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In 2007 all ECB publications feature a motif taken from the €20 banknote.

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Abstract

Using a structural VAR with time-varying parameters and stochastic volatility on post-WWII U.S. data, we document a striking 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 relative to a trend component. Using a standard sticky-price model, we show that a more aggressive policy stance towards inflation causes a decline in inflation predictability, providing a possible interpretation for the findings of the structural VAR.

Keywords: Bayesian time-varying VARs; sign restrictions; frequency domain; Great Inflation; predictability.

JEL codes: E37, E52, E58

Prepared for the invited session on 'Declining Macroeconomic Predictability' of the 2007 Meetings of the *European Economic Association*

Non Technical Summary

Changes in the data generating process for inflation have been the focus of much recent research. For the United States, in particular, Stock and Watson (2007) have documented a decrease in inflation's predictability over the most recent period, while Cogley and Sargent (2006) have identified a fall in the persistence of the 'inflation gap' around the time of the Volcker disinflation, where the gap is defined as the deviation of inflation from a slow-moving equilibrium level.

In this paper we attempt to provide a *structural* interpretation to these findings, by estimating a Bayesian time-varying parameters structural VAR with stochastic volatility for the post-WWII U.S., and then investigating the evolution of the long-run coefficient on inflation in the monetary rule, exploring in particular its co-movement with measures of the inflation gap's persistence and predictability. Given that, within our model, inflation is equal to the sum of the inflation gap and of a component evolving, to a first approximation, as a pure random walk, the fall in the inflation *gap*'s predictability automatically translates—conceptually in line with Stock and Watson (2007)—into a fall in the predictability of inflation *itself*.

We obtain three key results:

- we replicate Cogley and Sargent's finding of a fall in the persistence of the inflation gap, also in terms of timing;
- we replicate Stock and Watson's result of a decrease in U.S. inflation's predictability over the most recent period; and, crucially,
- we document a striking negative correlation between the long-run coefficient on inflation in the time-varying structural VAR's monetary rule and the previously mentioned time-varying persistence and predictability measures.

Based on a standard New Keynesian model, we show that this is *exactly* what theory predicts: a more aggressive stance towards inflation causes a fall in inflation persistence which, via standard time-series arguments, automatically translates into a decrease in its predictability. Our evidence is therefore compatible with the notion that both the fall in the persistence of the inflation gap around the time of the Volcker disinflation documented by Cogley and Sargent (2006), and the decrease in inflation's predictability over the most recent period documented by Stock and Watson (2007) may have been due to the FED's adoption, post-October 1979, of a more aggressivevely counter-inflationary stance.

1 Introduction

In a recent strand of research, Cogley and Sargent (2006) have shown that the persistence of the U.S. inflation gap, defined as the deviation of inflation from a trend component, declined remarkably around the time of the Volcker disinflation. Stock and Watson (2007) have shown that the predictability of U.S. inflation has fallen sharply over the post-1984 period. In this paper, we offer a *structural* interpretation for these two findings using a Bayesian structural VAR with time-varying parameters and stochastic volatility.

We document a striking negative correlation between the *evolution* of the long-run coefficient on inflation in the monetary rule of the structural VAR and the *evolution* of measures of persistence and predictability of U.S. inflation. Our estimates replicate the decline in the inflation gap persistence reported by Cogley and Sargent (2006), and the decline in inflation predictability reported by Stock and Watson (2007), and D'Agostino, Giannone and Surico (2006).

To interpret our results, we estimate a small scale sticky-price model and show that a more aggressive policy stance towards inflation *causes* a fall in inflation persistence and, therefore, a decrease in inflation predictability. Our evidence is consistent with the notion that the decline in both the persistence of the inflation gap and the predictability of inflation may have been due to the Fed adoption of a more aggressively counter-inflationary stance during the post-1979 period.

In Section 2, we present the results based on the structural VAR. In section 3, we present the results based on a widely used model of the business cycle.

2 Empirical evidence

2.1 A time-varying parameters VAR with stochastic volatility

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

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

where the notation is obvious, and Y_t is defined as $Y_t \equiv [r_t,\ \pi_t, y_t, m_t]'$, with r_t, π_t, y_t and m_t being the short-term interest rate, inflation, output growth and money growth.[1,2] For sake of comparability with earlier contributions, we set the lag order, p, to 2. The VAR time-varying parameters, collected in the vector θ_t, are postulated to evolve according to:

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

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

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

with $\eta_t \sim N(0,\, Q)$. The VAR reduced-form innovations in (1) are postulated to be zero-mean normally distributed, with time-varying covariance matrix Ω_t which,

[1] The source of data is as follows: 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.

[2] We add M2 growth to the other three variables in the VAR because, as discussed in Benati (2007b), under indeterminacy a VAR without money growth is mis-specified. The reason is that under indeterminacy the dynamics of the economy is driven by one additional unobserved state, for which money growth plays the role of an instrumental variable.

following established practice, we factor as

$$Var(\epsilon_t) \equiv \Omega_t = A_t^{-1} H_t (A_t^{-1})' \tag{4}$$

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

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

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

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

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 (2007a), we postulate the non-zero and non-one elements of the matrix A_t—which we collect in the vector $\alpha_t \equiv [\alpha_{21,t},\ \alpha_{31,t},\ ...,\ \alpha_{43,t}]'$—to evolve as driftless random walks,

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

and we assume the vector $[u_t',\ \eta_t',\ \tau_t',\ \nu_t']'$ to be distributed as

$$\begin{bmatrix} u_t \\ \eta_t \\ \tau_t \\ \nu_t \end{bmatrix} \sim N\left(0, V\right), \text{ with } V = \begin{bmatrix} I_4 & 0 & 0 & 0 \\ 0 & Q & 0 & 0 \\ 0 & 0 & S & 0 \\ 0 & 0 & 0 & Z \end{bmatrix} \text{ and } Z = \begin{bmatrix} \sigma_1^2 & 0 & 0 & 0 \\ 0 & \sigma_2^2 & 0 & 0 \\ 0 & 0 & \sigma_3^2 & 0 \\ 0 & 0 & 0 & \sigma_4^2 \end{bmatrix} \tag{8}$$

where u_t is such that $\epsilon_t \equiv A_t^{-1} H_t^{\frac{1}{2}} u_t$.[3] In line with Primiceri (2005), we adopt the additional simplifying assumption of a block-diagonal structure for S:

$$S \equiv \text{Var}\left(\tau_t\right) = \text{Var}\left(\tau_t\right) = \begin{bmatrix} S_1 & 0_{1\times2} & 0_{1\times3} \\ 0_{2\times1} & S_2 & 0_{2\times3} \\ 0_{3\times1} & 0_{3\times2} & S_3 \end{bmatrix} \tag{9}$$

[3] As discussed in Primiceri (2005), there are two justifications for assuming a block-diagonal structure for V_t. First, parsimony, as the model is already quite heavily parameterized. Second, 'allowing for a completely generic correlation structure among different sources of uncertainty would preclude any structural interpretation of the innovations'.

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

We estimate (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 hyperparameters 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 (2007a) 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 the Fourier-transformation of the estimated time-varying VAR:

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

where s_π is a row vector selecting inflation. Based on (10), we then compute persistence as the normalised spectrum of inflation at ω=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:

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

where F is the matrix of the VAR autoregressive coefficients in companion form.

In line with Cogley and Sargent (2006), the panels on the first row of 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.[4] As by definition inflation here is equal to the sum of the inflation gap and a trend component evolving to a first approximation as a random walk, 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 The 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 in Table 1 on the contemporaneous impacts of the structural shocks on the four endogenous variables. It can be shown that these restrictions are sufficient to identify uniquely the four shocks. We compute the time-varying structural impact matrix, $A_{0,t}$, via the procedure introduced by Rubio-Ramirez, Waggoner, and Zha (2005).[5]

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 focussing on median estimates, we notice that the results accord re-

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

[5] Specifically, let $\Omega_t = P_t D_t P_t'$ be the eigenvalue-eigenvector decomposition of the VAR time-varying covariance matrix Ω_t, and let $\tilde{A}_{0,t} \equiv P_t D_t^{\frac{1}{2}}$. We draw an $N \times N$ matrix, K, from the $N(0, 1)$ distribution, we take the QR decomposition of K—that is, we compute matrices Q and R such that $K{=}Q \cdot R$—and we compute the time-varying structural impact matrix as $A_{0,t}{=}\tilde{A}_{0,t} \cdot Q'$. If the draw satisfies the restrictions we keep it, otherwise we discard it and we keep drawing until the restrictions are satisfied, as in the Rubio-Waggoner-Zha code `SRestrictRWZalg.m`.

markably 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.[6] In the most recent period, the only two temporary drops in the long-run coefficient on inflation correspond to the 1990-1991 recession and the the collapse of the dotcom bubble.

In the bottom right panel of Figure 1, we show the main result of the paper: the medians of the distributions of the normalised spectrum of inflation at ω=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). To make easier the comparison of the evolution over time, the three series have been demeaned and standardised. 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. In the next section, we provide a possible interpretation for this result based on a simple monetary model of the business cycle.

3 Interpreting the empirical evidence

3.1 The model

The model we use in this section is given by

$$\pi_t = \frac{\beta}{1+\alpha\beta}\pi_{t+1|t} + \frac{\alpha}{1+\alpha\beta}\pi_{t-1} + \kappa y_t + \epsilon_{\pi,t} \quad (12)$$

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

$$r_t = \rho r_{t-1} + (1-\rho)[\phi_\pi \pi_t + \phi_y y_t] + \epsilon_{r,t} \quad (14)$$

[6] The fact that the long-run coefficient on inflation be—or not be—above one should be de-emphasised. As stressed by Lubik and Schorfheide (2004), (in)determinacy is a *system* property which 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.

where π_t, y_t and r_t are inflation, the output gap, and the Federal Funds rate, respectively.[7] 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—$\epsilon_{\pi,t}$, $\epsilon_{y,t}$, $\epsilon_{r,t}$—are postulated to evolve according to the AR(1) processes $\epsilon_{x,t} = \rho_x \epsilon_{x,t-1} + \eta_{x,t}$, $\eta_{x,t} \sim WN(0, \sigma_x^2)$, for $x = \pi$, y, r. As the model is log-linearised around its steady-state, π_t, y_t and r_t should be characterised as the inflation *gap*, the output gap, and the Federal Funds rate *gap*.

3.2 Bayesian estimation

We estimate (12)-(14) *via* Bayesian methods. DSGE models for the U.S. economy are routinely estimated over samples beginning in the late 1950s or early 1960s. In this paper, we restrict the estimation to the period 1983Q1-2005Q4.[8] The reason for this choice is that if the U.S. economy was indeed in an indeterminate equilibrium before, but not after, October 1979, then by estimating the model over the full sample period we would be mixing two quite different regimes obtaining biased estimates of the structural parameters.[9]

Following Lubik and Schorfheide (2004) and An and Schorfheide (2006), the pa-

[7] The output gap is the difference between the logs of GDPC1 ('Real Gross Domestic Product, 1 Decimal'), from the *Bureau of Economic Analysys*, and GDPPOT ('Real Potential Gross Domestic Product') from the *Congressional Budget Office*. The series are demeaned before estimation.

[8] Following Clarida, Gali, and Gertler (2000), we take the fourth quarter of 1982 to mark the end of the Volcker stabilisation.

[9] On artificial data, Surico (2006) shows that pooling into a full-sample 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.

rameters of the model are assumed mutually independent. The $4th$ and $5th$ columns of Table 2 reports the modes and the standard deviations of their prior densities. We maximise numerically the log posterior—defined as $ln\ L(\theta|Y) + ln\ P(\theta)$, where θ is the vector collecting the structural parameters, $L(\theta|Y)$ is the likelihood of θ conditional on the data, and $P(\theta)$ is the prior—*via* simulated annealing.[10] We generate draws from the posterior distributions of the parameters *via* the Random Walk Metropolis (henceforth, RWM) algorithm described in An and Schorfheide (2006). In implementing the RWM algorithm, we follow An and Schorfheide (2006, Section 4.1) with the single exception of the method we use to calibrate the covariance matrix's scale factor—the parameter c below—for which we follow the methodology described in Appendix C.2 of Benati (2007c).

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 $6th$ 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 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. We consider two grids of values for ϕ_π and ϕ_y over the intervals [0.5, 3] and [0.25,

[10] We implement simulated annealing *via* the algorithm proposed by Corana, Marchesi, Martini, and Ridella (1987), setting the key parameters as in Goffe, Ferrier, and Rogers (1994).

1].[11] 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 (10). The theoretical spectral density is then used to compute the normalised spectrum at frequency zero, which is our measure of persistence, and the R^2 in (11), which is our measure of predictability.

The last row of Figure 2 show, for different configurations of ϕ_π and ϕ_y, the number of explosive roots (one under indeterminacy and two under determinacy), the normalised spectrum of π_t at ω=0, and the R^2 of π_t. Several findings emerge from the three panels:

- 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 reveal that a policy shift towards a more aggressive monetary response to inflation can cause a decline in both the persistence and the predictability of inflation, thereby providing a possible interpretation for the VAR estimates of the previous section.

[11] 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).

4 Concluding remarks

In this paper, we have provided a tentative structural 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 the decrease in the U.S. inflation predictability documented by Stock and Watson (2007). Based on a time-varying VAR, we have uncovered a remarkable negative correlation between the evolution of both the inflation gap persistence and the inflation 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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Table 1: Sign restrictions imposed on the VAR

	Shock			
Variable:	ϵ_t^M	ϵ_t^D	ϵ_t^S	ϵ_t^{MD}
Federal Funds rate	+	+	x	+
inflation	−	+	−	−
output growth	−	+	+	−
M2 growth	−	+	x	+

x = left unconstrained

Table 2: Bayesian estimates of the structural parameters

			Prior distribution		Posterior distribution:
Parameter	Domain	Density	Mode	Standard deviation	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]

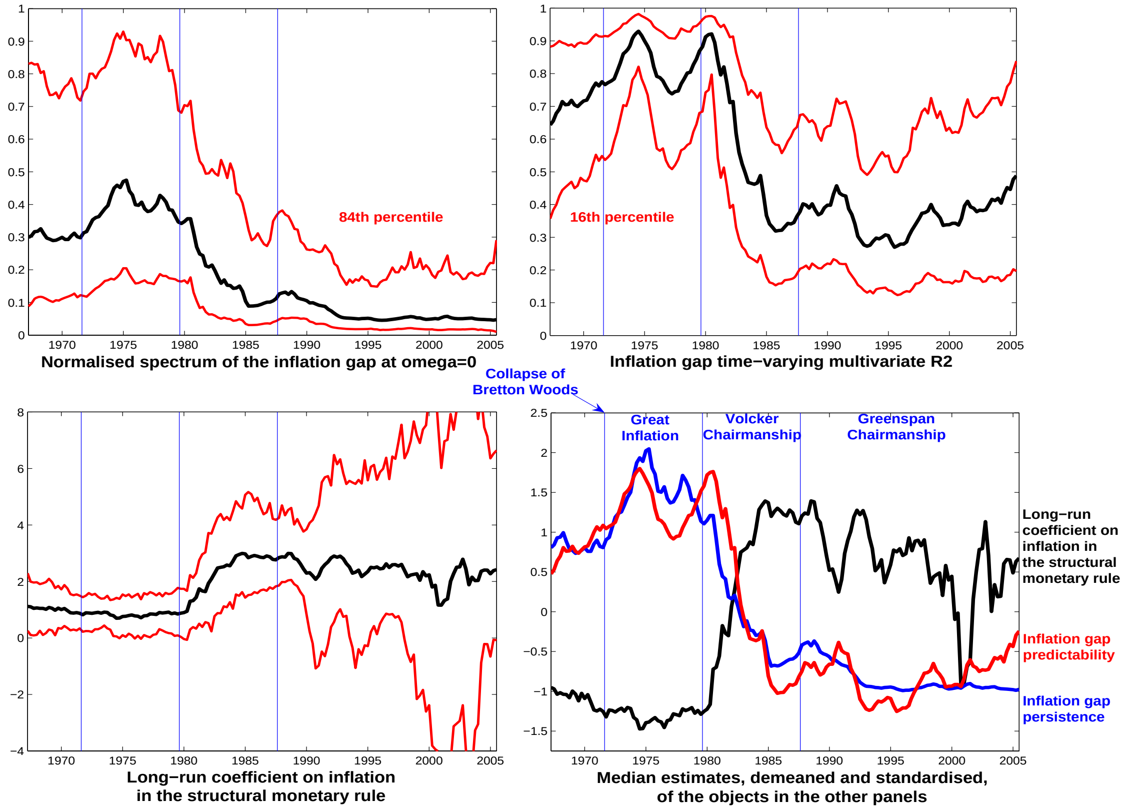

Figure 1: Evolving U.S. monetary policy, and inflation persistence and predictability

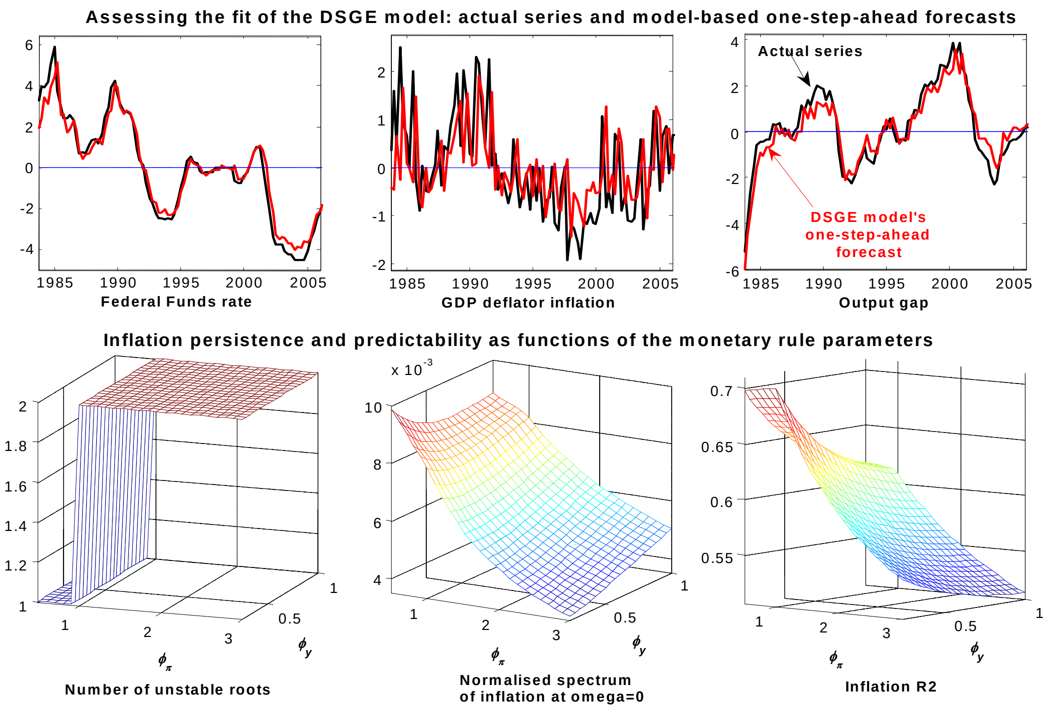

Figure 2: Fit of the DSGE model. Inflation persistence and predictability as functions of the monetary rule parameters

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