Evolving U.S. Monetary Policy and the Decline of Inflation Predictability

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Source: *Journal of the European Economic Association*, Apr. - May, 2008, Vol. 6, No. 2/3, Proceedings of the Twenty-Second Annual Congress of the European Economic Association (Apr. - May, 2008), pp. 634-646

Published by: Oxford University Press

Stable URL: https://www.jstor.org/stable/40282672

EVOLVING U.S. MONETARY POLICY AND THE DECLINE OF INFLATION PREDICTABILITY

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Abstract
Based on a structural VAR with time-varying parameters and stochastic volatility for the post-WWII U.S., we document a negative correlation between the evolution of the long-run coefficient on inflation in the structural monetary rule and the evolution of the persistence and predictability of inflation relative to a trend component. Using an estimated sticky-price model, we show that a more aggressive policy stance towards inflation causes a decline in inflation predictability. (JEL: E37, E52, E58)

1. Introduction

In a recent contribution, Cogley and Sargent (2006) showed that the persistence of the U.S. inflation gap, defined as the deviation of inflation from a trend component, declined sharply around the time of the Volcker disinflation. Stock and Watson (2007) and D'Agostino, Giannone, and Surico (2006) have shown that the predictability of U.S. inflation has fallen sharply over the post-1984 period. In this paper, we suggest that these two findings can be interpreted as the result of a more aggressive behavior of the Fed against inflation.

In fact, based on the estimation of a structural vector autoregressive (SVAR) model with time-varying coefficients and stochastic volatility, we document a strong negative correlation between the evolution of the long-run coefficient on inflation in the monetary rule and the evolution of the persistence and predictability of inflation. Based on an estimated dynamic stochastic general equilibrium (DSGE) model, we show that a more aggressive policy stance towards inflation causes a fall in both the persistence and predictability of inflation, thus providing a possible interpretation for the evidence uncovered via the VAR.

Acknowledgments: We wish to thank the editor (Roberto Perotti) and two anonymous referees for useful suggestions, Juan Rubio-Ramirez for kindly providing a MATLAB code for computing the VAR representation of a state-space form, and Lucrezia Reichlin for comments. Usual disclaimers apply. The opinions expressed herein are those of the authors and do not necessarily reflect the position of the European Central Bank or the Bank of England.

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Journal of the European Economic Association April–May 2008 6(2–3):634–646

1.1. Comparison with the Literature

Time-varying parameters or Markov-switching SVARs have been used by Canova and Gambetti (2005), Primiceri (2005), Sims and Zha (2006), and Gambetti, Pappa, and Canova (forthcoming) to investigate the sources of the Great Moderation in the United States. Gambetti, Pappa, and Canova do not analyze the evolution of the VAR's structural monetary rule, but Canova and Gambetti—who, as in the present work, achieve identification via sign restrictions—show the mean changes in the structural coefficients of the monetary rule. The problem is that it is not possible to recover the evolution of what truly matters—that is, the long-run coefficients—from the changes in the coefficients of the monetary rule. Using a Cholesky identification, Primiceri (2005) documents little change in the long-run response to the unemployment rate and some increase in the long-run response to inflation. Sims and Zha's best fit model (see Section 5) allows for changes only in the variance-covariance matrix, thus ruling out time variation in the structural monetary rule by definition. When they consider a model with a worse fit which allows for changes in the variance-covariance matrix and the policy rule, the evidence is difficult to assess[1] but does not point toward dramatic changes in the monetary rule.

So a first contribution of the present work is to provide novel evidence on changes in the long-run response to inflation in the structural monetary rule within an SVAR framework.[2] Second, we relate such changes to changes in the persistence and predictability of inflation, and we interpret the two sets of facts in the light of an estimated DSGE model.[3] In a paper conceptually related to ours, Boivin and Giannoni (2006) focus on changes in the effectiveness of monetary policy (and especially they largely focus on impulse-response functions), eschewing the issues of inflation persistence and predictability. Hence this paper should be essentially regarded as complementary to theirs.

The paper is organized as follows. In Section 2 we present results based on the structural VAR, and in Section 3 we present results based on a widely used model of the business cycle. Our evidence is consistent with the notion that at least part of the decline in both the persistence of the inflation gap and the predictability of inflation may be attributable to the Fed's adoption of a more aggressively counter inflationary stance after October 1979. Section 4 concludes.

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1. This is because for the Volcker and Burns regimes "the responses of the Federal Funds rate are, variable by variable, so ill-determined that we instead present responses to money growth" (Sims and Zha 2006, p. 65).

2. Cogley and Sargent (2002) tackle the same issue by interpreting the interest rate equation in their reduced-form time-varying VAR as a time-varying Taylor rule. Boivin (2004), instead, estimates a time-varying Taylor rule.

3. Within the DSGE literature, Lubik and Schorfheide (2004) and Justiniano and Primiccri (2008) have presented sub-sample analyses based on estimated DSGE models; Clarida, Galí, and Gertler (2000) have focused on changes in the Taylor rule within a single-equation framework.

2. Empirical Evidence

2.1. A VAR with Time-Varying Parameters and Stochastic Volatility

In this section, we work with the following time-varying parameters VAR(p) model

$$Y_t = B_{0,t} + B_{1,t}Y_{t-1} + \ldots + B_{p,t}Y_{t-p} + \varepsilon_t \equiv X_t'\theta_t + \varepsilon_t, \tag{1}$$

where Y_t is defined as $Y_t \equiv [r_t, \pi_t, y_t, m_t]'$ and r_t, π_t, y_t and m_t, respectively, are the short-term interest rate, inflation, output growth, and money growth.[4] For sake of comparability with earlier contributions, we set the lag order, p, to 2. It is postulated that the VAR time-varying parameters, collected in the vector θ_t, evolve according to

$$p(\theta_t \mid \theta_{t-1}, Q) = I(\theta_t) \cdot f(\theta_t \mid \theta_{t-1}, Q), \tag{2}$$

where $I(\theta_t)$ is an indicator function rejecting unstable draws—thus enforcing a stationarity constraint on the VAR—and where $f(\theta_t \mid \theta_{t-1}, Q)$ is given by

$$\theta_t = \theta_{t-1} + \eta_t, \tag{3}$$

with $\eta_t \sim N(0, Q)$. The VAR reduced-form innovations in equation (1) are postulated to be zero-mean normally distributed with time-varying covariance matrix Ω_t; following established practice, we factor this as

$$\operatorname{Var}(\varepsilon_t) \equiv \Omega_t = A_t^{-1} H_t \left(A_t^{-1}\right)'. \tag{4}$$

The time-varying matrices H_t and A_t are defined as follows:

$$H_t \equiv \begin{bmatrix} h_{1,t} & 0 & 0 & 0 \\ 0 & h_{2,t} & 0 & 0 \\ 0 & 0 & h_{3,t} & 0 \\ 0 & 0 & 0 & h_{4,t} \end{bmatrix}, \quad A_t \equiv \begin{bmatrix} 1 & 0 & 0 & 0 \\ \alpha_{21,t} & 1 & 0 & 0 \\ \alpha_{31,t} & \alpha_{32,t} & 1 & 0 \\ \alpha_{41,t} & \alpha_{42,t} & \alpha_{43,t} & 1 \end{bmatrix}, \tag{5}$$

with the elements $h_{i,t}$ evolving as geometric random walks:

$$\ln h_{i,t} = \ln h_{i,t-1} + \nu_{i,t}. \tag{6}$$

4. Data sources: Federal Funds rate ('FEDFUNDS, Effective Federal Funds Rate, Board of Governors of the Federal Reserve System, Monthly, Percent'), which we convert to the quarterly frequency by taking averages within the quarter; GDP deflator inflation based on GDPDEF ('Gross Domestic Product: Implicit Price Deflator, Quarterly, Seasonally Adjusted'); the output growth, computed as the log difference of GDPC1 ('Real Gross Domestic Product, 1 Decimal'), from the *Bureau of Economic Analysis*; and the money growth, computed as the log difference of M2 ('Money Stock, M2SL, Board of Governors of the Federal Reserve System, Seasonally Adjusted, Monthly, Billions of Dollars') from the St. Louis FED.

For future reference, we define $h_t \equiv [h_{1,t}, h_{2,t}, h_{3,t}, h_{4,t}]'$. Following Primiceri (2005), and in line with Benati and Mumtaz (2007) and Benati (2008a), we postulate that the non-zero and non-one elements of the matrix A_t—which we collect in the vector $\alpha_t \equiv [\alpha_{21,t}, \alpha_{31,t}, \ldots, \alpha_{43,t}]'$—to evolve as driftless random walks,

$$\alpha_t = \alpha_{t-1} + \tau_t. \tag{7}$$

We likewise assume that

$$\begin{bmatrix} u_t \\ \eta_t \\ \tau_t \\ \nu_t \end{bmatrix} \sim N(0, V),$$

with

$$V = \begin{bmatrix} I_4 & 0 & 0 & 0 \\ 0 & Q & 0 & 0 \\ 0 & 0 & S & 0 \\ 0 & 0 & 0 & Z \end{bmatrix} \quad \text{and} \quad Z = \begin{bmatrix} \sigma_1^2 & 0 & 0 & 0 \\ 0 & \sigma_2^2 & 0 & 0 \\ 0 & 0 & \sigma_3^2 & 0 \\ 0 & 0 & 0 & \sigma_4^2 \end{bmatrix}, \tag{8}$$

where u_t is such that $\varepsilon_t \equiv A_t^{-1} H_t^{1/2} u_t$. In line with Primiceri (2005), we adopt the additional simplifying assumption of a block-diagonal structure for S:

$$S \equiv \mathrm{Var}(\tau_t) = \mathrm{Var}(\tau_t) = \begin{bmatrix} S_1 & 0_{1\times 2} & 0_{1\times 3} \\ 0_{2\times 1} & S_2 & 0_{2\times 3} \\ 0_{3\times 1} & 0_{3\times 2} & S_3 \end{bmatrix}, \tag{9}$$

where $S_1 \equiv \mathrm{Var}(\tau_{21,t})$, $S_2 \equiv \mathrm{Var}([\tau_{31,t}, \tau_{32,t}]')$, and $S_3 \equiv \mathrm{Var}([\tau_{41,t}, \tau_{32,t}, \tau_{43,t}]')$, implying that the non-zero and non-one elements of A_t that belong to different rows evolve independently. As discussed in Primiceri (2005, Appx. A.2), this assumption drastically simplifies inference, because it enables Gibbs sampling on the non-zero and non-one elements of A_t equation by equation.

We estimate equations (1)–(9) via Bayesian methods. The details of the methodology—including the choices for the priors, the Markov-Chain Monte Carlo algorithm used to simulate the posterior distribution of the hyper-parameters, and the states conditional on the data, and the method we use to assess the convergence of the Markov chain—are identical to those used in Benati (2008a) and Benati and Mumtaz (2007), to which the interested reader is referred.

2.2. Evolving Persistence and Predictability of the U.S. Inflation Gap

We approximate the time-varying spectral density of the inflation gap by Fourier-transforming the estimated time-varying VAR:

$$f_{\pi,t|T}(\omega) = s_\pi \left(I_4 - \sum_{k=1}^{p} B_{k,t|T} e^{-ik\omega} \right)^{-1} \frac{\Omega_{t|T}}{2\pi} \left[\left(I_4 - \sum_{k=1}^{p} B_{k,t|T} e^{ik\omega} \right)^{-1} \right]' s_\pi', \tag{10}$$

where s_π is a row vector selecting inflation. Based on equation (10), we then compute persistence as the normalized spectrum of inflation at $\omega = 0$. Following Cogley and Sargent (2006), predictability is measured as a function of the ratio between the conditional and the unconditional variance of inflation, which we approximate as

$$R^2_{\pi,t} \simeq 1 - \frac{s_\pi \Omega_t s_\pi'}{s_\pi \left[\sum_{h=0}^{\infty} F_t^h \Omega_t \left(F_t^h \right)' \right] s_\pi'}. \tag{11}$$

Here F is the companion matrix of the VAR.

It is worth emphasizing that, unlike in the univariate case, in a multivariate context the mapping between inflation persistence and inflation predictability is not one-to-one, so the two features ought to be investigated from both a theoretical and an empirical standpoint.

In line with Cogley and Sargent (2006), the panels in Figure 1 show that the U.S. inflation gap has become less persistent as well as less predictable since the end of the Volcker disinflation.[5] Because inflation is here, by definition, equal to the sum of the inflation gap and a trend component evolving to a first approximation as a random walk, it follows that the fall in the inflation gap predictability translates into a fall in the predictability of inflation itself.

What did cause these changes? Could monetary policy have played a role? To provide a tentative answer, we need to identify a structural monetary rule.

2.3. Evolution of the Structural Monetary Rule

We identify four structural shocks—monetary policy (ε_t^M), supply (ε_t^S), demand non-policy (ε_t^D), and money demand (ε_t^{MD})—by imposing the sign restrictions reported in Table 1 on the contemporaneous impacts of the structural shocks on the four endogenous variables. It can be trivially shown that these restrictions

5. The link between the persistence and predictability of a series have been discussed by Granger and Newbold (1986) and Barsky (1987), among others.

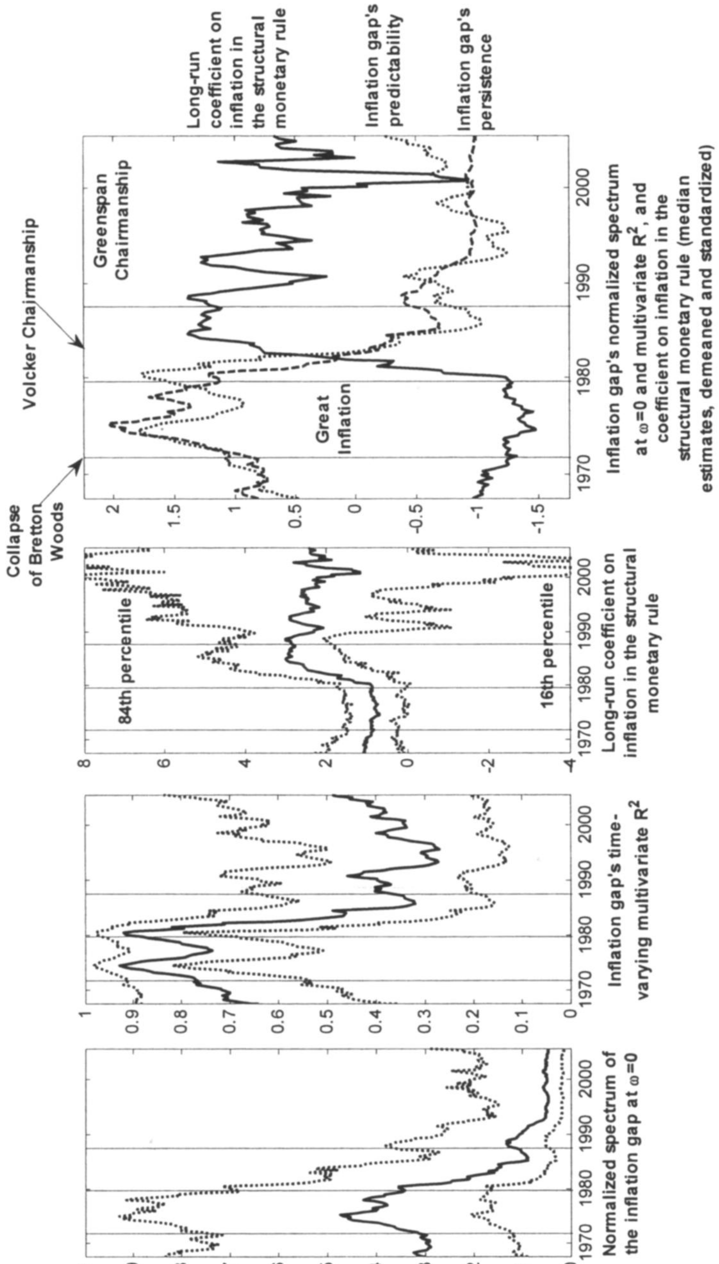

FIGURE 1. U.S. evolving monetary policy and inflation's persistence and predictability.

TABLE 1. Sign restrictions imposed on the VAR.

	Shocks			
Variable	ε_t^M	ε_t^D	ε_t^S	ε_t^{MD}
Federal Funds rate	+	+	x	+
Inflation	−	+	−	−
Output growth	−	+	+	−
M2 growth	−	+	x	+

Note: x = left unconstrained.

are sufficient to identify uniquely the four shocks.[6] We compute the time-varying structural impact matrix, $A_{0,t}$, via the procedure introduced by Rubio-Ramirez, Waggoner, and Zha (2005).

Sign restrictions are appealing here because they allow us to impose a set of theory-consistent restrictions. In contrast, a recursive identification scheme, as used for instance by Primiceri (2005), would make it difficult to interpret the SVAR evidence using the standard New-Keynesian model as we do in Section 3.

The bottom left panel of Figure 1 plots the median and the 16th and 84th percentiles of the distribution of the long-run coefficient on inflation in the structural monetary rule. Abstracting from the econometric uncertainty of the second half of the sample and focusing on median estimates, we notice that the results accord remarkably well with the narrative account of the post-WWII U.S. monetary history: The reaction of the federal funds rate to inflation after 1979 is markedly more aggressive than the reaction before that date.[7]

In the bottom right panel of Figure 1 we show the main result of this paper: The medians of the distributions of the normalized spectrum of inflation at $\omega = 0$, its time-varying R^2, and the long-run coefficient on inflation in the structural monetary rule (i.e., the black lines in the first three panels). In order to facilitate the comparison of the evolution over time, the three series have been de-meaned and standardized. A striking negative correlation between the long-run coefficient on inflation, on the one hand, and the persistence and predictability of the inflation gap, on the other, is readily apparent.

To the extent that the structural VAR correctly captures the evolution of the underlying structural relationships in the economy, this evidence suggests that the evolution of the U.S. monetary policy stance might have caused a change in

6. These restrictions are exactly the same as those used in Benati (2008a).

7. Whether the long-run coefficient on inflation exceeds 1 should be de-emphasized. As stressed by Lubik and Schorfheide (2004), (in)determinacy is a system property that depends on the interplay between all the coefficients of the model; as such, it bears no clear-cut relationship with the value taken by a single (policy or non-policy) coefficient.

both the persistence of inflation, and its extent of predictability. These findings therefore motivate an investigation of the impact of changes in the coefficients of the monetary rule within a New Keynesian model, which we perform in the next section.

3. Interpreting the Empirical Evidence

3.1. The Model

The model we use in this section is given by

$$\pi_t = \frac{\beta}{1+\alpha\beta}\pi_{t+1|t} + \frac{\alpha}{1+\alpha\beta}\pi_{t-1} + \kappa y_t + \varepsilon_{\pi,t}, \tag{12}$$

$$y_t = \gamma y_{t+1|t} + (1-\gamma)y_{t-1} - \sigma^{-1}(r_t - \pi_{t+1|t}) + \varepsilon_{y,t}, \tag{13}$$

$$r_t = \rho r_{t-1} + (1-\rho)[\varphi_\pi \pi_t + \varphi_y y_t] + \varepsilon_{r,t}, \tag{14}$$

where π_t, y_t, and r_t denote inflation, the output gap, and the Federal Funds rate, respectively.[8] The parameter $\alpha \in [0, 1]$ is price setters' extent of indexation to past inflation; $\gamma \in [0, 1]$ is the forward-looking component in the intertemporal IS curve; κ and σ are the slope of the Phillips curve and the elasticity of intertemporal substitution in consumption; ρ, φ_π, and φ_y are the smoothing parameter and the coefficients on inflation and the output gap in the monetary rule. The three structural disturbances $\varepsilon_{\pi,t}$, $\varepsilon_{y,t}$, $\varepsilon_{r,t}$ are postulated to evolve according to the AR(1) processes $\varepsilon_{x,t} = \rho_x \varepsilon_{x,t-1} + \eta_{x,t}$, with $\eta_{x,t} \sim WN(0, \sigma_x^2)$, for $x = \pi, y, r$. Because the model is log-linearized around its steady-state, it follows that π_t, y_t and r_t should be characterized as the inflation gap, the output gap, and the Federal Funds rate gap.

3.2. Bayesian Estimation

We estimate equations (12)–(14) via Bayesian methods. Note that DSGE models for the U.S. economy are routinely estimated over samples beginning in the late 1950s and early 1960s. In this paper, we restrict the estimation to the period 1983Q1–2005Q4.[9] The reason for this choice is that, if the U.S. economy was indeed in an indeterminate equilibrium before but not after October

8. The output gap is the difference between the logs of GDPC1 ('Real Gross Domestic Product, 1 Decimal'), from the *Bureau of Economic Analysis* and GDPPOT ('Real Potential Gross Domestic Product') from the *Congressional Budget Office*. The series are de-meaned before estimation.

9. Following Clarida, Galí, and Gertler (2000), we take the fourth quarter of 1982 to mark the end of the Volcker stabilization.

1979, then by estimating the model over the full sample period we would be mixing two quite different regimes, thus obtaining biased estimates of the structural parameters.[10]

Following Lubik and Schorfheide (2004) and An and Schorfheide (2007), the parameters of the model are assumed to be mutually independent. The fourth and fifth columns of Table 2 report the modes and the standard deviations of their prior densities. We maximize numerically the log posterior—defined as $\ln L(\theta \mid Y) + \ln P(\theta)$, where θ is the vector collecting the structural parameters, $L(\theta \mid Y)$ is the likelihood of θ conditional on the data, and $P(\theta)$ is the prior—via simulated annealing.[11] We generate draws from the posterior distributions of the parameters via the Random Walk Metropolis (henceforth, RWM) algorithm described in An and Schortheide (2007). In implementing the RWM algorithm, we follow An and Schorfheide (2007, Sec. 4.1) with the single exception of the method we use to calibrate the covariance matrix's scale factor—the parameter c—for which we follow the methodology described in Benati (2008b, Appendix C.2).

We run a burn-in sample of 200,000 draws which we then discard. After that, we run a sample of 500,000 draws, keeping every draw out of 10 in order to decrease the autocorrelation of the draws. In the sixth column of Table 2 we report the modes and the 90%-coverage percentiles of the posterior distributions of the parameters. The first row of Figure 2 shows the fit of the DSGE model

TABLE 2. Bayesian estimates of the structural parameters.

			Prior Distribution		Posterior Distribution:
Parameter	Domain	Density	Mode	St. dev.	mode and 90%-coverage percentiles
σ_R^2	$\mathbb{R}^+$	Gamma	1	20	0.404 [0.332; 0.569]
σ_π^2	$\mathbb{R}^+$	Gamma	5	20	0.293 [0.227; 0.398]
σ_y^2	$\mathbb{R}^+$	Gamma	2	20	0.154 [0.117; 0.223]
κ	$\mathbb{R}^+$	Gamma	0.05	0.01	0.031 [0.025; 0.048]
σ	$\mathbb{R}^+$	Gamma	10	5	28.312 [20.909; 36.581]
α	[0, 1]	Beta	0.75	0.05	0.698 [0.614; 0.794]
γ	[0, 1]	Beta	0.25	0.05	0.521 [0.496; 0.557]
ρ	[0, 1)	Beta	0.8	0.05	0.811 [0.779; 0.858]
φ_π	$\mathbb{R}^+$	Gamma	1.5	0.25	1.924 [1.558; 2.344]
φ_y	$\mathbb{R}^+$	Gamma	0.5	0.25	0.558 [0.306; 0.905]
ρ_π	[0, 1)	Beta	0.25	0.05	0.321 [0.222; 0.384]
ρ_y	[0, 1)	Beta	0.25	0.05	0.203 [0.148; 0.288]
ρ_r	[0, 1)	Beta	0.25	0.05	0.300 [0.222; 0.393]

10. On artificial data, Surico (2006) shows that pooling the observations generated under the indeterminacy and determinacy regimes produces upward biased estimates of the backward-looking component of the Phillips curve. A similar argument can be made for the IS schedule.

11. We implement simulated annealing via the algorithm proposed by Corana et al. (1987), setting the key parameters as in Goffe, Ferrier, and Rogers (1994).

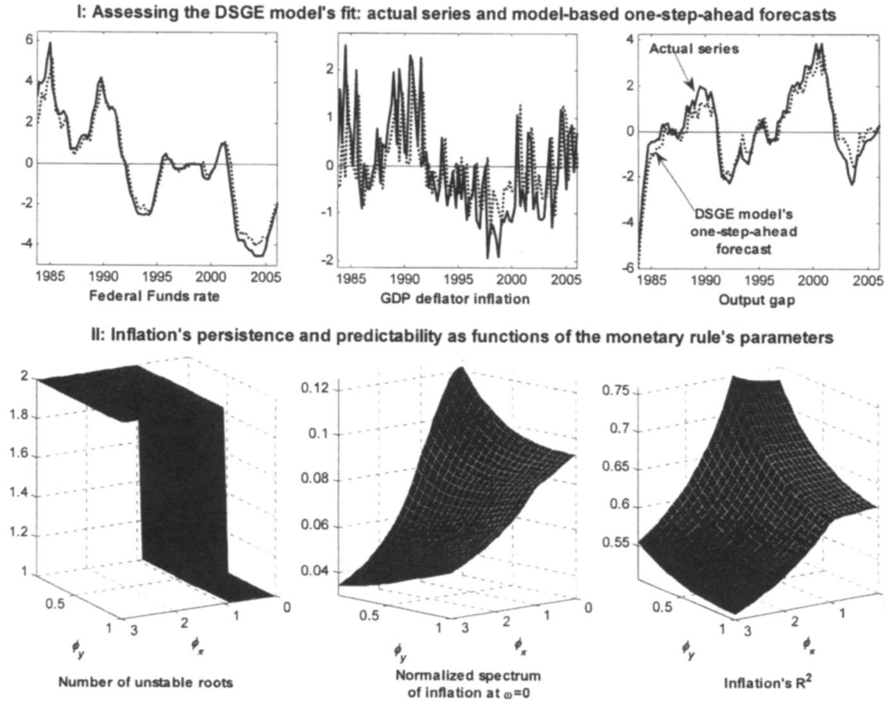

FIGURE 2. DSGE models fit and inflation's persistence and predictability as functions of the parameters of the monetary policy rule.

by plotting the actual series together with the one-step-ahead forecasts of the model.

3.3. *Monetary Policy and Inflation Gap Persistence and Predictability*

Based on the estimated DSGE model, we now explore the extent to which the persistence and predictability of inflation vary with the parameters of the monetary rule.[12] We consider two grids of values for φ_π and φ_y over the intervals [0.5, 3]

12. While revising the paper we became aware of a conceptually related earlier contribution by Adolfson and Soderstrom (2003). In that paper, the authors document changes in the reduced-form properties of the Swedish economy after the introduction of inflation targeting via time-series methods, and try to interpret them in terms of changes in the conduct of monetary policy based on a calibrated DSGE model. Key differences with the present work are: (i) their focus is on Sweden; (ii) their model is calibrated, whereas ours is estimated; (iii) as mentioned, our stylized facts are generated based on a Bayesian time-varying parameters VAR, whereas they use time-invariant methods.

and [0.25, 1].[13] For each combination of the policy parameters, we compute the theoretical spectral density of inflation by (i) expressing the DSGE model in state-space form; (ii) computing the VAR representation of the model for π_t, y_t and r_t; and (iii) Fourier-transforming the VAR using the formula in equation (10). The theoretical spectral density is then used to compute the normalized spectrum at frequency 0, which is our measure of persistence, and the R^2 in equation (11), which is our measure of predictability. The last row of Figure 2 shows, for different configurations of φ_π and φ_y, the number of explosive roots (one under indeterminacy and two under determinacy), the normalized spectrum of π_t at $\omega = 0$, and the R^2 of π_t. Several findings emerge:

- Irrespective of the specific value taken by φ_y, the persistence of the inflation gap is consistently and monotonically decreasing in φ_π under both determinacy and indeterminacy.
- Under determinacy, predictability of π_t is monotonically decreasing in φ_π, irrespective of the specific value taken by φ_y. Under indeterminacy, inflation predictability is close to invariant to changes in φ_π irrespective of the specific value taken by φ_y.

Our findings show that a shift toward a more aggressive response to inflation causes a decline in both the persistence and the predictability of inflation, thus providing a possible explanation for at least part of the decline in persistence and predictability previously identified via the time-varying VAR. The second panel of Figure 1 shows, based on median estimates, that the U.S. inflation's predictability declined from between 0.7 and 0.9 before October 1979, to between 0.3 and 0.5 after the Volcker stabilization. Conditional on the values of the coefficients of the monetary rule estimated for the post-1982 period, the extent of inflation predictability implied by the DSGE model is 0.55, about 0.15 higher than the mid point of the interval for the period following the end of the Volcker stabilization. Holding φ_y at the value estimated for the post-1982 period and decreasing φ_π, the New Keynesian model generates values for inflation's predictability of about 0.67–0.68 under indeterminacy, whereas also allowing for a decrease in φ_y further increases predictability beyond 0.7, and toward 0.75. Therefore, (i) changes in φ_π can by themselves, can replicate about one third of the overall decline in inflation's predictability; (ii) the fraction tends to one half when plausible changes in φ_y are allowed; but (iii) further additional changes in the model's structural features appear however to be needed in order to fully replicate what we see in the data.

13. The lower limits of the two grids have been purposefully chosen so as to explore also the indeterminacy region. Under indeterminacy, we solve the model using the 'continuity' identifying assumption proposed by Lubik and Schorfheide (2004).

As for persistence, the extent of variation in the normalized spectrum of inflation at $\omega = 0$ implied uniquely by changes in the coefficients in the monetary policy rule is quite small (see the second panel in the bottom row of Figure 2), and it cannot replicate what we see in the data; hence changes in additional features (again, first and foremost, indexation) are here even more important.[14]

An important point to stress, however, is that the changes shown in the second and the third panels in the bottom row of Figure 2 should be regarded as only lower bounds for the true extent of variation in persistence and predictability. Indeed, the results reported in Benati (2008b) clearly show that, historically, the adoption of more aggressively counter inflationary monetary rules has been associated with a drastic fall in inflation's indexation within a panel of countries.[15]

4. Concluding Remarks

In this paper we have provided a tentative interpretation for the decline in the persistence of the U.S. inflation gap, documented by Cogley and Sargent (2006) around the time of the Volcker disinflation, and for the decrease in the U.S. inflation predictability documented by Stock and Watson (2007) and D'Agostino, Giannone, and Surico (2006). Based on a time-varying VAR, we have uncovered a strong negative correlation between the evolution of the inflation gap persistence and inflation's predictability, and the evolution of the long-run coefficient on inflation in the structural monetary rule. We have shown that the negative correlation between the policy response on inflation and the predictability of inflation accords very well with the prediction of a standard sticky-price model.

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14. Based on a DSGE model very similar to the one used herein, Benati and Surico (2007) show that changes in the parameters of the monetary rule along the lines of Clarida, Galí, and Gertler (2000) and Lubik and Schorfheide (2004) are capable, by themselves, of replicating the dramatic fall in macroeconomic volatility associated with the transition from the Great Inflation to the Great Moderation along two key dimensions: first, the series' volatilities; second, their innovation variances. See, for example, Justiniano and Primiceri (2008) for a different view on this issue.

15. To put it differently, the indexation parameter clearly appears not to be structural in the sense of Lucas (1976).

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