A Time-Varying Parameter Model of Inflation in India[†]

Sudhanshu Kumar[††]
Indira Gandhi Institute of Development Research, Mumbai

Naveen Srinivasan
Indira Gandhi Institute of Development Research, Mumbai

M. Ramachandran
Pondicherry University

Abstract

For much of the 1970s and 1980s, India experienced recurrent bouts of high inflation together with sub-par economic performance. In contrast her inflation record over the past decade or so has been far better. What explains this turnaround? Standard models of central bank optimization predict that the central bank's preference parameters and the slope of the Phillips curve are key determinant of inflation dynamics. Hence, time variation of these parameters should be reflected in changes in inflation dynamics. Therefore, a time-varying parameter model for inflation is proposed and is estimated using the median-unbiased estimator. The estimated time paths of the reaction function coefficients suggest that while monetary policy and structural change have played a non-trivial role in inflation reduction, good luck and exchange rate regime played a primary role.

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[††] Corresponding Author: Indira Gandhi Institute of Development Research, Gen. A.K. Vaidya Marg, Goregaon (East), Mumbai 400065, INDIA. Email: sudhanshu@igidr.ac.in

1. Introduction

One of the most striking events of the past two decades has been the remarkable decline in inflation in both developed and developing countries, in sharp contrast to the period immediately preceding it.[1] Interestingly, the behaviour of inflation in India broadly exhibits such a pattern. For much of the 1970s and 1980s, India experienced recurrent bouts of high inflation together with sub-par economic performance. Since the 1990s not only has growth picked up but there has also been a marked decline in mean inflation (see Figures 1 and 2).[2] What explains this turnaround? And more importantly, will this progress on inflation be sustained, or is the recent improvement only a temporary lull?

There are several plausible explanations for this turnaround. In a nutshell, there are four main candidates: good policy, structural change, good luck and exchange rate regime. The suggestion that the moderation in inflation is down to "good policy" argues that the main contrast between the two broad periods of inflation experience in Figures 1 and 2 can be attributed to a significant

[1] Most empirical and theoretical studies on inflation dynamics have concentrated on developed economies. Our study pays attention to the international dimension of the issue.

[2] Figure 1 plots quarterly WPI (all commodities) inflation (year-on-year percentage change) for the period 1955Q1-2008Q1. Visual shifts in mean inflation can be shown more formally by a rolling 10-year regression of inflation (π_t) on a constant (c) in the form: $\pi_t = c + \varepsilon_t$, where ε_t is a random disturbance term. The estimated constant, $\hat{c}$, along with the 5% critical value are plotted in Figure 2. The critical value shown in the figure is 1.96*SE, where SE is the standard error of the constant estimated from the rolling regressions.

change in central bank responses to inflation. Specifically, it suggests that during periods in which the central bank's resolve to stabilize inflation is strong (less relative weight on output stabilization), inflation would fall.

In contrast, the structural change explanation puts weight on the idea that fundamental reforms to product and labour markets as a result of globalization and trade openness, has increased the flexibility of economies (altered the slope of the Phillips curve). This, in turn, makes it less likely that a given shock to costs or to demand results in inflationary pressure.

The third explanation for the more recent experience of low inflation emphasises "good luck". According to this view economies have not in more recent times been subjected to too many inflationary cost shocks of the kind that we saw in the 1970s . There is a substantial reduction in both the mean and volatility of supply shocks. This, so the story goes, has diminished the challenges faced by policymaker's charged with controlling inflation.[3]

Finally, an alternative interpretation especially relevant for emerging market economies (EME's) such as India focuses on the nature of the exchange rate

[3] However, note that these explanations are unlikely to be independent of each other. In particular, if monetary policy changes have credibly lowered and stabilized inflation, the resulting drop in expected inflation may have given central bankers more flexibility to respond to adverse supply shocks (see Bean, 2006). Consequently, it becomes difficult to disentangle the good luck hypothesis from the theory that monetary policy has more effectively responded to shocks. Similarly, the structural change explanation is not independent of monetary policy. The slope of the Phillips curve depends on the underlying monetary policy framework among other things (see Lucas, 1973 and Ball et al., 1988).

regime. Specifically, if the objective of the central bank is to curb exchange rate volatility, then an increase (or decrease) in U.S. inflation forces domestic monetary authorities to allow domestic inflation to rise (or fall), so as to avoid a depreciation (or appreciation) of the domestic currency (see Doyle and Falk, 2008). As a result inflation may have spilled over from the U.S. to India due to dislike on the part of the Reserve Bank of India (RBI) of large nominal exchange rate movements. Given that the current currency regime in India betrays symptoms of pegging to the U.S. dollar (see Calvo and Reinhart (2002) and Reinhart and Rogoff (2002)), it would seem natural to test whether the spillover hypothesis has any empirical relevance in the Indian context.[4]

To this end, we construct a reduced form econometric model for inflation with drifting coefficients that encompasses all the candidate explanations described above. While our model, by its very nature, does not allow one to uncover the deep structural sources behind the moderation, we believe it can still be helpful at ruling out some hypotheses and shedding light on the relative merits of alternative explanations, while imposing a minimal structure. In this model the central bank's preference parameters (namely, the

[4] In fact, exchange-rate targeting has been successfully used in the past to control inflation in industrialized countries too. Both France and the United Kingdom, for example, successfully used exchange-rate targeting to lower inflation by tying the value of their currencies to the German mark (see Mishkin, 1999). Among emerging market countries an important recent example has been Argentina, which in 1990 established a currency board arrangement, requiring the central bank to exchange U.S. dollars for new pesos at a fixed exchange rate of 1 to 1. Inflation which had been running at over a one-thousand percent annual rate in 1989 and 1990 fell to under 5% by the end of 1994, and economic growth was rapid, averaging almost 8% at an annual rate from 1991 to 1994.

relative weight placed on output and exchange rate stabilization) and the slope of the Phillips curve are key determinants of inflation dynamics. Hence, time variation of these parameters should be reflected in changes in inflation dynamics.

Therefore, a time-varying parameter (TVP) model for inflation is proposed. The time variation is modelled as driftless random walks, and is estimated using the median-unbiased estimator proposed by Stock and Watson (1998). The estimated time paths of the reaction function coefficients suggest gradual changes in the rule coefficients, unlikely to be captured by the usual split-sample estimation. Our findings ascribe to monetary policy and structural change explanation quite a modest role in the moderation of inflation. In sum, while monetary policy and structural change have played a non-trivial role in moderating inflation, exchange rate regime along with favourable supply-side developments (a reduction in both the mean and volatility of shocks), seem to have played the dominant role.

The paper is organized as follows. Section 2 outlines the framework linking the dynamics of inflation to the preferences and optimization problem of the central bank. We also provide the intuition and the main ideas necessary to interpret the quantitative results. Section 3 describes the data, the estimation method, results and discusses policy implications. Section 4 provides concluding remarks.

2. Theoretical Framework

The theoretical framework consists of a stylized model in which the central bank aims to minimize a quadratic loss function with inflation, the output gap and nominal exchange rate as arguments. Specifically, we extend the textbook time-inconsistency model (policymakers' target for output growth is greater than its trend growth rate) by also allowing the policymakers' target for output growth to be a nonlinear function of the underlying supply shocks.[5] We assume that supply shocks follow a general autoregressive process (with drift). Finally, the model also allows for the possibility that inflation in one country is transmitted to another country via the effect on exchange rates by incorporating exchange rates explicitly in the policymakers' objective function.

While the first extension permits us to relax the standard certainty equivalence result and allows moments of higher order than the mean to affect equilibrium inflation,[6] the second extension permits the level of supply

[5] In the textbook time-inconsistency model the explanation for sub-optimal inflation relies on the presumption that policymakers use monetary policy to raise output (or employment) above its natural level. From an empirical standpoint the time-inconsistency model implies that the higher (lower) the natural rate of unemployment is, the higher (lower) the equilibrium inflation rate is (see Parkin, 1993 and Ireland, 1999).

[6] Despite the time-inconsistency model's popularity it has been questioned by both policymakers as well as by some academics on the grounds of realism (McCallum (1997) and Blinder (1998)). Such questioning led to the emergence, since the late nineties, of a new body of literature that incorporates the possible existence of asymmetries in the objective functions of central banks - the new inflation bias hypothesis, exemplified by Ruge-Murcia (2003a,b, 2004) and Cukierman and Gerlach (2003). More precisely, this literature demonstrates that

shocks to affect equilibrium inflation because of the nature of monetary institutions.[7] Finally, the third extension allows us to test the role of exchange rate regime in bringing down inflation. The testable implication we draw from this is that changes in U.S. inflation also cause changes in the domestic inflation rate.

In order to understand the implications of such preferences for optimal policy we solve the policymakers' objective function subject to a linear expectations augmented Phillips curve. The solution to the central bank's optimization problem generates a reduced form solution for inflation in which equilibrium inflation is determined by the interaction of the structural parameters of the model with the exogenous processes driving the system (namely, the trend growth rate of output, the level and volatility of supply shocks, and foreign inflation rate). This set up has proved useful in thinking about monetary

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when the central bank is also expected to engage in stabilization of output (or employment), some uncertainty about the future state of the economy and asymmetric concerns about positive and negative output gaps combine to create an inflation bias. Here, a bias arises in spite of policymakers targeting the natural rate of output. From an empirical standpoint the new inflation bias hypothesis implies that the slope parameter in a regression of inflation on the *conditional variance of the supply shocks* should be significant.

[7] An alternative hypothesis, such as those of Chari et al. (1998) and Clarida et al. (2000), suggest that a bad supply shock (e.g. increase in crude oil prices) triggers a jump in expected inflation, which then became transformed into permanently high inflation because of the nature of monetary regime in existence. This is the so called *expectations trap hypothesis*. An interesting question that arises here is what caused inflationary expectations to rise in the first place? According to the expectations trap hypothesis, the cause lies with the nature of monetary institutions themselves. From an empirical standpoint the expectations trap hypothesis suggests that if the policymaker does not find a way to credibly commit to not validating high inflation expectations, then the slope parameter in a regression of inflation on the level of supply shock should be significant.

policy and can help us evaluate the relative merits of alternative explanations advanced to explain the recent inflation moderation.

2.1 Model

We assume that the central bank chooses a sequence of short-term interest rates (i_t) in order to minimize the present discounted value of its loss function. The policymakers' objective is not only to stabilize inflation, π_t (around a given constant long-run target, π^*) and the rate of growth of output, y_t (around a time-varying target, y_t^T) but also to stabilize the nominal exchange rate.[8] Formally, the central bank faces the following problem:

$$L(\pi_t, y_t, \Delta e_t) = \frac{1}{2}\left\{(\pi_t - \pi^*)^2 + \lambda_t (y_t - y_t^T)^2 + \phi_t (\Delta e_t)^2\right\}, \qquad \textbf{(2.1)}$$

where, $0 < \lambda_t < 1$ and $0 < \phi_t < 1$ are the relative weight on output and exchange rate stabilization and Δe_t is the change in the log of the nominal exchange rate.[9] The parameter 'λ_t' is crucial for evaluating the role played by

[8] Calvo and Reinhart (2002) and Reinhart and Rogoff (2002) classify the current currency regime in India as a "peg to the US dollar". Furthermore, monitoring the nominal exchange rate, as opposed to the real exchange rate, has been the official policy (see Jalan, 1999).

[9] In our baseline model we assume that absolute purchasing power parity (PPP) theory holds, so that $e_t = p_t - p_t^f$, where 'p_t' is the domestic price level in logs and p_t^f is log of the foreign price level. This implies that, $\Delta e_t = \pi_t - \pi_t^f$, where π_t^f denotes foreign inflation. We assume that foreign inflation follows a unit root process i.e., $\pi_t^f = \pi_{t-1}^f + w_t$, where w_t is a white noise disturbance term.

monetary policy in bringing inflation down. Note that all coefficients have a '*t*' subscript to emphasize that they are potentially time-varying.

Moreover, instead of just assuming that the policymakers' target for output growth is greater than its trend growth rate (the conventional inflation bias hypothesis), we also assume that it is a (nonlinear) function of the underlying supply shock (see also Gerlach, 2003). The later assumption allows for the possibility that policymakers raise (contract) their target for output growth in response to contractionary (expansionary) supply shocks. That is,

$$y_t^T = (1+k_t)y_t^* + \frac{1}{\gamma_t}\left(e^{\gamma_t(u_t)} - 1\right), \qquad k_t > 0 \qquad \textbf{(2.2)}$$

where, y_t^* is its trend growth rate of output, 'γ_t' is the asymmetric preference parameter.[10] The supply disturbance, u_t, in turn, fluctuates over time in response to a random shock, ε_t, according to the autoregressive process (with drift),

$$u_t = \delta_t + \rho_t u_{t-1} + \varepsilon_t, \qquad \textbf{(2.3)}$$

[10] Notice that when, $\gamma_t > 0$, this specification captures recession aversion preference. To see this note that, $dy_t^T / \partial u_t = e^{\gamma_t u_t} > 0$. That is, policymakers' attempt to offset the contractionary effects of an adverse shock ($u_t > 0$), by raising their target for output growth. Also note that, $\partial^2 y_t^T / \partial u_t^{\,2} = \gamma_t e^{\gamma_t u_t} > 0$, so that they respond more strongly to large than to smaller magnitude shocks. Thus, both the sign and the magnitude of the shock are relevant to the policymaker when preferences are asymmetric.

where, $0 < \rho_t < 1$, $\delta_t > 0$. We shall also assume that the supply disturbance is conditionally hetroscedastic, $\sigma_{u,t}^2$. While the drift term in eq. (2.3) permits persistent supply shocks to affect equilibrium inflation, the conditional volatility assumption allows changes over time in the volatility of the structural shocks to affect equilibrium inflation. These assumptions in turn allow us to evaluate the good luck hypothesis.

The private sector behavior is characterized by an expectations augmented Phillips curve:

$$\pi_t = \pi_t^e + \alpha_t \left(y_t - y_t^*\right) + u_t, \qquad \alpha_t > 0 \tag{2.4}$$

where, π_t^e denotes expectations conditional upon the information available at time $t-1$.[11] The slope of the Phillips curve is crucial for evaluating the structural change explanation as discussed in detail below.

Finally, the central bank affects π_t through a policy instrument. We can interpret this instrument as the rate of growth of a monetary aggregate or as a short-term nominal interest rate. The instrument is imperfect in the sense that in a stochastic world, it cannot determine inflation completely, as in:

$$\pi_t = i_t + \xi_t, \tag{2.5}$$

[11] The trend growth rate of output is assumed to follow a unit root process, $y_t^* = y_{t-1}^* + v_t$, where v_t is a white noise disturbance term.

where, i_t is the policy instrument and $\xi_t \sim N(0,\ \sigma_\xi^2)$ is a control error uncorrelated with u_t and v_t. Since there is no private information in the model, the government's and central bank's information set coincide with the public's and are given by, Ω. Since i_t is chosen in the previous period, $i \in \Omega$.

In order to understand the implications of this model for equilibrium inflation we minimize the period loss function subject to the constraints provided by the structure of the economy, which yields:

$$\pi_t = \left(\frac{\alpha_t^{\,2}}{\lambda_t + \alpha_t^{\,2}(1+\phi_t)}\right)\left(\pi^* + \frac{\lambda_t k_t}{\alpha_t}\left(y_t^*\right) + \frac{\lambda_t}{\alpha_t^{\,2}}\left(\pi_t^e + u_t\right) + \frac{\lambda_t}{\alpha_t \gamma_t}\left\{e^{\gamma_t(u_t)} - 1\right\} + \phi_t \pi_t^f\right). \quad \textbf{(2.6)}$$

Here eq. (2.6) is the optimal inflation response to the developments in the economy. Linearization of the exponential term in (2.6) by means of a second-order Taylor series expansion around $u_t = 0$ yields, $e^{\gamma_t(u_t)} \cong 1 + \gamma_t(u_t) + \frac{\gamma_t^{\,2}(u_t)^2}{2}$. Substituting this approximation above and taking expectations conditional upon information available in period *t-1* yields:

$$\pi_t^e = a_{0t} + a_{1t} y_{t-1}^* + a_{2t} u_{t-1} + a_{3t}\sigma_{u,t}^2 + a_{4t}\pi_{t-1}^f\,, \quad \textbf{(2.7)}$$

where, $a_{0t} = \left(\frac{\alpha_t^{\,2}\pi^* + \lambda_t \delta_t (1+\alpha_t)}{\alpha_t^{\,2}(1+\phi_t)}\right)$, $a_{1t} = \left(\frac{\lambda_t k_t}{\alpha_t(1+\phi_t)}\right)$, $a_{2t} = \left(\frac{\lambda_t \rho_t (1+\alpha_t)}{\alpha_t^{\,2}(1+\phi_t)}\right)$, $a_{3t} = \left(\frac{\lambda_t \gamma_t}{2\alpha_t(1+\phi_t)}\right)$, $a_{4t} = \left(\frac{\phi_t}{1+\phi_t}\right)$ and $\sigma_{u,t}^2$ is the conditional variance of supply shock. The solution for expected inflation depends on the underlying parameters of the model and the exogenous processes driving the system. This is the sense in which the dynamics of inflation are linked to the preferences of the central bank and the structure of the economy.

An interesting feature of the model is that it nests several plausible explanations for the inflation moderation. In terms of the four special cases discussed above, the "good policy" hypothesis refers to the possibility that the increased macroeconomic stability experienced by the Indian economy in recent years can be accounted for by a substantial change in central bank responses to inflation. This has been discussed extensively in the recent academic and policy literature (see Mishkin, 2007). Indeed, over the past few decades, most of the world's major central bankers have sought to keep inflation low, some going so far as to adopt explicit inflation targeting regimes. Proponents of inflation targeting see the great moderation as evidence of these policies effectiveness.

The intuition is straightforward. The greater the apparent concern shown by the central bank for the real economy (higher, λ), the greater is the risk of falling into an expectations trap. This is because in response to an adverse supply shock (for example, the oil price shocks of the 1970s) the private sector rightly believes the central bank would accommodate a rise in inflation expectations so as to avoid a full-blown recession. This gives rise to sunspot equilibrium, as it leaves open the possibility of bursts of inflation and output that result from self-fulfilling changes in expectations (see Clarida et al., 2000). The view that monetary policy matters argues that the main contrast between the two broad periods of inflation experience in Figures 1 and 2 can be attributed to a significant change in central bank's response to inflation. Specifically, it suggests that during periods in which the central bank's resolve to stabilize inflation is strong (lower, λ), inflation would fall.

In contrast, the "structural change" explanation puts weight on the idea that increased competition due to trade openness has increased the flexibility of economies which has fundamentally altered the short-run dynamics of the inflation process. A recent study carried out at the Bank for International Settlements (Borio and Filardo, 2006) finds some empirical support for the view that globalization has flattened the short-run trade-off between inflation and the domestic output gap (lowered 'α'). This could happen through a variety of channels.

First, the increased trade and specialisation associated with globalization reduces the response of inflation to domestic output gap, and at the same time potentially makes it more sensitive to the balance between demand and supply in the rest of the world. Second, increased competition can reduce the cyclical sensitivity of profit margins, as businesses have less scope to raise their prices when domestic demand increases. Assuming marginal costs rise with output, we would expect that the mark-up of price over marginal cost will tend to be squeezed more when demand rises. Such factors could work to flatten the Phillips curve.[12]

It is worth mentioning one consequence of greater openness that might work in the opposite direction. Influential work by Romer (1993) pointed out that the Phillips curve is steeper in relatively more open economies. The intuition that the slope of the Phillips curve is related to openness is based on models of small open economies with nominal rigidities. In such models, unanticipated monetary expansion typically leads to real currency depreciation. There are potentially two effects on the trade-off. When inflation is measured in terms of a consumer price index, the effect of the depreciation on the domestic price of imports will add to the inflation cost of a monetary expansion.

[12] In this regard a survey by the Federation of Indian Chamber of Commerce and Industry (FICCI) on 'Emerging Oil Price Scenario and the Indian Industry' conducted during the month of October-November 2004 is quite revealing. The response of these firms suggests that strengthening of competition in the product market since liberalization has limited the extent to which oil prices and induced wage effects can be passed on to customers. The survey covered companies with a wide geographical and sectoral spread. The turnover of the companies that participated in the survey ranged from Rs. 1 crore to Rs. 60, 000 crore. The survey (conducted at a time when oil prices shot up to $50 a barrel) revealed that as many as 77% of the 147 companies studied said their cost of production had risen by up to 20% due to rising oil prices. However, despite this increase in costs, a majority 60% reported that this incremental cost was being absorbed internally instead of increasing their product prices. 38% reported that they are taking in a part of the incremental cost internally and passing the rest to the consumer. Only 2% were found to pass it on fully to the consumers through increased prices.

Meanwhile, if wages are partially indexed to a consumer price index, or if foreign goods are used as intermediate inputs in domestic production, the output gain to a given monetary expansion will be reduced. Both effects mean that the Phillips curve is likely to be steeper in relatively open economies. As a result policymakers have less incentive to pursue expansionary policies since the output gain from a given monetary expansion will be reduced. Hence, inflation would be lower in more open economies.

A similar line of reasoning also applies to the relationship between the trend growth rate in output (y^*_{t-1}) and inflation in our model. Thus, for example, when trend growth rate of output is low, like the 1960s and 1970s (popularly dubbed the Hindu rate of growth) policymakers' have greater temptation to generate surprise inflation. This occurs because in the absence of a credible commitment to price stability the time inconsistency model implies that changes in the trend growth rate influence policymakers' temptation to create surprise inflation. In contrast, the rise in the trend growth rate (especially in the 1990s), owing to reform measures undertaken to liberalize the economy, reduces the policymakers' reflationary instincts. As a result inflationary pressures subside.

The third explanation for the more recent experience of low inflation emphasises "good luck". The good luck hypothesis refers to the possibility that the increased macroeconomic stability experienced by the Indian economy in recent years can be accounted for by a substantial reduction in both the mean and variance of supply shocks. The model in this case predicts that a decrease in commodity prices (u_{t-1}) leads to a fall in expected inflation provided the policymaker is concerned about output stability ($\lambda > 0$). By the same token a reduction in commodity price volatility ($\sigma^2_{u,t}$) has diminished the challenges faced by policymakers charged with controlling inflation.

Finally, the exchange rate regime hypothesis corresponds to the case where, $\phi_t > 0$. The model in this case predicts a systematic relationship between domestic inflation and foreign inflation. In particular, the model implies that a fall in U.S. inflation can translate into a reduction in domestic inflation due to a dislike on the part of monetary authorities in emerging market economies of large nominal exchange rate movements. In this regard we would point out that while the spillover hypothesis has the potential to explain the moderation in inflation in the 1990s, it cannot however explain the take-off in inflation witnessed during the 1970s.

This is because until the early 1990s India was a relatively closed economy. As a result, it seems reasonable to expect the central bank to place less weight on the exchange rate objective prior to liberalization. In contrast, in the post-liberalization era trade and capital flows have expanded at a rapid pace. It seems plausible that unwelcome variability in exchange rates is perceived as more costly in relatively more open economies. With higher perceived costs of exchange rate variability, the central bank assigns greater weight on exchange rate stabilization (higher, ϕ).

In sum, all these theories are capable of accounting for the shift in inflation dynamics although the mechanism by which this shift arises differs from one theory to the other. Therefore disentangling the relative importance of each is our next challenge.

2.2 Inflation reduced-form

We now proceed to empirically evaluate our model. To do this, we substitute eq. (2.7) in eq. (2.5). Thus, our benchmark reduced-form model for inflation is given by,

$$\pi_t = a_{0t} + a_{1t} y^*_{t-1} + a_{2t} u_{t-1} + a_{3t} \sigma^2_{u,t} + a_{4t} \pi^f_{t-1} + \xi_t \ . \qquad \textbf{(2.8)}$$

Notice that the reduced-form parameters (a_{it}'s) depend on a number of key structural parameters such as, the relative weight on output gap (λ) and exchange rate stabilization (ϕ), the slope of the Phillips curve (α), the extent of supply-side distortions (k), etc. There is no strong reason to believe that these parameters have remained constant throughout the sample period.[13] If

[13] For example, a number of authors have investigated the possibility that the central banker's preference parameters, λ and k, may have shifted, perhaps due to increases in central bank

true, this calls for a technique that incorporates time-varying properties of coefficients into the estimation procedure. Kalman filtering is applied here to estimate the time-varying coefficients of the time-varying parameter (TVP) model.

3. Data, methodology, empirical results, and policy implication

3.1 Data

We use seasonally adjusted quarterly data for India spanning the period 1955:Q1-2008Q1. The published quarterly time series data on GDP is available only from 1996:Q2 and hence, quarterly data for the earlier period is obtained from annual real GDP at factor cost (seasonally adjusted) using an optimal interpolation procedure.[14] The annual point-to-point percentage change in wholesale price index (WPI) is used as dependent variable. For policy formulation the WPI (π_t) is the main measure of inflation in India since it has a broader coverage and is available at weekly frequency. Estimates of year-on-year percentage change in quarterly trend GDP (y^*_{t-1}) are obtained using the Hodrick and Prescott (1997) filter.

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independence (Alesina and Summers (1993), Campillo and Miron (1997), and Mishkin (2007, 2008). These factors, and others like them, represent potential sources of parameter changes.

[14] The quarterly data are derived by solving a quadratic optimization problem which is an inbuilt program in RATS software using the procedure DISTRIB.

With regard to supply shock researchers have traditionally used energy prices as a proxy. So we use year-on-year percentage change in quarterly WPI index for fuel, power, light and lubricants (u_{ft-1}) as a proxy. Finally, for foreign inflation (π_{t-1}^{f}) we use the year-on-year percentage change in quarterly U.S. GDP deflator. The basic data are collected from various issues of the RBI Handbook of Statistics on Indian Economy, RBI Handbook of Monetary Statistics on Indian Economy, the Report on Currency and Finance and the Federal Reserve Bank of St. Louis.

Moreover, our model predicts that the conditional variance of supply shock helps forecast inflation if policymakers' preferences are asymmetric. This prediction can only be examined in a time series framework if supply shock is conditionally heteroskedastic. Hence, before proceeding further, it is important to test whether the conditional variance of our proxy for supply shock (fuel inflation) is indeed time-varying. For this we fit an appropriate ARMA process for fuel inflation, collect the residuals and regress the square of residuals on a constant and one to four of its lags. Lagrange Multiplier (LM) test statistics for neglected ARCH to test the null hypothesis that the coefficients on lagged squared residuals are zero is rejected at conventional level of significance.

However, the conditional variance of supply shock is not directly observable. In order to address this issue, we construct an estimate of conditional volatility using a parametric ARCH model of Engle (1982) with oil inflation as our proxy for supply shock. Since the conditional variance ($\sigma_{u,t}^{2}$) then becomes a generated regressor in eq. (2.8), one must consider its effect on the efficiency and consistency of the estimates (see Ruge-Murcia, 2003b). Especially, where the conditional variance is computed using an ARCH-type model, the Maximum Likelihood estimator could be biased and inconsistent if the ARCH model is mis-specified.

Pagan and Ullah (1988) suggest specification tests to assess whether the chosen ARCH model is valid. The standard misspecification test for ARCH models is the Ljung-Box Q test for detecting autocorrelation and the LM test for neglected ARCH applied to standardised squared residuals. The mean equation contained a constant and eleven lags of dependent variable and the corresponding variance equation is specified as ARCH(1) process. These results are not reported here to save space, but they are available from us upon request. The estimated Q statistics using four lags of standardised residuals is 3.1370 with p-value of 0.535; thus, accepting the null of no autocorrelation. The LM test for neglected ARCH is 0.984 with p-value of 0.32. These results indicate that a parsimonious ARCH(1) model adequately captures the conditional heteroskedasticity when fuel inflation is used as a proxy for supply shock.

3.2 Estimation Methodology

How did the parameters of the model change and by how much? Answering these questions requires modelling the time variation in the parameters. A discrete break in any of the parameters could reflect fundamental change in the way monetary policy is conducted and/or fundamental change in the structure of the economy. However, there is no guarantee that the implied changes in the parameters were discrete. For instance, time variation stemming from an evolving view of the central bank on the economy would suggest gradual and continuous drifts in the policy parameters (see Boivin, 2006). So before proceeding further we explicitly test whether the model parameters have remained constant throughout the sample period. We use the sequential Chow test for this purpose.

As a statistical measure Sequential Chow test (Quandt Likelihood Ratio test) can be used to detect a single discrete break, multiple discrete breaks, and/or slow evolution of the regression coefficients. It tests the hypothesis that the intercept and the coefficients on the dependent variables are constant against the alternative that there is a break in all or some of the coefficients at a given date for the central 70% of the sample. Figure 3 plots the Sequential Chow test F-statistic, calculated from Generalised Least Square (GLS) estimation of the model using central 70% of the sample (15% trimming from both ends of the sample). There is evidence that at least one of the five coefficients have changed over the sample - for the periods where *F*-Statistic is above the critical value (these are the critical values provided byndrews (1993), for number of restrictions equal to five.).

In order to allow for the variation in parameters we estimate a time varying parameter (TVP) model.[15] Individual law of motion of each parameter is given in the transition equation, while their relationship to observed variables is governed by the measurement equation. Measurement equation is given by the reduced form model:

$$\pi_t = a_{0t} + a_{1t} y^*_{t-1} + a_{2t} u_{t-1} + a_{3t} \sigma^2_{u,t} + a_{4t} \pi^f_{t-1} + \xi_t = \beta^{'}_t Z_t + \xi_t \qquad \textbf{(3.2.1)}$$

where, $\xi_t \sim N(0, \ \sigma^2_\xi)$, β_t is the collection of parameters and Z_t the corresponding regressors.

Following Cooley and Prescott (1976), and most of the subsequent empirical literature which allows for time variation in parameters, we assume that the parameter vector follow random walk (without drift):

$$\beta_t = \beta_{t-1} + \eta_t, \qquad \textbf{(3.2.2)}$$

where, $\eta_t \sim N(0, Q)$. This is the model proposed by Cooley and Prescott (1976), partly as a way to empirically account for the Lucas (1976) critique on the

[15] TVP model gives an alternative to the often assumed discrete break models. Unlike standard structural stability tests such as Quandt Likelihood Ratio test or Exponential Wald test, TVP model can discriminate between different forms of instability. For a detailed discussion on Kalman Filter and the TVP model see Anderson and Moore (1979), Harvey (1989), and Durbin and Koopman (2001).

inappropriateness of stable econometric models for policy evaluations. It has been widely used in forecasting applications, and Cogley and Sargent (2001) use this specification for the parameters of their reduced-form VAR model. Since the parameters are no longer constrained to have a fixed mean, the model can accommodate fairly fundamental changes in structure of the economy and monetary policy regimes.

All the parameters of the model, including the variance of η_t, can be estimated jointly by maximum likelihood estimation (MLE) using the Kalman Filter algorithm. Provided with an estimate of the variance of, η_t, the time series of the parameters, $\{\beta_t\}$, can be obtained using the Kalman filter. Some of the variances of η_t obtained this way were tending towards zero, leading to non-convergence of MLE. This is consistent with the pile-up problem identified by Stock and Watson (1998). That is, if the variances of the state specification (the variances of η_t) are small, its maximum likelihood estimates are biased towards zero. Thus, the MLE has a (large) point mass at 0.

To avoid the pile-up problem, Stock and Watson (1998) suggest an alternative way of estimating the variance of, η_t. Their approach called the Median Unbiased Estimate (MUE) of the coefficient variance is explicitly designed to account for the deficiency of the MLE when the parameters' variances are small. The estimation method exploits the fact that the distribution of a stability test, under the alternative of a TVP model, depends on the variance of the parameters. Hence, if all other parameters are known or consistently estimable, an estimate of the variance of these parameters can be inferred from a realization of a given stability test.

Rewriting the time varying parameters expressed in equation (3.2.2) as:

$$\Delta\beta_t = \eta_t = \tau_t \upsilon_t ,$$

where η_t and υ_t are serially and mutually uncorrelated zero mean random disturbance terms and τ is scalar, which governs the size of the variance of the parameter. The median unbiased estimation procedure focuses on estimation of τ when it is small using the nesting $\tau = \lambda / T$, where T is the sample size and, λ is obtained by inverting the hetroskedaticity robust version of the QLR test, using the lookup table provided by Stock and Watson (1998). To estimate the heteroscedasticity-robust estimate of $\text{var}(\Delta\beta_t)$ we have followed the steps given in Boivin (2006):

1. An estimate of λ is obtained by inverting the heteroskedasticity-robust version of the QLR_T test, which can be performed by using a lookup table in Stock and Watson (1998);

2. Σ_{ZZ} is estimated with $T^{-1}\sum_{t=1}^{T} Z_t Z_t^{'}$ and, based on the Stock and Watson (1998) results, under the local-to-zero time-variation in the parameters, Ω can

be estimated by $T^{-1}\sum_{t=1}^{T}\hat{\varepsilon}_t^2 Z_t Z_t^{'}$ which is the White estimator of $E[Z_t\varepsilon_t\varepsilon_t^{'}Z_t^{'}]$, based on the OLS residuals, $\hat{\varepsilon}_t$.

3. $\hat{\Sigma}_{\upsilon\upsilon}$ is constructed from $\hat{\Sigma}_{ZZ}^{-1}\hat{\Omega}\hat{\Sigma}_{ZZ}^{-1}$.

Given this the variances, $\hat{\mathrm{var}}(\Delta\beta_t) = (\hat{\lambda}/T)^2\hat{\Sigma}_{\upsilon\upsilon}$ are calculated. Now the time series of $\left\{\hat{\beta}_t\right\}$ is obtained by the MLE, conditional on $\hat{\mathrm{var}}(\Delta\beta_t) = (\hat{\lambda}/T)^2\hat{\Sigma}_{\upsilon\upsilon}$. That is, $\mathrm{var}(\Delta\beta_t)$ is fixed to the values obtained through the method discussed above before estimation. This gives time varying estimate of the parameters.

3.3 Estimation Results and Analysis

The joint stability test on all the coefficients gives QLR test value of 7.5901. Using lookup Table 3 in Stock and Watson (1998), the median unbiased estimate of λ implied by the joint stability test is 6.96197, which suggests a relatively small period to period variation in the parameters. This estimate is within the range of the values for which the MLE of λ turns into trouble. According to Table 1 in Stock and Watson (1998), for the local level model and for $\lambda = 7$ the pile probability that $\hat{\lambda} = 0$ for MLE is 0.48. It is only 0.16 for the median unbiased estimate. Given this value of $\hat{\lambda}$ we estimated the heteroscedasticity robust version of the MUE. Conditional on this, we estimated the remaining parameters using MLE. The estimates of the standard deviation are reported in Table 1. Smoothed estimate of the time varying parameters for the benchmark model with two standard error confidence bands are shown in Figure 4.[16]

Table 1: Estimate of Standard Deviation (Fuel)

	σ_{ξ}	$\sigma_{\eta 0}$	$\sigma_{\eta 1}$	$\sigma_{\eta 2}$	$\sigma_{\eta 3}$	$\sigma_{\eta 4}$
Estimates	3.50 (0.000)	0.14932	0.024457	0.014126	0.015151	0.044629

Value in parenthesis (#) is the corresponding p-value. Since state variances are estimated using the MUE, we don't have the corresponding standard errors or p-value

The response coefficients on trend growth rate of output (a_{1t}) and the level and volatility of supply shocks (a_{2t} and a_{3t}) have gradually decreased over the sample period. Specifically, these responses remained strong until the late 1970s and then started to gradually decrease. What explains the time variation in these coefficients? In our model the central bank's preference parameters and the slope of the Phillips curve are key determinant of inflation dynamics. Hence, time variation of these parameters should be reflected in changes in inflation dynamics.

Time variation in the preference for output and exchange rate stability can be motivated in a number of ways. First, an incoming governor may hold views

[16] For our purpose looking at the smoothed estimates is more appropriate, as our purpose is not to use Kalman Filter to produce forecasts. Rather our objective is to capture accurate information about the path followed by the time-varying coefficients.

about the desired degree of output stability (and exchange rate stability) that differ from their incumbents' opinions. Second, political pressure may at times be expressed about the desirable degree of output (and exchange rate) volatility, particularly when output losses are needed to offset inflationary cost-push shocks. Third, the central bank's perception of the relationships in the economy may change over time. For example, as evidence accumulates that lower inflation is conducive to better economic outcomes, a central bank may develop a taste for stabilizing inflation (lower λ and ϕ). Thus a plausible argument for the moderation in inflation can be attributed to a significant change in central bank's response to inflation.

In contrast, the "structural change" explanation puts weight on the idea that increased competition has increased the flexibility of economies and has fundamentally altered the short-run dynamics of the inflation process. For example, an increase in the elasticity of inflation with respect to the output gap (higher, α) would, all else constant, reduce these coefficients. While our model, by its very nature, does not allow one to uncover the deep structural sources, it can at the very least shed light on the importance of these explanations. Thus, for example, a discrete break in any of the parameters could reflect fundamental change in the way monetary policy is conducted and/or signify fundamental change in the structure of the economy. Is there evidence of a discrete break in these coefficients?

Figure 4 suggests that the evolution of these coefficients is unidirectional and at best gradual. The implied changes in the parameters were not discrete. Rather what we observe is gradual evolution of the regression coefficients. Thus our findings would suggest that changes in the underlying monetary policy framework and/or structural change cannot fully account for the recent moderation in inflation in India. In sum, our findings ascribe to monetary policy and structural change explanation quite a modest role in inflation stabilization.

Finally, the response coefficient on foreign inflation (a_{4t}) has gradually risen, especially since the 1990s. In this regard, we would point out that until the early 1990s the Indian economy was relatively closed to both trade and capital flows. In response to the external debt crisis which surfaced in 1991 the government set in motion a process of economic liberalization and structural reforms which sought to increase the degree of openness of the economy.[17] Subsequently, both trade and capital flows have expanded considerably. Foreign capital surged in, creating pressures for exchange rate appreciation.

[17] The process began with the introduction of convertibility on trade as quantitative restrictions on imports, except for consumer goods, were dismantled and tariff levels were reduced. It was combined with a liberalization of the regimes for foreign investment and foreign technology. And restrictions on international economic transactions, including capital movements, were progressively reduced. Subsequently, there was a gradual dismantling of controls on capital outflows. In particular restrictions on overseas investment and lending and the prepayment of foreign loans have been gradually relaxed. This process was also influenced by the gathering momentum of globalization which was associated with increasing economic openness in trade flows, investment flows and financial flows.

With the rupee appreciating, concerns about export competitiveness mounted, forcing the RBI to actively intervene in the foreign exchange market to curb "excessive" exchange rate variability (Ramachandran and Srinivasan, 2007).[18] Under such circumstances, a decrease in U.S. inflation forces domestic monetary authorities to allow domestic inflation to fall, so as to avoid an appreciation of the domestic currency. Thus, our results are consistent with the view that, lower U.S. inflation (Figure 5 plots the time-varying mean of U.S. inflation) may have spilled over to India due to dislike on the part of the RBI of large nominal exchange rate volatility.

To add to this the Indian economy has not in more recent times, been subjected to two many inflationary cost shocks of the kind that we saw in the 1970s. Figures 6 and 7 are plots of the time-varying mean and volatility of fuel inflation. Notice that unlike the 1970s both the mean and the volatility of supply shocks have been fairly benign in recent years. This could clearly contribute to the turnaround on the inflation front in our model. That is, even if the institutional structure governing monetary policy were pretty much the same, our model predicts that favourable supply-side developments can

[18] According to official pronouncements India moved away from a fixed exchange rate regime to a market determined exchange rate regime in March 1993. Nevertheless, the RBI actively intervenes in the foreign exchange market with the goal of 'containing volatility', and influencing the market value of the exchange rate. As Governor Reddy (1997) states, "*In the context of large capital flows (inflows as well as outflows) within a short period, it may not be possible to prevent movements in the exchange rate away from the fundamentals. Hence, the management of rate fluctuations becomes passive i.e., one of preventing undue appreciation in the context of large inflows and providing supply of dollars in the market to prevent sharp depreciations*".

prevent inflationary expectations from gaining a permanent hold. So another plausible explanation for the moderation in inflation is that good luck has diminished the challenges faced by policymakers charged with controlling inflation.

3.4 Robustness

We next explore the robustness of our findings along two dimensions. We begin by considering an alternative proxy for supply shock and its conditional volatility, namely the year-on-year percentage change in quarterly WPI index for primary commodities (u_{ct-1}). We note that the RBI reports quarterly index number of WPI for primary commodities only from 1971:Q2. Nevertheless, Chandhok (1978) reports yearly index numbers of primary commodities from 1947 to 1987. So we interpolate this data (using the procedure 'DISTRIB') to generate quarterly index number for primary commodities up to 1971:Q1.[19]

The joint stability test on all the coefficients gives QLR test value of 16.48. Using look up Table 3 in Stock and Watson (1998), the median unbiased estimate of λ implied by the joint stability test is 12.62. According to Table 1

[19] Note that unlike conditional volatility estimates with fuel inflation, when the year-on-year percentage change in quarterly WPI index for primary commodities is used as a proxy for supply shock a parsimonious GARCH(1, 2) model adequately captures the conditional volatility of supply shock. The results are not reported here but are available from the author(s) upon request. We use these estimates as our proxy for conditional volatility of supply shocks in our robustness exercise. Figures 8 and 9 are plots of the time-varying mean and volatility of primary inflation.

in Stock and Watson (1998), for the local level model and for $\lambda = 12$ the pile-up probability that $\hat{\lambda} = 0$ for MLE is 0.24. It is only 0.06 for the median unbiased estimate. Given this value of $\hat{\lambda}$ we estimated the heteroscedasticity robust version of the MUE. Conditional on this, we estimated the remaining parameters using MLE. The estimates of the standard deviation are reported in Table 2. Smoothed estimate of the time varying parameters with two standard error confidence bands are shown in Figure 10.

Table 2: Estimate of Standard Deviation (Primary)

	σ_{ξ}	$\sigma_{\eta 0}$	$\sigma_{\eta 1}$	$\sigma_{\eta 2}$	$\sigma_{\eta 3}$	$\sigma_{\eta 4}$
Estimates	1.971331 (0.0000)	0.200871	0.034103	0.023938	0.039408	0.061475

Value in parenthesis (#) is the corresponding p-value. Since state variances are estimated using the MUE, we don't have the corresponding standard errors or p-value

In the baseline model we assumed the central bank targets the nominal exchange rate and absolute PPP holds. One could argue that while the nominal exchange rate, especially relative to the U.S. dollar, is the focus of much attention, what really matters for overall export competitiveness is the real (effective) exchange rate. In our robustness exercise we replace nominal exchange rate with real exchange rate to check whether this makes any difference to our baseline results.[20]

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[20] Formally, the central bank faces the following problem: $L(\pi \ , y_t, \Delta e_t) = \frac{1}{2}\left\{(\pi_t - \pi^*)^2 + \lambda_t (y_t - y_t^T)^2 + \phi_t (\Delta R_t)^2\right\}$, where, $R_t = e_t + p_t - p_t^f$, is

The joint stability test on all the coefficients gives QLR test value of 7.9546. Using look up Table 3 in Stock and Watson (1998), the median unbiased estimate of λ implied by the joint stability test is 7.2142. According to Table 1 in Stock and Watson (1998), for the local level model and for $\lambda = 7$ the pile probability that $\hat{\lambda} = 0$ for MLE is 0.48. It is only 0.16 for the median unbiased estimate. Given this value of $\hat{\lambda}$ we estimated the heteroscedasticity robust version of the MUE. Conditional on this, we estimated the remaining parameters using MLE. The estimates of the standard deviation are reported in Table 3. Smoothed estimate of the time varying parameters with two standard error confidence bands are shown in Figure 11.

Table 3: Estimate of Standard Deviation (Real Exchange Rate)

	σ_{ξ}	$\sigma_{\eta 0}$	$\sigma_{\eta 1}$	$\sigma_{\eta 2}$	$\sigma_{\eta 3}$	$\sigma_{\eta 4}$
Estimates	3.572416 (0.0000)	0.157731	0.025825	0.015075	0.016221	0.012778

Value in parenthesis (#) is the corresponding p-value. Since state variances are estimated using the MUE, we don't have the corresponding standard errors or p-value

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the real exchange rate in logs and e_t is the nominal exchange rate in logs. In order to understand the implications of this model for equilibrium inflation we minimize the period loss function subject to the constraints provided by the structure of the economy, which yields the following reduced form solution for inflation: $\pi_t = a_{0t} + a_{1t} y^*_{t-1} + a_{2t} u_{t-1} + a_{3t} \sigma^2_{u,t} + a_{4t}\left(\Delta e_{t-1} + \pi^f_{t-1}\right) + \xi_t$, where $\left(\Delta e_{t-1} + \pi^f_{t-1}\right)$ is the % change in rupee value of foreign price level.

These results demonstrate that in each instance, the insights from the baseline model remain largely intact. The key result from the baseline case is robust to the use of primary goods inflation as a proxy for supply shock. Importantly, the conclusions about the general evolution of the coefficients remain largely intact: the changes are gradual and the timing is essentially the same. Nevertheless, when we use real exchange rate (instead of nominal exchange rate) in our baseline model the spillover hypothesis is not validated. A plausible explanation for this is that monitoring the nominal exchange rate, as opposed to the real exchange rate, has been the official policy.[21]

Indeed a confluence of forces has in recent years put enormous pressure on India's real exchange rate to appreciate. So what should the RBI do when foreign capital surges in, creating pressures for exchange rate appreciation, over and above other pressures stemming from productivity growth and excess domestic demand? As Rajan and Prasad (2008) argue the medium term steps are naturally to work harder on reducing domestic demand and increasing supply. In the short term, though, legitimate concerns arise about export growth and the loss of job growth that any loss of external

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[21] In fact, the former Governor of the Reserve Bank of India Jalan (1999) states: "*From a competitive point of view and also in the medium term perspective, it is the REER, which should be monitored as it reflects changes in the external value of a currency in relation to its trading partners in real terms. However, it is no good for monitoring short-term and day-to-day movements as 'nominal' rates are the ones which are most sensitive of capital flows. Thus, in the short run, there is no option but to monitor the nominal rate.*"

competitiveness would entail. These political pressures have led the RBI to try and manage the nominal exchange rate by intervening in foreign exchange markets. Since monitoring the nominal exchange rate, as opposed to the real exchange rate, has been the official RBI policy, it is not at all surprising that our results are not robust with real exchange rate.

3.5 Policy Implications

We began this paper by recalling that during the 1970s and 1980s the Indian economy experienced recurrent bouts of high inflation together with sub-par economic performance. Since then there has also been a marked decline in mean inflation. We then asked two questions: What explains this turnaround? And more importantly, will this progress on inflation be sustained, or is the recent improvement only a temporary lull?

The candidate explanations for the inflation moderation include better monetary policy, structural change, good luck and exchange rate regime. Sorting out the relative merits of these potential causes has important implications for policymakers. If the moderation happened as a result of improved monetary policy, and the extent that this moderation has been beneficial, then policy should continue to be made this way. In contrast, if the moderation is down to structural changes, then public policy should encourage such competition and flexibility. Finally, if good luck or exchange rate regime played an important role, then policymakers' should recognize that this may not last indefinitely and they should prepare for a turn for the worse.

Thus for example, a central bank which is burdened with other objectives such as exchange rate stability cannot guarantee price stability simply because these goals are in direct conflict with the objective of price stability. A threat to the credibility for low inflation tends to arise if an appreciation of the exchange rate hurts export competitiveness and results in job losses. Under such circumstances the central bank is forced to pursue expansionary monetary policy (buying U.S. dollars in the foreign exchange market) to prevent the domestic currency from appreciating. This creates a doubt in the public's mind about whether a central bank can sustain domestic price stability if the pressure on exchange rate continues. The exchange rate must be allowed to adjust flexibly if a country is to enjoy the benefits that monetary policy can deliver.

In sum our empirical results suggest that both good luck and exchange rate regime have played a major role in the moderation of inflation. This interpretation suggests that to prevent a resurgence of 1970s-style inflation, the central bank should reinforce as much as possible its commitment to low inflation by institutional, operational, and rhetorical means. Otherwise, sooner or later, luck will dry out and high inflation could return to haunt us.

4. Summary and Concluding Remarks

Since the mid-1990s the Indian economy has experienced only mild inflation. This is in sharp contrast to the period immediately preceding it. An important question is what ultimately brought about this improved economic outcome. There are several plausible explanations for this: better monetary policy, luck, structural change, exchange rate regime, etc. Some or all these factors likely played a role, and disentangling the relative importance of each is an important challenge. To this end, we construct a reduced form econometric model with drifting coefficients for inflation that nests all these candidate explanations.

The time variation in parameters is modeled as driftless random walks, and is estimated using the median unbiased estimator. We find that while better monetary policy and structural change have played a non-trivial role, good luck and exchange rate regime have played a major role in the moderation of inflation in the 1990s. It follows that, to prevent a resurgence of inflation the RBI should reinforce as much as possible its commitment to price stability by institutional, operational, and rhetorical means. For this, it must receive a mandate that allows it to place less weight on developments in the real side of the economy and to commit more credibly to price stability. Moreover, for this reform to be credible we need an independent central bank free from political interference. Otherwise, sooner or later luck will dry out and high and unstable inflation could return to haunt us.

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Figure 1: Quarterly WPI (All Commodities) Inflation 1955Q1-2008Q1

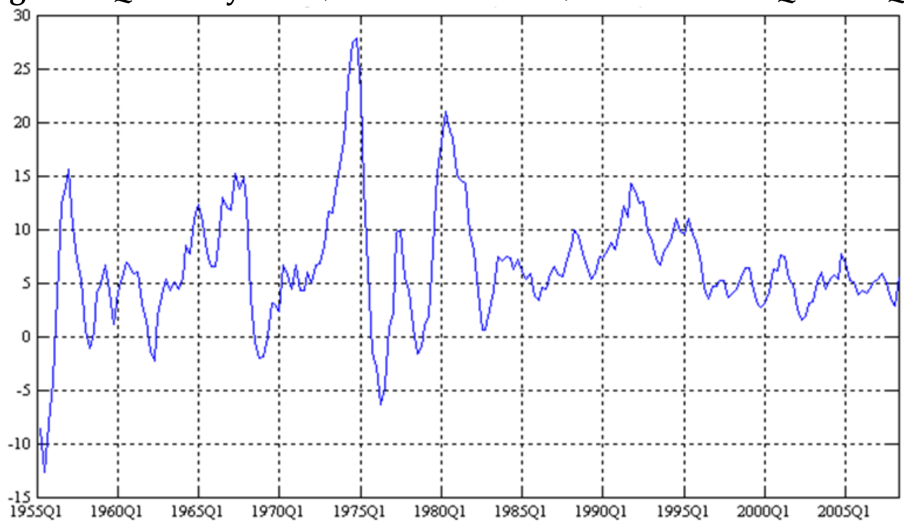


Figure 2: Estimated Constant from rolling 10-year WPI (All Commodities) Inflation regression

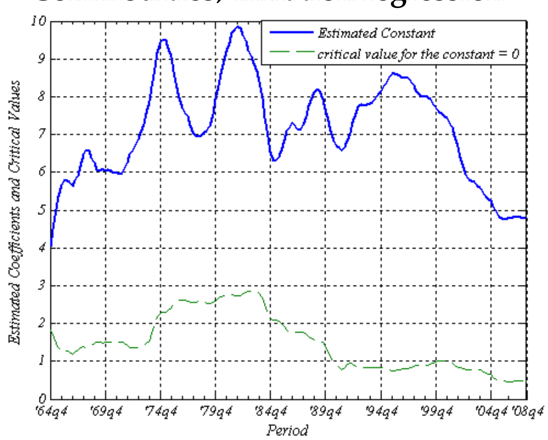
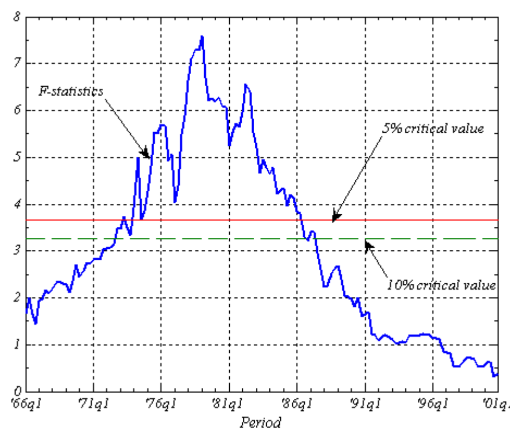


Figure 3: *F*-Statistics Testing for a Break in Equation 3.2.1 at Different Dates*



*critical values are from Andrews, 1993

Figure 4: TVP estimates of Benchmark Model

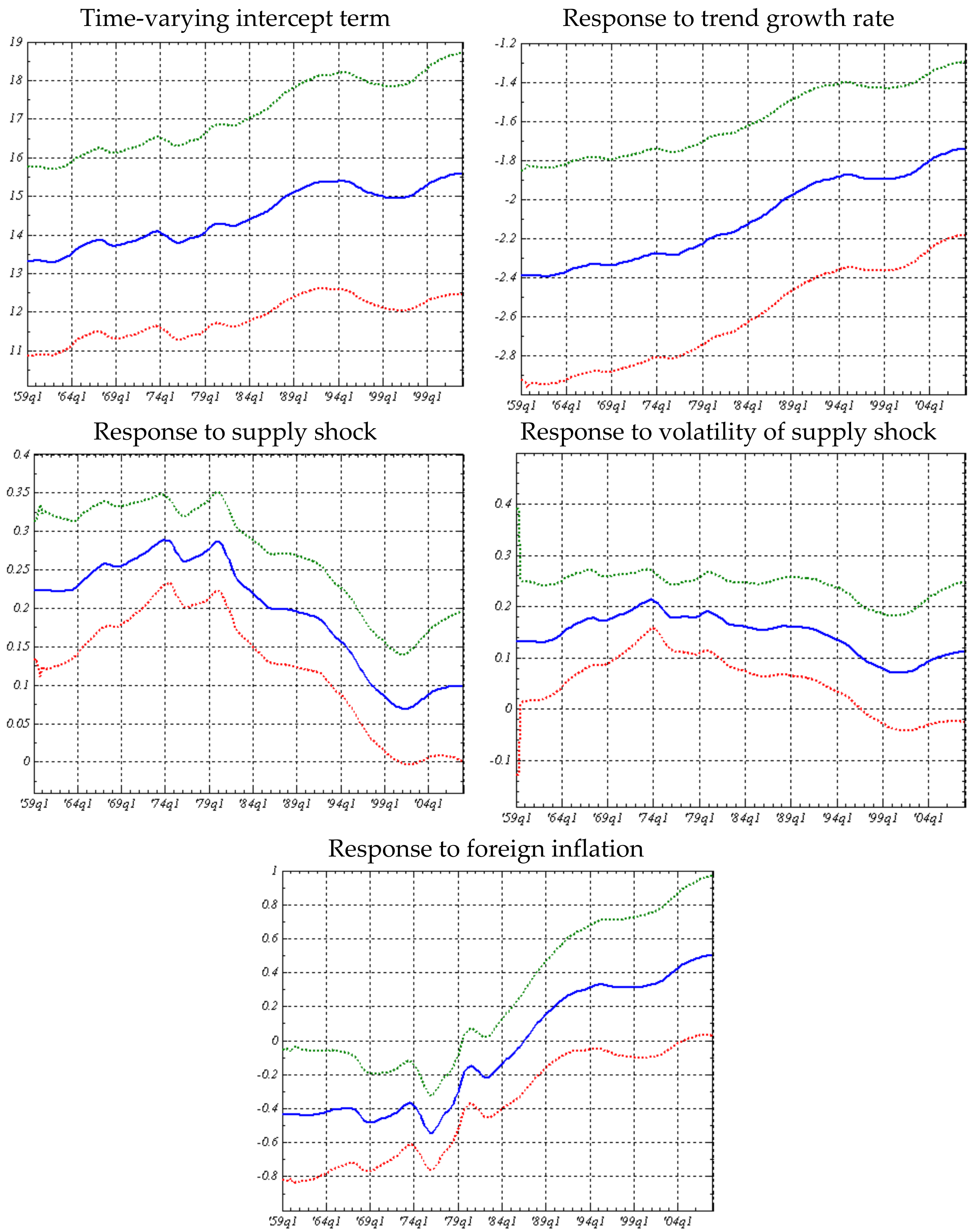


Figure 5: Estimated Constant from rolling 10-year USA Inflation regression

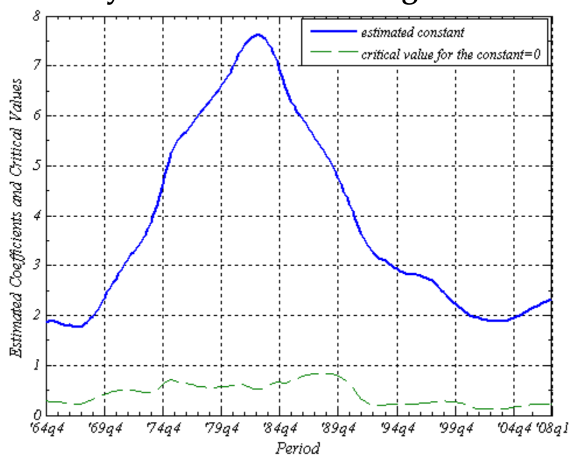


Figure 6: Estimated Constant from rolling 10-year WPI (Fuel) Inflation regression

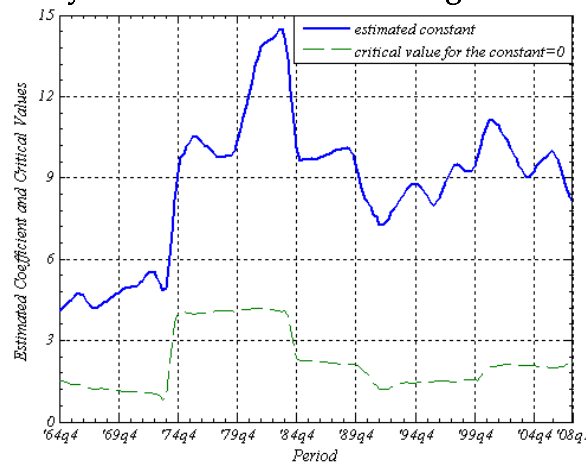


Figure 7: Time-Varying volatility of WPI Fuel inflation (rolling 10-year standard deviation)

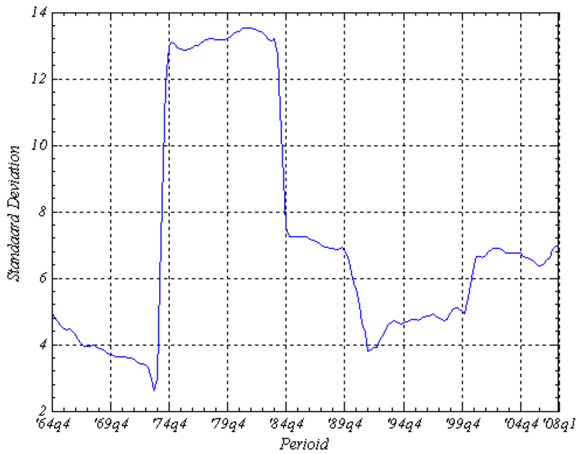


Figure 8: Estimated Constant from rolling 10-year WPI (Primary) Inflation regression

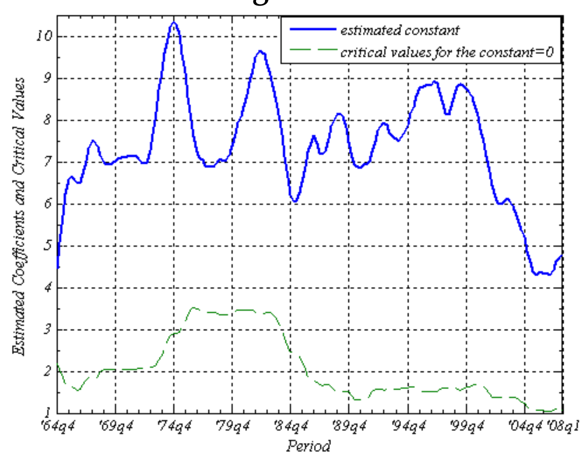


Figure 9: Time-Varying volatility of WPI Primary inflation (rolling 10-year standard deviation)

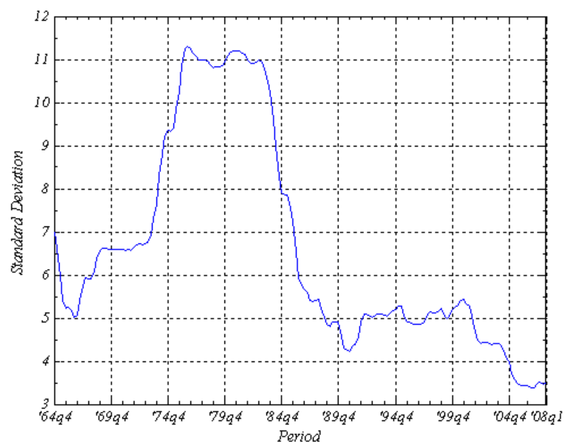


Figure 10: TVP estimates with Primary Inflation as proxy for supply shock

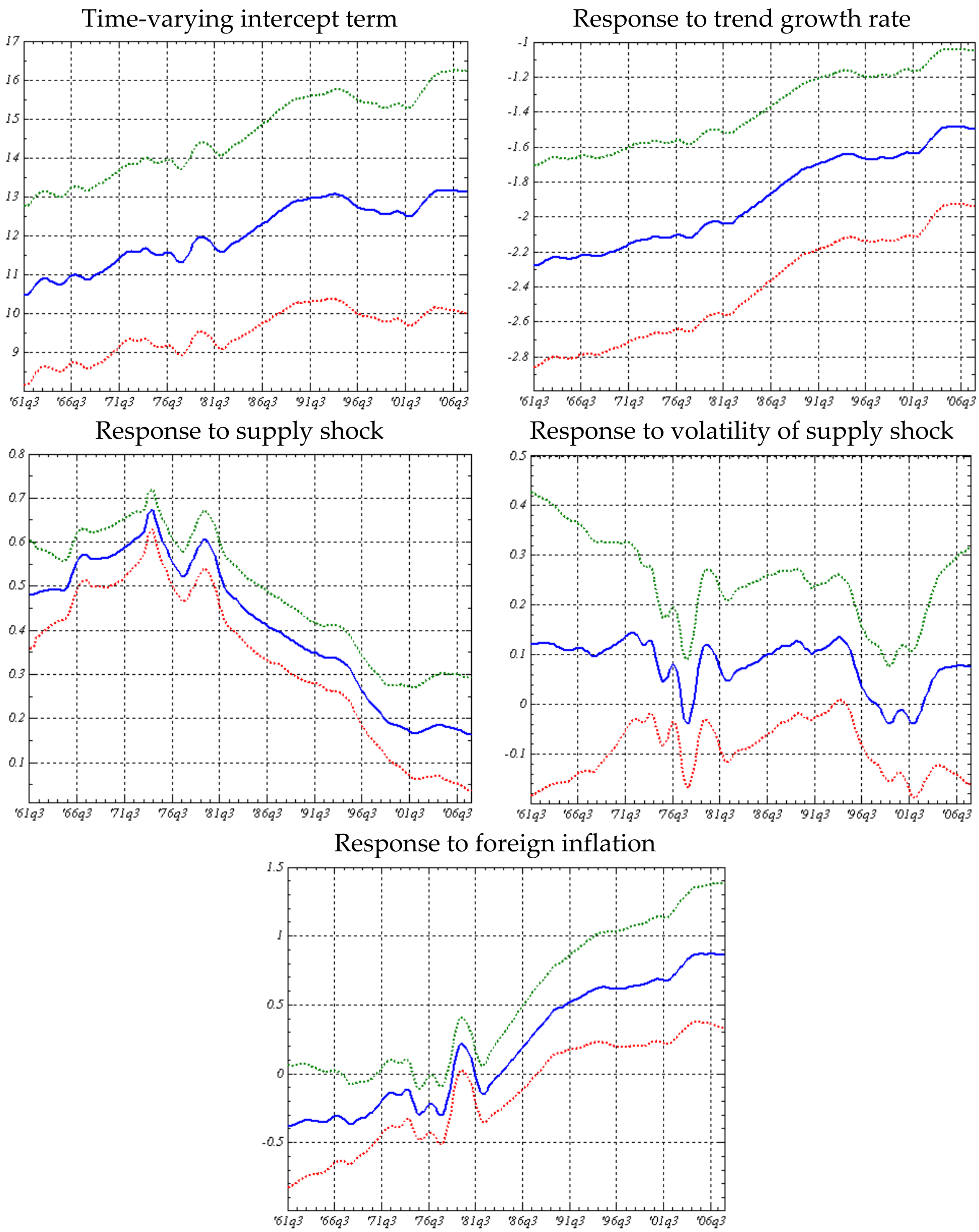
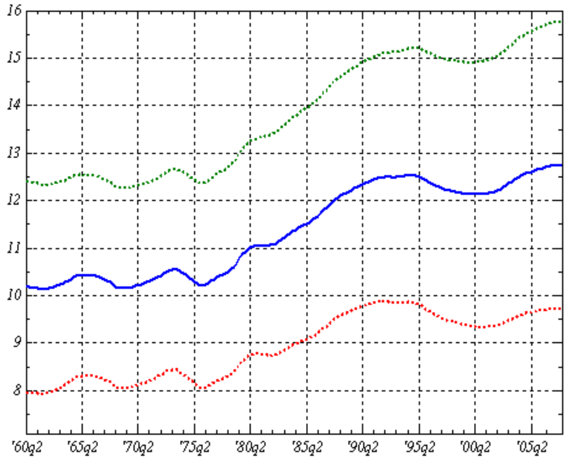
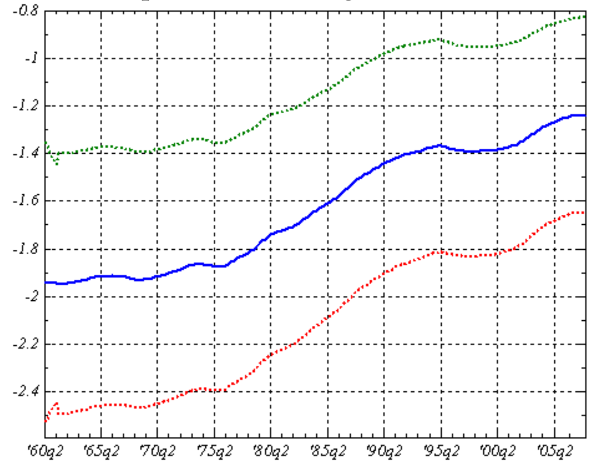


Figure 11: TVP estimates with Real Exchange Rate

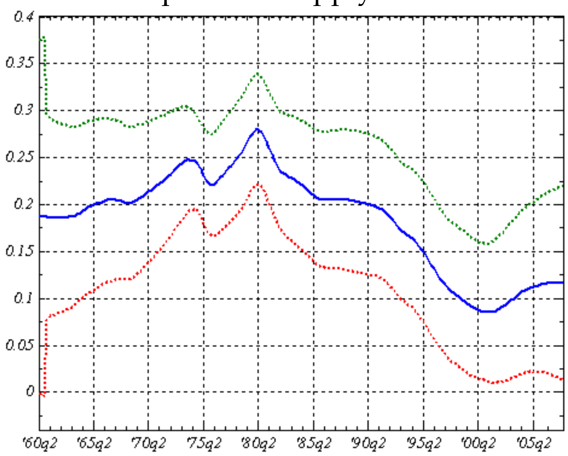
Time-varying intercept term



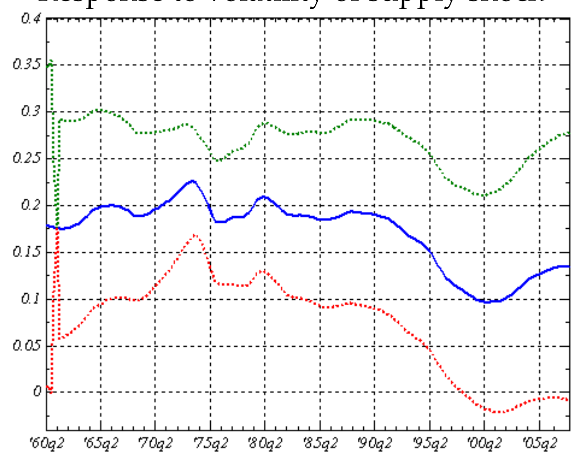
Response to trend growth rate



Response to supply shock



Response to volatility of supply shock



Response to % change in rupee value of the foreign price level