

Dickey, David, and Wayne Fuller, 1981, "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root," *Econometrica*, Vol. 50, pp. 1057–72.

Dornbusch, Rudiger, 1987, "Exchange Rates and Prices," *American Economic Review*, Vol. 77 (March), pp. 93–106.

———, 1988, *Exchange Rates and Inflation* (Cambridge, Massachusetts: MIT Press).

Edwards, Sebastian, 1989a, *Real Exchange Rates, Devaluation, and Adjustment: Exchange Rate Policy in Developing Countries* (Cambridge, Massachusetts: MIT Press).

———, 1989b, *Exchange Rate Misalignment in Developing Countries* (Baltimore: Johns Hopkins University Press).

Engle, Robert, and Clive Granger, 1987, "Co-Integration and Error Correction: Representation, Estimation, and Testing," *Econometrica*, Vol. 55, (March), pp. 251–76.

Feenstra, Robert, 1995, "Estimating the Effects of Trade Policy," in *Handbook of International Economics*, Vol. III, ed. by Gene Grossman and Kenneth Rogoff (Amsterdam: Elsevier Science).

Fuller, Wayne, 1976, *Introduction to Statistical Time Series* (New York: John Wiley).

Ghysels, Eric, 1990, "Unit Root Tests and the Statistical Pitfalls of Seasonal Adjustment: The Case of U.S. Post-War Real Gross National Product," *Journal of Business and Economics Statistics*, Vol. 8, No. 2, pp. 145–52.

Granger, Clive, 1969, "Investigating Causal Relations by Econometric Models and Cross-Spectral Methods," *Econometrica*, Vol. 37 (January), pp. 424–38.

Harberger, Arnold, 1963, "The Dynamics of Inflation in Chile," in *Measurement in Economics: Studies in Mathematical Economics in Memory of Yehuda Grunfeld,* ed. by Carl Christ (Stanford, California: Stanford University Press).

Hossain, Akhtar, 1996, *Macroeconomic Issues and Policies: The Case of Bangladesh* (New Delhi: Sage Publications).

———, 2000a, *Exchange Rates, Capital Flows and International Trade* (Dhaka: University Press Limited).

———, 2000b, "Anatomy of *Hartal* Politics in Bangladesh," *Asian Survey*, Vol. 40, (May/June), pp. 508–29.

———, 2002, "An Examination of Some Key Issues in the Conduct of Monetary Policy in Bangladesh," in *Studies on Socio-Economic and Political Development of Bangladesh*, ed. by Ashraf Ali, Ruhul Kuddus, and Syed Andaleeb (Dhaka: University Press Limited, forthcoming).

———, and Anis Chowdhury, 1996, *Monetary and Financial Policies in Developing Countries: Growth and Stabilisation* (London: Routledge).

International Monetary Fund, 1997, *World Economic and Financial Surveys: World Economic Outlook*, October (Washington: International Monetary Fund).

Islam, Nurul, 1981, *Foreign Trade and Economic Controls in Development: The Case of United Pakistan* (New Haven, Connecticut: Yale University Press).

Johansen, Soren, 1988, "Statistical Analysis of Cointegration Vectors," *Journal of Economic Dynamics and Control*, Vol. 12, pp. 231–54.

———, 1991, "Estimation and Hypothesis Testing of Cointegrating Vectors in Gaussian Vector Autoregressive Models," *Econometrica*, Vol. 59, pp. 1551–80.

———, and Katarina Juselius, 1990, "Maximum Likelihood Estimation and Inference on Cointegration with Application to the Demand for Money," *Oxford Bulletin of Economics and Statistics*, Vol. 52 (May), pp. 169–210.

Kamin, Steven, 1988, "Devaluation, External Balance, and Macroeconomic Performance: A Look at the Numbers," Princeton Studies in International Finance 62 (Princeton, New Jersey: International Finance Section, Department of Economics, Princeton University).

Kenen, Peter, and Clare Pack, 1980, "Exchange Rates, Domestic Prices, and the Adjustment Process: A Survey of Recent Evidence," Occasional Papers 1 (New York: Group of Thirty).

Krugman, Paul, 1987, "Pricing to Market When the Exchange Rate Changes," in *Real-Financial Linkages Among Open Economies,* ed. by Sven Arndt and J. David Richardson (Cambridge, Massachusetts: MIT Press).

———, and Lance Taylor, 1978, "Contractionary Effects of Devaluations," *Journal of International Economics*, Vol. 8, pp. 445–56.

Mishkin, Frederic, 1998, *The Economics of Money, Banking, and Financial Markets* (New York: Addison-Wesley).

Pesaran, M. Hashem, and Bahram Pesaran, 1997, *Working with Microfit 4.0: Interactive Econometric Analysis* (London: Oxford University Press).

Phillips, Peter, 1988, "Optimal Inference in Cointegrated Systems," University of Auckland Discussion Paper No. 42 (Auckland, New Zealand: University of Auckland).

Prachowny, Martin, 1975, *Small Open Economies* (Lexington, Massachusetts: Lexington Books).

Reserve Bank of New Zealand, 1991, *Reserve Bank Model of the New Zealand Economy* (Auckland: Reserve Bank of New Zealand).

Roca, Eduardo, 2000, *Price Independence Among Equity Markets in the Asia-Pacific Region: Focus on Australia and ASEAN* (Aldershot, England: Ashgate).

Sachs, Jeffrey, and Felipe Larrain, 1993, *Macroeconomics in the Global Economy* (New York: Harvester Wheatsheaf).

Taylor, Lance, 1979, *Macro Models for Developing Countries* (New York: McGraw-Hill).

———, 1983, *Structuralist Macroeconomics* (New York: Basic Books).