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Measuring Inflationary Pressure in Bangladesh: The P-Star Approach

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February 2009

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Measuring Inflationary Pressure in Bangladesh: The P-Star Approach

Abstract

The paper estimates the P model for the Bangladesh economy and tests its forecasting ability through generating recursive forecasts. The model puts together long run determinants of price level based on the classical quantity theory of money and short term changes in current inflation. The empirical results show that the model performs relatively well and contains additional information regarding future rates of inflation. The price and output gap models fare consistently better than the velocity gap model which brings out the importance of non-monetary factors in explaining inflation dynamics in Bangladesh. The out of sample forecasts show that the price gap model performs better followed by the output gap model and the velocity gap model. With financial sector liberalization and reforms, it is likely that the scope for the P* model to play a more proactive role would be ramified in Bangladesh.*

I. INTRODUCTION

Over the years, rising inflation has emerged as one of the major concerns across the countries in the world. Although the claim of the traditional theory is that there is a trade-off between inflation and growth (alternatively between inflation and unemployment) and inflation control is the most important objective of monetary policy, recent evidence shows that such relationships are somewhat elusive and unstable and, even if such relationship exists, it is more likely to be non-linear and the desired rate of inflation can change in space and over time (Bruno 1995, Bruno and Easterly 1996, Pollin and Zhu 2005).

Despite such findings, concerns regarding high inflation are significant in countries like Bangladesh due both to potential threat of rising inflation to macroeconomic stability and its negative welfare consequences on different socioeconomic groups especially on the poor and vulnerable groups in society. It is argued that if prices of necessities rise faster than those of other commodities, the poor will be more adversely affected than the non-poor. Moreover, inflation is considered costly for the poor since their purchasing power is eroded especially for those groups whose earnings are fixed in nominal terms, their assets are devalued more as they hold a larger share of their assets in liquid form compared with the non-poor, and it is difficult for the poor to hedge against inflation due to their limited access to the financial system (see Mujeri 2008, Mujeri and Mortaza 2008).

The fight against inflation requires effective policies to control undue rise in prices and credible tools to predict future movements in inflation and identify the causal factors. In this respect, among alternative traditions, the P-star (P*) class of models has gained popularity in recent years especially due to its close link with the long tradition of mainstream monetary theory and firm roots in the so-called quantity theory of money. After the model was first developed by Hallman et. al. (1989, 1991), a wide body of literature has emerged on its theoretical underpinnings and empirical validation across different countries (see, for example, Christano 1989, Hallman and Anderson 1993). On the other hand, several researchers have criticized the approach for its unreasonable dynamic behavior and low predictability in forecasting performance (see, Pecchenino and Rasche 1990, Hoeller and Poret 1991, Funke and Hall 1994, Hall and Milne 1994). Although the P* model has widely been applied to major developed countries, its application to less developed countries remains somewhat limited. [1]

In this paper, we have tested the performance of the P* model for the Bangladesh economy. For the purpose, the model has been estimated with annual data covering the period 1980-2008 and its forecasting ability tested through generating recursive forecasts. Our results lend some ground for confidence in the use of P* models for gauging inflationary pressures in the Bangladesh economy.

The paper is organized as follows. After the introductory remarks of this section, Section II provides a short review of the empirical estimation of the P* models in different countries and the credibility of the model as inflation forecaster in these countries. The basic methodological underpinning of the P* model is set out in Section III while the details of the empirical estimation of the model with Bangladesh data and the discussion of the results are the subject matter of Section IV. The forecasting performance of the estimated model is discussed in Section V. Finally, some conclusions and policy implications are spelt out in Section VI.

II. REVIEW OF P* MODEL ESTIMATION

Despite the fact that the class of P* models had a very recent origin, the model has been widely applied to many countries. The model, with suitable modifications, was applied in the 1990s to UK (Allen and Hall 1990), France (Bordes et. al. 1992), Germany (Tödter and Reimers 1994), and other OECD countries (Hoeller and Poret 1991). Tatom's (1992) application of P* model to Austria reports problems in identifying permanent shocks to potential output and/or velocity leading to the rejection of such models. Although he finds evidence of a long-run relationship between Austrian inflation and money growth, the tests favor rejection of even the first-difference version of the model. Since Austria has a small economy closely tied to Germany, Tatom also investigates the hypothesis as

[1] Using the P* model as an indicator of inflation, central banks in countries like Australia, Canada, Finland, New Zealand, Spain, Sweden, and UK have shifted their approach to monetary policy to publicly announced inflation targeting in the 1990s.

to whether Austrian prices are linked to German P* measure which is also rejected. He, however, reports statistically significant long-run relationship between Austrian and German inflation.

Hoeller and Poret (1991), while examining the inflationary pressure in OECD countries, computed trend values of output and velocity by using linear time trend and statistical filters. They find that P* equations outperform other simple financial market based inflation models for most countries although P* equations are found less satisfactory for the purpose of short term inflation forecasting. Except for the US and Germany, the P* forecasts were clearly dominated by predictions from an auto-regressive model. They held that the poor forecasting performance of the P* approach is mainly due to the difficulty in making *ex ante* discrimination between transitory and permanent changes in both trend velocity and potential output.

For Germany, Tödter and Reimers (1994) have defined the equilibrium price level *P** as money per unit of potential output at equilibrium velocity. In this case, deviation between *P** and the actual price level (the price gap) is taken as an indicator of future price movements. To measure equilibrium velocity, the authors integrate a long-run money demand function into the *P** approach. The empirical results of the model provide a stable link between M3 and the price level. The authors, however, fail to establish comparable evidence for Ml and M2. Svensson (1999), on the other hand, argues that the P* model implies that inflation is determined by the level of and changes in the 'real money gap' which is an important indicator for future inflation. He finds that the P* model does not provide any rationale for either a Bundesbank-style money growth target or a Euro system-style money growth indicator. While examining the relationship between money and prices in Estonia over 1997 to 2003 by applying the P* model, Dabusnskas (2005) finds that the money gap measures dominate the output gap in explaining inflation in the short run.

The application of the P* approach to less developed countries is somewhat limited. In India, Nachane and Lakshmi (2002) have estimated the P* model using both annual and quarterly data for the period 1955 to 1990. The results of their analysis show that income velocity of money is trend stationary and the velocity gap models fare consistently better providing support for a distinctly monetarist explanation of inflation. They report a perverse sign for the output gap for both annual and quarterly versions of the model along with relatively poor forecasting performance of the output and price gap models. Nevertheless, their view contradicts several other studies favoring structuralist explanations of inflation in India (e.g. Bhattacharya 1984, Balakrishnan et. al.1994).

Qayyum and Bilquees (2005) have used the P* model for Pakistan to calculate the leading indicator of inflation. They also test the forecasting performance of the P* model the results of which show that compared with simple autoregressive model and

M2 growth augmented model, the P* model performs better mainly due to its use of additional information about future inflation.

III. METHODOLOGICAL CONSIDERATIONS

The central objective of the P* approach to modeling is to assess the inflationary potential of an economy through evaluating the characteristics of its monetary conditions. The P* type of models is based on the well-known classical quantity theory of money and combines long term determinants of the price level with short term changes in inflation in the economy. In this section, we develop the P* model which closely follows the methodology provided by Hallman et. al. (1989, 1991).

As mentioned earlier, the P* concept is part of the general quantity theory of money whose basic identity is the equation of exchange. We, therefore, start with the famous equation of exchange which gives a relationship between the price level and the quantity of money in the economy:

$$MV = PY \qquad \text{(1)}$$

where M is a suitable measure of money supply, V is the (income) velocity of money, P is the aggregate price level, and Y is real GDP. For the present analysis, we assume that real GDP fluctuates around its potential level and there exists an equilibrium level for the income velocity of money.[2]

The model links the behavior of P to growth in M adopting the hypothesis that Y fluctuates around its potential value while V has an equilibrium level that is independent of time and tracks the long run. Let us define Y* as the potential real GDP and V* as the equilibrium level of velocity. Then equilibrium price level P* may be defined, for a given level of money supply M, as:

$$P^* = MV^*/Y^* \qquad \text{(2)}$$

Taking logarithm and using lower case letters to denote the transformation, equations (1) and (2) can be written as:

$$m + v = p + y \qquad \text{(3)}$$

[2] The empirical existence of equilibrium velocity (v*), however, is marred with controversy. In their pioneering work, Hallman et. al. (1989) considered that the velocity of M2 was stationary for the US indicating that shocks on velocity represent transitory shocks on the level of the velocity. On the other hand, such an assertion has been refuted in several other countries. See, for example, Bordes et. al 1992, Tödter and Reimers 1994, Atta-Mensah 1996.

$m + v^* = p^* + y^*$ ……………………………………………………………………………… (4)

From (3) and (4), it follows that

$(p – p^*) = (v – v^*) – (y – y^*)$ ………………………………………………………………… (5)

In equation (5), $(p – p^*)$ provides the 'price gap' whereas $(v – v^*)$ and $(y – y^*)$ measure 'velocity gap' and 'output gap' respectively.

Following the postulates of the P* model, we assume that actual price level (P) has a tendency to move toward the equilibrium price level (P*) and the difference between the actual and the equilibrium price level acts as a good predictor of inflation. The implication is that if, at any point in time, the quantity of money (M) supports a price level which is above its equilibrium value, then it is either driving actual velocity above its equilibrium level or depressing actual output below its potential level or both.

Thus equation (5) can be used to predict the movements of the rate of inflation. For instance, if actual inflation (π) exceeds $(p – p^*)$ which may be taken as the equilibrium inflation rate (π^*), then π should fall until it reaches π^*. The movements would be reversed in case $\pi < \pi^*$. Therefore, the difference between p and p* provides a leading indicator for future acceleration or deceleration of inflation.

Estimation of y* and v*

The critical issue in the present analysis is the estimation of potential GDP (y*) and equilibrium velocity (v*). In estimating y*, alternative techniques are available using differing assumptions. One approach used by the researchers is to combine a Phillips curve based estimate of the natural unemployment rate with Okun's law.[3] Another alternative is to use a linear time trend to calculate the potential output (see Christiano 1989, Hannah and James 1989, Hallman et. al. 1991, Nachane and Laxmi 2002). A widely used statistical method is to apply the filter approach for which two versions are available—the Kalman filter and the Hodrick-Prescott (H-P) filter (see Hodrick and Prescott 1980).[4] The H-P filter, for instance, is a two-sided linear filter that uses a long term symmetric moving average to de-trend the time series.

For our purpose, we have tested two alternative methods for estimating y* and v*. Under the first method, we define the equilibrium velocity (v*) such that v(t) is stationary (that is integrated of order zero, I(0)) either around a constant v_0 or around a linear trend. If v(t) is considered stationary around $v_{0,}$ the equilibrium velocity (in log

[3] For application of this approach to estimating Y*, see Ebrill and Fries 1990, Pacchenino and Rasche 1990.
[4] For examples of using Kalman filter, see Bomhoff 1990, Kuttner 1992, Fisher and Fleissing 1995. The H-P filter has been used, among others, by Hoeller and Paret 1991, Gibbs 1995, McMarrow and Poeger 2001, Qayyum and Bilquees 2005.

form) would be taken as $v^* = v_0$. In the second case, v* would be taken as variant over time with $v^*(t) = \gamma + \varphi t$. The potential GDP (y*) would also be measured using a similar methodology through examining if y is I(0). For the second method, we would adopt H-P filter approach to estimate y* and v*. More specifically, the method adopts constrained minimization of the sum of the squared deviations of y (or v) from its trend. For instance, in the case of potential GDP, H-P method chooses y* (lnY*) that minimizes

$$\Sigma(\ln Y_t - \ln Y_t^*)^2 + \lambda\, \Sigma[(\ln Y_{t+1} - \ln Y_t^*)-(\ln Y_t^* - \ln Y_{t-1})]^2 \quad \text{.................................(6)}$$

where λ is Lagrange multiplier and y* approaches a linear trend as $\lambda \rightarrow \infty$. As practiced in the literature, we set $\lambda = 100$ keeping in view the fact that a smaller λ implies shorter cycles and smaller output gaps.

Cointegration of p and p*

Once y* and v* are estimated, p* can be calculated through substitution in equation (5). For our purpose, it is important to test whether p and p* are cointegrated (Engle and Granger 1987, Johansen and Juselius 1990). If the two are cointegrated, then $p^* > p$ would indicate accelerating inflation in future while $p^* < p$ would imply deceleration of future inflation. Moreover, the above cointegration would indicate stationarity of velocity and output gaps since equation (5) involves a linear relationship.

Specifying inflation dynamics

The price gap as given in equation (5) does not contain any dynamics of adjustment of p toward p*. For the purpose, we follow Hallman et al (1991) who generalized the approach of McCallum (1980) and Mussa (1981). For our purpose, we adopt the following general specification:

$$\Delta\pi(t) = \alpha + \beta\,(v_{t-1} - v^*_{t-1}) + \gamma\,(y_{t-1} - y^*_{t-1}) + \delta\,\pi_{t-1} + \sum_{j=1}^{q} \phi_j \Delta\pi_{t-j} + \varepsilon_t \quad \text{........................ (7)}$$

In equation (7), Δ is the one period difference operator, q is the chosen lag, and ε_t satisfies white noise properties. The above formulation (7) allows the possibility of velocity and output gaps impinging differently on changes in inflation while the inclusion of lagged inflation permits incomplete adjustments.

However, while using in specific contexts, it may be desirable to use specific versions of the above model in order to derive more robust results. One possibility, while estimating equation (7), is to test the hypothesis $\delta = 0$ using Dickey-Fuller non-standard statistic τ_μ (Dickey and Fuller 1979, Fuller 1976, Hallman et. al. 1991). Based on the test outcome, two versions of model (7) arise.

In case δ is not significant, equation (7) can be reformulated as follows:

$$\Delta\pi(t) = \alpha + \beta\,(v_{t-1} - v^*_{t-1}) + \gamma\,(y_{t-1} - y^*_{t-1}) + \sum_{j=1}^{q} \phi_j \Delta\pi_{t-j} + \varepsilon_t \quad \text{.................................... (8)}$$

On the other hand, if δ is significant, the estimating model becomes:

$$\pi(t) = \alpha + \beta\,(v_{t-1} - v^*_{t-1}) + \gamma\,(y_{t-1} - y^*_{t-1}) + \sum_{j=1}^{q} \phi_j \Delta\pi_{t-j} + \varepsilon_t \quad \ldots\ldots\ldots\ldots\ldots\ldots (9)$$

The above equations allow the possibility of inclusion of different lags in estimation with both annual and quarterly data.

IV. EMPIRICAL ESTIMATION AND RESULTS

The data

The class of models underlying the P* approach is considered more suitable to generate short term forecasts of inflation making these models more appealing to the policy makers. Obviously, the relevance of the model output to reality depends on quality of the data used for estimating the model. For the present model, we have used published data from different sources. A major limitation, however, is that the model could not be estimated using quarterly data due to lack of availability of quarterly estimates of GDP. Data on the annual real GDP series (Y) were taken from the national accounts statistics published by the Bangladesh Bureau of Statistics (BBS).

Changes in the aggregate price level (P) have been measured by the twelve-month average national consumer price index (CPI) inflation computed by BBS on a monthly basis. Since inflation figures for earlier years are not available for the currently used 1995/96 base, these were appropriately adjusted in order to bring consistency in the data. The money supply measure selected is broad money (M2) for which data are available from the Bangladesh Bank over the entire period of analysis. The income velocity of money (V) has been calculated as the ratio of nominal GDP to money supply where money supply is the quantity of broad money (M2). The annual model was estimated with time series data covering the financial years (FY) FY80 to FY08.

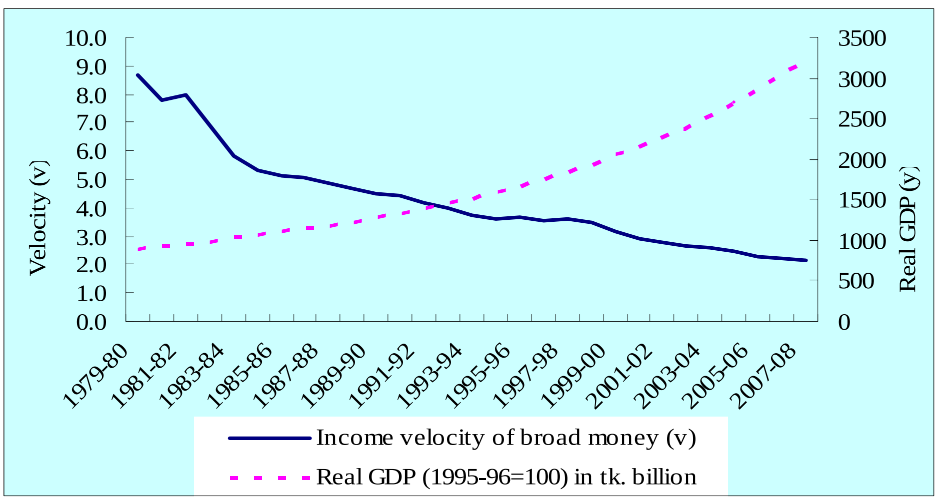
Estimation of equilibrium velocity and potential output

As mentioned above, one of the basic tenet of the P* approach is the existence of an equilibrium level for the velocity of money (v^*) toward which the current velocity (v_t) converges in the long run. Therefore, one of the first steps in our empirical analysis is to estimate v^*. For the purpose, the first task is to test whether v is I(0) by conducting unit root tests of the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) varieties through applying a sequential procedure (Holden and Perron 1994). Accordingly, the following equation was first estimated:

$$\Delta v(t) = \alpha + \beta t + \rho v(t-1) + \sum_{j=1}^{q} \Delta\pi_{t-j} + \varepsilon(t) \quad \ldots\ldots\ldots\ldots\ldots\ldots (10)$$

with q chosen large enough to purge $\varepsilon(t)$ of serial correlation. A value of q=3 was found adequate for the purpose. In order to test the joint hypothesis β=0, ρ=0, we used the ϕ_3 statistic for the ADF test and the $Z(\phi_3)$ statistic for the PP test. It has, however, been argued that DF types of tests have low power to reject the null hypothesis of unit roots against stable autoregressive series with roots near unity in finite samples (Kwiatkowski et. al 1992). Therefore, we have also conducted the KPSS test which stipulates the null hypothesis of stationarity against the alternative of a unit root. The results of testing the stationarity assumption for v (and y) are given in Appendix Table I. From the tests, stationarity is rejected for v (as well as y) indicating the absence of a unit root. This can also be visualized from the yearly movements of v (and y) during the sample period (Figure I). It appears that the rapid progress in financial deepening and innovations that have characterized the Bangladesh economy especially since the 1990s have produced non-stationarity in the velocity data. A similar explanation of structural transformation and rapid changes in sectoral composition of output may be applied for the output data. Hence we use the earlier mentioned Hodrick-Prescott (H-P) detrending method which considers the stochastic tendencies in a series (see King and Rebelo 1993). Since output (y) is also a I(1) series in our case, the estimation of y* followed a similar approach.

Figure 1: Movements in V and Y, 1980-2008



Source: Bangladesh Bureau of Statistics and Bangladesh Bank

Estimation of cointegration and causality

The next empirical issue to test is the presence of any tendency for the actual price (p) to gravitate toward its long run equilibrium value (p*). Since only two variables are involved, we invoke the Engle-Granger (1987) methodology.[5] The results are presented in Appendix Table 2 which affirm the existence of cointegration of p and p*.

[5] The procedure works as follows. First, a cointegrating regression is performed which is a simple OLS regression between p and p*. The Durbin-Watson (DW) statistic of this regression yields Cointegrating Regression DW (CRDW) statistic. In the next step, the residuals of the above regression are tested for unit roots with Dickey-Fuller (DF) or Augmented DF (ADF) approach. The critical values, however, are not the usual ones and these have been specifically constructed by Engle and Woo (1987).

Before actual estimation of the inflation dynamics model, it is appropriate to decide whether we should specify the right hand side of the general model (7) in level form $\pi(t)$ or in changed form $\Delta\pi(t)$. As mentioned in Section III, this involves estimation of equation (7) followed by a significance test on δ (the coefficient of π_{t-1}) using the Wald test. The results are presented in Appendix Table 2 which show that the hypothesis $\delta=0$ can be rejected so that equation (9) is the appropriate prototype model in our case.

Finally, one of the key assumptions that is implicit in the P* approach is the postulate inherent in the monetarist view that money supply affects the price level. It is important therefore to reveal such causality prior to estimating the model with data considered valid for the analysis. [6] We have examined the causality issue following Granger (1969, 1988) through testing for the null hypothesis of non-causality from M to P using the following equation:

$$\Delta P(t) = \alpha_0 + \sum_{j=1}^{q} \alpha_{1j}\, \Delta P_{t-j} + \sum_{j=1}^{q} \alpha_{2j}\, \Delta M_{t-j} + \varepsilon_t \quad \ldots\ldots (11)$$

We also tested an equivalent expression for M not being caused by P. The results are given in Appendix Table 3 which show that the causality between money and prices works in the direction expected by the P* approach given the rejection of the null hypothesis of non-causality from the monetary aggregates to prices in Bangladesh.[7]

Model estimation and results

Since equation (9) turns out to be the appropriate model in our case, we have provided the results using both inflation level and inflation change formats to facilitate comparisons since the original P* model has inflation change specification.

Before discussing the estimated results, a few observations may be made. As we have noted earlier, the price gap (p – p*) in the model consists of two components: the velocity gap (v – v*) and the output gap (y – y*). A positive velocity gap (v > v*) at any point of time reflects a situation in which current velocity is higher than the equilibrium velocity and hence the tendency for v is to fall. On the other hand, a negative velocity gap is analogous to a 'liquidity surplus' situation when the expectation is that the velocity would rise toward its equilibrium level. In the case of the output gap, a positive value (y > y*) leads to the expectation that output would fall to its equilibrium level while the situation would be reversed in the case when the gap is negative. Thus, if equation (9) is the chosen model, both a positive velocity gap and a negative output gap would lead to a decline in the rate of inflation while, under such situations, inflation rate

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[6] It is argued that the validity of the P* approach to represent the data should not be taken as synonymous to validity of the underlying assumption of the model. See Hall and Milne 1994.

[7] The conclusions from the results from such causality tests, however, should be treated with caution. For instance, it has been argued that the results are often quite sensitive to the choice of the optimal time lag. See Hansen 1989.

would decelerate if inflation change (equation 8) is the appropriate model. The underlying mechanisms of bringing about the change are, however, different. While the aggregate demand curve undergoes a shift to the left in the former case, the aggregate supply curve shifts to the right in the case of the latter. Thus, one would expect a negative sign for the estimated coefficient of the velocity gap while the a priori sign is positive in the case of the output gap.

The results of empirical estimation are presented in Table 1 for both level (Model 1) and change (Model 2) forms of inflation. In both models, estimated coefficients of velocity and output gaps have the expected signs and the P* model does not seem to fail in the Bangladesh context. In Model 1, coefficients of both velocity and output gaps have correct signs and are significant (at 10 percent level). It may be noted that the absolute value of the coefficient of the output gap (1.006) is appreciably higher than that (0.205) of the velocity gap. In this context, the above results may be interpreted in terms of neo-classical explanations of inflation which largely run in terms of the output gap while the monetarist approach focuses on the velocity gap. In the case of Model 2, although both the coefficients have the right signs, the output gap coefficient turns out to be insignificant. Thus the above results suggest that both velocity and output gaps are important determinants of changes in inflation in Bangladesh.

Table 1: P* Models of Inflation Dynamics

Variable	Model 1 (Dependent variable π)	Model 2 (Dependent variable $\Delta\pi$)
v_{t-1}-v^*_{t-1}	-0.2052	-0.2623
	(-1.8495)	(-1.8464)
y_{t-1}-y^*_{t-1}	1.0055	0.2158
	(1.8956)	(0.3498)
π_{t-1}	-	-0.5266
		(-2.0798)
$\Delta\pi_{t-1}$	0.5656	0.3153
	(2.5109)	(1.1588)
$\Delta\pi_{t-2}$	-0.1920	-0.2521
	(-1.1961)	(-0.9381)
$\Delta\pi_{t-3}$	0.1115	0.2805
	(0.5511)	(0.9972)
$\Delta\pi_{t-4}$	-	0.2439
		(1.0211)
$\Delta\pi_{t-5}$	-	0.2251
		(1.1106)
α	0.0645	0.0341
	(13.2271)	(1.9435)
Adj. R-squared	0.4493	0.4729
Darbin-Watson	1.5387	2.2496

Note: Figures in parentheses show the t-statistic.
Source: Authors' calculations.

In order to examine the relative significance of velocity and output gaps further, we have estimated the following two 'specialized' versions of Model 1:

(i) Velocity gap model (γ=0 in equation 9)
(ii) Output gap model (β=0 in equation 9)

The results are presented in Table 2. Both velocity and output gaps turn out to be highly significant with their expected signs. This reaffirms the earlier conclusion regarding the importance of both monetary and real factors in the inflationary process in Bangladesh.

Table 2: Velocity and Output Gap Models

Variable	Version 1 (Velocity gap model)	Version 2 (Output gap model)
v_{t-1}-v^*_{t-1}	0.2795	-
	(2.5263)	
y_{t-1}-y^*_{t-1}	-	1.3522
		(2.5665)
$\Delta\pi_{t-1}$	0.5660	0.6161
	(2.3597)	-2.5952
$\Delta\pi_{t-2}$	-0.1975	-0.2367
	(1.1940)	(1.4061)
$\Delta\pi_{t-3}$	0.1746	0.0985
	(0.8212)	(0.4591)
α	0.0618	0.0671
	(12.4123)	(13.5974)
Adj. R-squared	0.3742	0.3792
Darbin-Watson	1.223	1.440

Note: Figures in parentheses show the t-statistic.
Source: Authors' calculations.

V. FORECASTING PERFORMANCE

The accuracy of out-of-sample forecasts of a model is an important indicator of its robustness. This section provides an assessment of the out-of-sample performance of our estimated models. For the purpose, we consider the following models:

Model I: Price gap model (Equation 9)
Model II: Velocity gap model (γ=0 in equation 9)
Model III: Output gap model (β=0 in equation 9)

However, in order to gauge the forecasting performance of the models, we need a benchmark model for comparison for which we have chosen the following AR(2) model based on AIC which is referred to as Model IV [8]:

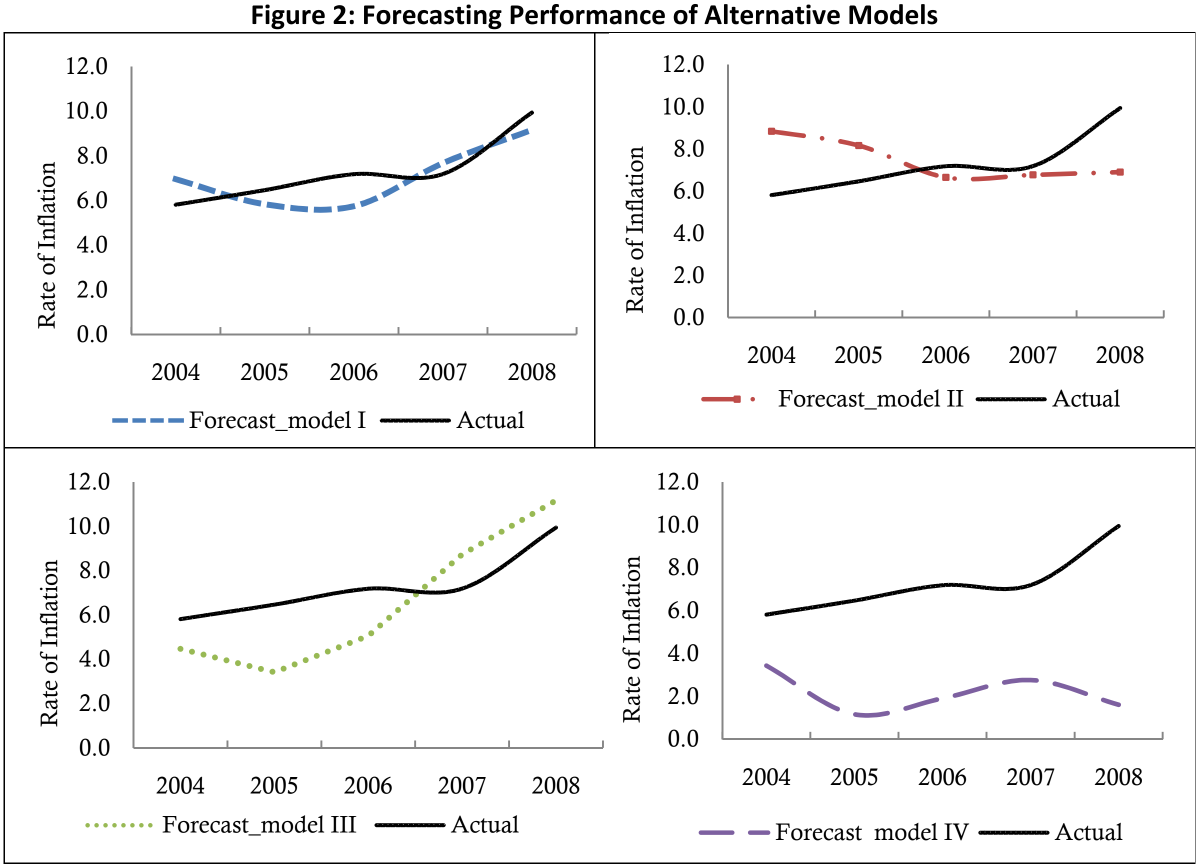
[8] See Box and Jenkins (1970) for details on such models.

$$\Delta\pi_t = -0.0008 - 0.0125\,\Delta\pi_{t-1} - 0.5832\,\Delta\pi_{t-2}, \quad \text{adjusted } R^2 = 0.36 \quad .... (12)$$
$$(-0.175) \quad (-0.079) \qquad (-3.925)$$

(Figures in parentheses denote t values)

The models I-IV are estimated over the entire sample. To evaluate the out-of-sample performance, we have re-estimated each model recursively, beginning with the year 2004 to 2008 introducing successively a new year at each recursion. The one year ahead forecasts made at each simulation are then noted and compared with corresponding actual value.[9] The out of sample forecasts have been generated from 2004 to 2008 (Figure 2). This shows that the price gap model performs better followed by the output gap model and the velocity gap model. The forecasting performance of the benchmark model, however, does not turn out to be satisfactory.

Figure 2: Forecasting Performance of Alternative Models



Source: Authors' calculations.

[9] The methodology is as follows. At first recursion, based on the coefficients of the model estimated over 1980-2008, a one-year ahead forecast is made for 2004. The model is then re-estimated over 1980-2004 and using the revised set of coefficients, a forecast is made for 2005 and so on.

The relative merits of each model have been judged using two criteria. The first is the well known root mean square error (RMSE) criterion and the second is the Chong-Hendry (1986) 'forecast encompassing' criterion. The RMSE for any model is calculated as:

$$RMSE = [\ 1/N \sum e(t)^2\ ]^{½} \quad \text{.......} \quad (13)$$

where e(t) is the forecast error at time t and N is the total number of periods over which the forecasts are being compared.

Table 3 presents the RMSE for the four models used in this forecasting exercise. The results bring out two important points. *First,* the forecasting performance of the benchmark model (model IV) is significantly inferior to the performance of the three alternative models. *Second,* the price gap model (model I) outperforms the velocity and output gap models while the output gap model marginally outclasses the velocity gap model. These results are consistent with our earlier observations.

Table 3: RMSE Comparison of Alternative Models

Estimation Period	Forecasting Period	h-step	RMSE				Ratio of RMSE				
			I	II	III	IV	I/IV	II/IV	III/IV	I/II	I/III
1980-2003	2004-2008	5	0.0090	0.0194	0.0186	0.0514	0.1751	0.3774	0.3619	0.4639	0.4839
1980-2004	2005-2008	4	0.0085	0.0164	0.0198	0.0574	0.1481	0.2857	0.3449	0.5183	0.4293
1980-2005	2006-2008	3	0.0092	0.0166	0.0156	0.0593	0.1551	0.2799	0.2631	0.5542	0.5897
1980-2006	2007-2008	2	0.0059	0.0200	0.0128	0.0632	0.0934	0.3165	0.2025	0.2950	0.4609
1980-2007	2008-2008	1	0.0071	0.0280	0.0111	0.0789	0.0900	0.3549	0.1407	0.2536	0.6396

Source: Authors' calculations.

According to the second criterion, if we are interested in comparing the forecasting performance of two models A and B, then we need to compute the forecast errors of the models at time t, $e(t)^A$ and $e(t)^B$. Now if $g(t)^A$ and $g(t)^B$ are the forecasts of the two models, we perform the following regressions:

$$e(t)^A = \delta\ [g(t)^B - g(t)^A\ ] + \varepsilon_t \quad \text{.......} \quad (14)$$

$$e(t)^B = \rho\ [g(t)^A - g(t)^B\ ] + \eta_t \quad \text{.......} \quad (15)$$

Now if t_δ and t_ρ denote the t values of δ and ρ respectively and if the former is not significant but the latter is, then model A 'forecast encompasses' model B. The interpretation is that model A contains useful information beyond that contained in the forecasts generated by model B.

The results are given in Table 4. These show that (i) the price gap model (model I) forecast encompasses the benchmark AR model (model IV); (ii) the velocity gap model (model II) forecast encompasses the benchmark AR model; (iii) the output gap model (model III) forecast encompasses the benchmark AR models and the velocity gap model

(model II) though at close to 10 percent level; (IV) the AR model is forecast encompassed by all three models.

Table 4: Forecast Encompassing Comparisons

Model	I	II	III	IV
I	--	0.699 (0.490)	1.431 (0.166)	-0.128 (0.899)
II	0.876 (0.390)	--	1.709 (0.101)	0.966 (0.345)
III	-0.025 (0.979)	0.581 (0.567)	--	-1.362 (0.188)
IV	2.089 (0.048)	1.946 (0.064)	2.224 (0.030)	--

Note: The entry (i,j) gives the t value corresponding to A being model i and B being model j. Figures in parentheses are p values.

Source: Authors' calculations.

VI. CONCLUSIONS AND POLICY IMPLICATIONS

The P* model, which puts together long run determinants of the price level based on the classical quantity theory of money and short term changes in current inflation, has gained wide popularity in recent years to explain inflation dynamics especially in developed countries. Empirical studies have also lent support in favor of the model in different countries. In this paper, we have applied the P* approach to Bangladesh. The model was estimated with annual data and its forecasting performance tested applying alternative methodologies.

The results show that the model performs relatively well with annual data and the model contains additional information regarding the future rates of inflation. The price and output gap models fare consistently better than the velocity gap model which brings out the importance of non-monetary factors in explaining inflation dynamics in Bangladesh. These findings have implications for the monetary targeting policy framework currently pursued by the Bangladesh Bank.

The design of the P* models, however, puts emphasis on short term inflation forecasting so that the use of more frequent (e.g. quarterly or monthly) data is called for. Unfortunately, such an exercise could not be undertaken for Bangladesh due to paucity of relevant data. While availability of reliable data, especially on quarterly estimates of actual and potential GDP, would help in expanding the model and deriving more robust results, its usefulness in policy analysis can be enhanced through incorporating other relevant indicators including growth in credit and interest rate spreads.

The P* models can have wide applications in policy analysis. The model can be used in national income analysis and in deriving macroeconomic implications of different monetary policy rules through estimating impulse response functions. Since income stabilization has not been an important concern of the monetary authorities in Bangladesh in the past within a policy environment where monetary policy could not play its desired role, we have kept such an analysis outside the purview of the present paper. However, with financial sector liberalization and reforms and greater autonomy for the Bangladesh Bank (the central bank), it is likely that the policy relevance of the P* type of models will be ramified in the country. This would create the scope for the P* model to play a more proactive role in formulating monetary policy in Bangladesh.

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APPENDIX

Appendix Table 1

Unit Root Tests for v(t) and y(t)

Variable	Φ_3	$z(\Phi_3)$	KPSS
v(t)	-2.59	-2.75	0.12
y(t)	0.22	0.36	0.19
5% critical value	-3.58	-3.58	0.15
10% critical value	-3.22	-3.22	0.12

Source: Authors' calculations.

Appendix Table 2
Test of Cointegration of P and P*

a. Estimation of the long run equation

Dependent variable p

Variable	Coefficient	t-statistic	Prob.
p*	1.0031	57.8170	0.0000
Constant	-0.0154	-0.1999	0.8430

Adj. R-squared 0.992 Durbin-Watson stat 0.834

b. Unit root tests for ε_t

Variable	Φ_3
ε_t	-3.0964
5% critical value	-1.9539
10% critical value	-1.6096

c. Equilibrium correction model

Dependent variable Δp

Variable	Coefficient	t-statistic	Prob.
Δp*	0.3384	2.9665	0.0065
ε_{t-1}	-0.2735	-2.5187	0.0185
Constant	0.0463	4.7779	0.0001

Adj. R-squared 0.274 Durbin-Watson stat 1.23

b. Significance test for δ in equation 7

Test statistic	Value	df	Prob.
F-statistic	4.3258	(1, 13)	0.0579
Chi-square	4.3258	1	0.0375

Source: Authors' calculations.

Appendix Table 3

Granger Causality Tests between P and M

Null hypothesis	F-statistic	Prob.
P does not Granger cause M	0.4628	0.6355
M does not Granger cause P	9.1625	0.0013

Source: Authors' calculations.