DISCUSSION PAPER SERIES

No. 4696

MONETARY MAGIC? HOW THE FED IMPROVED THE FLEXIBILITY OF THE ECONOMY

Tamim Bayoumi and Silvia Sgherri

INTERNATIONAL MACROECONOMICS

Centre for Economic Policy Research

www.cepr.org

Available online at: **www.cepr.org/pubs/dps/DP4696.asp**

ISSN 0265-8003

MONETARY MAGIC? HOW THE FED IMPROVED THE FLEXIBILITY OF THE ECONOMY

Tamim Bayoumi, International Monetary Fund (IMF) and CEPR
Silvia Sgherri, International Monetary Fund (IMF)

Discussion Paper No. 4696
October 2004

Centre for Economic Policy Research
90–98 Goswell Rd, London EC1V 7RR, UK
Tel: (44 20) 7878 2900, Fax: (44 20) 7878 2999
Email: cepr@cepr.org, Website: www.cepr.org

This Discussion Paper is issued under the auspices of the Centre's research programme in **INTERNATIONAL MACROECONOMICS**. Any opinions expressed here are those of the author(s) and not those of the Centre for Economic Policy Research. Research disseminated by CEPR may include views on policy, but the Centre itself takes no institutional policy positions.

The Centre for Economic Policy Research was established in 1983 as a private educational charity, to promote independent analysis and public discussion of open economies and the relations among them. It is pluralist and non-partisan, bringing economic research to bear on the analysis of medium- and long-run policy questions. Institutional (core) finance for the Centre has been provided through major grants from the Economic and Social Research Council, under which an ESRC Resource Centre operates within CEPR; the Esmée Fairbairn Charitable Trust; and the Bank of England. These organizations do not give prior review to the Centre's publications, nor do they necessarily endorse the views expressed therein.

These Discussion Papers often represent preliminary or incomplete work, circulated to encourage discussion and comment. Citation and use of such a paper should take account of its provisional character.

CEPR Discussion Paper No. 4696

October 2004

ABSTRACT

Monetary Magic? How the Fed Improved the Flexibility of the Economy*

Extending recent theoretical contributions on sources of inflation inertia, we argue that monetary policy uncertainty helps determine the sluggish adjustment of expectations to nominal disturbances. Estimating a model in which rational individuals learn over time about shifts in US monetary policy and the Phillips curve, we find strong evidence that this link exists. These results question the standard approach for evaluating monetary rules by assuming unchanged private sector responses, help clarify the role of monetary stability in reducing output variability in the US and elsewhere, and tell a subtle and dynamic story of the interaction between monetary policy and the supply-side of the economy.

JEL Classification: C51, E31 and E52
Keywords: inflation dynamics, Kalman filter and monetary policy

Tamim Bayoumi
International Monetary Fund
700 19th Street, NW
Washington DC, 20431
USA
Tel: (1 202) 623 6333
Fax: (1 202) 589 6333
Email: tbayoumi@imf.org

For further Discussion Papers by this author see:
www.cepr.org/pubs/new-dps/dplist.asp?authorid=116357

Silvia Sgherri
International Monetary Fund
700 19th Street, NW
Washington DC, 20431
USA
Tel: (1 202) 623 5969
Fax: (1 202) 623 6343
Email: ssgherri@imf.org

For further Discussion Papers by this author see:
www.cepr.org/pubs/new-dps/dplist.asp?authorid=139803

*The views expressed in this Paper are those of the authors and not necessarily those of the IMF. We are grateful for comments to Morten Ravn and Shigeru Iwata, participants at the CEPR conference on ‘Macroeconomic Policy Design for Monetary Unions’ and Australasian Meeting of the Econometric Society, and seminars at the IMF Institute and Federal Reserve Board. All remaining errors are our own.

Submitted 14 September 2004

"Given that the structure of an econometric model consists of optimal decision rules of economic agents, and that optimal decision rules vary systematically with changes in the structure of series relevant to the decision maker, it follows that any change in policy will systematically alter the structure of econometric models"

(Lucas, 1976 p.41)

I. Introduction

One of the most striking economic developments of recent decades has been the success in restoring low inflation in the United States and elsewhere. While this was aided by a variety of factors—including more prudent fiscal policies, structural reforms, and declining oil and commodity prices—there is a large consensus that changes in monetary policy have played a central role.[2] Estimated monetary policy rules suggest that through much of the 1970s the Federal Reserve pursued a policy that accommodated inflationary shocks in the United States, leading to instability in the economy as real rates responded perversely to inflationary disturbances.[3] This practice ended with Paul Volker's appointment as Chairman in 1979, when the policy response to expected inflation became "sufficiently" strong and monetary stability was thereby restored.

[2] This largely reflected the recognition by the public and politicians that high inflation was associated with bad economic performance, as well as the recognition by central bankers that policies aimed at systematically exploiting the short-run output/inflation tradeoff to increase output beyond potential were ineffective and self-defeating (Viñals, 2001), based on the model of Barro and Gordon (1983).

[3] See Taylor (1999) and Clarida, Galí, and Gertler (1998, 2000). Christiano and Gust (2000) emphasize that a high inflation expectations trap may arise if policy accommodates inflation as suggested by empirical estimates of the U.S. monetary policy reaction function. On the other hand, Orphanides (1998, 2000), Orphanides and Williams (2002), and McCallum (2001) argue the policy error was more related to an overemphasis on flawed estimates of the output gap.

This paper reexamines the experience with inflation and disinflation since the late-1960s focusing on the interaction between monetary policy and nominal flexibility/supply-side responses of the private sector. Recent theoretical papers have concluded that imperfect information about the future path of monetary aggregates can directly affect the speed of response of aggregate supply in models in which inflation inertia is the result of noisy signals about the future path of nominal aggregate demand (Mankiw and Reis, 2001, Woodford, 2003, Amato and Shin, 2003, based on the original insights provided by Lucas, 1979, and Phelps, 1983). We extend this model by linking the uncertainty about aggregate demand to instability in monetary policy. Because the real interest rate is an important driver of spending, this extension provides a clear potential link between the conduct of monetary policy and the degree of inertia in supply-side response. To capture the dynamics of this relationship, we estimate how individuals' perceptions of the Phillips curve (i.e., the supply function) and the monetary reaction function have evolved since the early 1970s employing Kalman filters for the learning process, and then use these results to examine the link between monetary stability and inflationary persistence.

Anticipating our conclusions, we find strong evidence that reductions in uncertainty about the path of the real interest rate do indeed produce a gradual reduction of the nominal inertia in the Phillips curve. This link helps to explain the role of the Federal Reserve in some of the recent improvements in supply-side responses of the U.S. economy, such as the fall in output volatility over recent years and the widespread belief that the costs of reducing inflation have also fallen.[4] In doing so, we question what might be termed the central dogma

[4] See Blanchard and Simon (2001), Boivin and Giannoni (2002), and Stock and Watson (2003) on the fall of output variability in the United States since the 1980s.

of modern monetary economics;[5] namely the belief that changes in monetary policies affect only the aggregate demand side of the economy.[6] This assumption has generated an extensive literature on evaluation the outcomes and robustness of alternative monetary rules.[7] While such an approach can be useful in evaluating minor changes in policy or short-term responses, our results indicate it may be seriously flawed for longer-term analysis. We would also note that although we focus on U.S. data, the increase in transparency and predictability of central bankers' behavior has been a general phenomena across countries, as has been the fall in inflationary persistence.[8] Hence, our analysis has implications across a wide range of countries.

While we are unaware of any other empirical work using the theoretical link we make, our theoretical framework can be related to a number of earlier contributions. Erceg and Levin (2003) simulate a calibrated micro-founded model with staggered contracts to suggest that realistic changes in inflation persistence can be generated by agents' inability to disentangle permanent from transitory shifts in the policy target of the central bank's reaction function (see also Gertler, 1982). On the empirical side, Cogley and Sargent (2001), using a

---

[5] The phrase "central dogma" was coined by Francis Crick in biology to describe the assumption that DNA affected RNA but not vice versa. This assumption was a useful first approximation, although it has subsequently become clear that RNA can indeed affect DNA.

[6] There is little controversy that changes in monetary policy can affect supply responses in extreme cases, such as at the end of hyper-inflations (Sargent, 1993) or over the great depression (Friedman and Schwartz, 1963).

[7] See, for example, the book edited by Taylor (1999) and the comprehensive survey in Walsh (1998).

[8] See Clarida, Galí, and Gertler (1998) on monetary policies and IMF (2002) for a more general overview.

non-linear Bayesian VAR with time-varying coefficients, provide evidence on the positive correlation between inflation inertia and the monetary authority's evolving view about the economy, but do not link the two in any systematic manner[9].

The remainder of the paper is organized as follows. Section II develops the analytical framework for analyzing the mechanism through which shifts in monetary policy predictability alter the nature of the inflation process. In Section III, we present estimates for the United States of the changing parameters of the model and provide evidence of the long-run relationship linking monetary policy uncertainty to inflation persistence. In the last section, we summarize the findings and discuss policy implications.

II. THE MODEL

A. Theory

Modern monetary models of business fluctuations are generally derived within a New-Keynesian framework allowing for price stickiness through staggered timing of price adjustment (Taylor, 1980; Calvo, 1983) or via quadratic costs of price adjustment (Rotemberg, 1996). Assuming full information and rational expectations, such pricing models give rise to an aggregate supply relation of the form:

[9] A Bayesian VAR model of U.S. monetary policy allowing for discrete regime shifts is also used in Sims (1999). For a structural time-varying-parameter model see, for instance, the seminal contribution by Kim and Nelson (1989). Fuhrer and Hooker (1993) examines the link between inflation persistence and learning about regime shifts in the parameters of a monetary reaction function by means of stochastic simulations.

$$\pi_t = \beta E(\pi_{t+1}|\Omega_t) + \gamma\left(y_t - y^*{}_t\right) + \varepsilon_t^\pi \tag{1}$$

where π_t is the quarterly change in the log of consumer prices, $\left(y_t - y^*{}_t\right)$ are the deviations of log of real GDP from its flexible-price level, and ε_t^π is a supply shock assumed to be white noise. *E* is the mathematical expectation operator , Ω_t is the full information set available in the economy at time *t*, β is the discount factor, and γ is a parameter measuring the degree of real rigidities.

On empirical grounds, this aggregate supply model fails to generate realistic degrees of *inflation* persistence and disinflation costs (Ball, 1994; Roberts, 1998, 2001). To account for more sluggish expectations adjustment to nominal shocks, equation (1) needs to be augmented with a backward looking element involving past inflation:

$$\pi_t = \lambda\pi_{t-1} + (1-\lambda)\beta E(\pi_{t+1}|\Omega_t) + \gamma\left(y_t - y^*{}_t\right) + \varepsilon_t^\pi \tag{1'}$$

Although such Phillips curves are standard in the empirical literature (Clarida, Galí, and Gertler, 1999; King and Wolman, 1999; Levin, Wieland and Williams, 1999), the theoretical justification for these additional lags has been a source of contention.[10]

[10] Buiter and Jewitt (1981) and Fuhrer and Moore (1995) argue that there is a structural interpretation using overlapping *relative* real wage contracts. Alternative approaches have assumed imperfect credibility of monetary authority's announcements (Ball, 1995) or that some agents use simple autoregressive rules of thumb to forecast inflation instead of perfectly rational expectations (Roberts, 1998; Ball, 2000; Ireland, 2000; Gali and Gertler, 1999). Departures from an optimizing-agent framework are, however, unpalatable to some involved in the microfoundation approach to macroeconomics (Rotemberg and Woodford, 1997 and 1999).

In addition, an empirical conundrum with such models is that the disinflation of the 1980s and 1990s was accompanied by an increase in the coefficient on forward-looking inflationary expectations—in other words, a fall in inflationary persistence.[11] It is difficult to see why a reduction in inflation and inflationary uncertainty would be accompanied by lower persistence in a model relying only on staggered contracts or menu costs to explain nominal inertia. Lower and more stable inflation would seem to more likely result in a lengthening of contracts, implying greater persistence in inflation. Similarly, costs of adjustment would be lower as inflation is reduced and stabilized, again implying greater inflationary persistence.

Fortunately, recent advances in theory produce a motivation for inflation inertia through another mechanism, namely imperfect information about the future path of nominal aggregate demand. As discussed in Woodford (2003) and Mankiw and Reis (2001), and building on original insights by Lucas (1972) and Phelps (1983), persistent real effects of nominal shocks can also be generated in a model that assumes that fully rational individuals have only access to noisy information about the state of nominal aggregate demand. More specifically, it is assumed that each individual receives a different and imperfect signal about changes in aggregate demand conditions. The uncertainty associated with this signal makes inflation expectations adjust only sluggishly to nominal disturbances even when rationally interpreted by economic agents, an effect which for plausible parameter values is reinforced by the anticipation that others are behaving in the same manner (Amato and Shin, 2003).[12]

[11] Erceg and Levin (2003) make the same point.

[12] Amato, Morris, and Shin (2002) discuss more extensively the impact of public information in economies with imperfect common knowledge and the signalling role of monetary policy in financial markets.

The crucial implication of such theoretical models (at least for this paper) is that they are able to explain why and how the degree of inflation inertia varies over time. Specifically, pricing models embedding imperfect information imply that the coefficient on lagged inflation (λ) rises as uncertainty about the signal rises relative to uncertainty from other sources—the signal to noise ratio. Using equation (1'), the resulting Phillips curve can be rewritten as:

$$\pi_t = \lambda(\kappa_t)\pi_{t-1} + (1-\lambda(\kappa_t))E(\pi_{t+1}|\Omega_t) + \gamma\left(y_t - y^*{}_t\right) + \varepsilon_t^{\pi} \qquad (1'')$$

where κ_t represents the portion of the overall uncertainty about the state of demand that is due to uncertainty in predicting the signal, while we have simplified the equation by assuming that β=1 to ensure money superneutrality.[13]

As the interest rate stance is an important factor in determining aggregate demand, it follows that uncertainty about the evolution of interest rate policy translates into uncertainty about the path of aggregate demand. To illustrate this link, let us posit an extremely simple aggregate demand relationship in which the output gap depends linearly on uncorrelated demand shocks and on deviations of the current short-term real interest rate from its equilibrium real rate (alternative dynamic specifications can also be considered, but add little to the intuition provided by this simple example):

[13] Structural information-based pricing models may also call for higher degree of persistence in the relevant measure of real activity. Moreover, changes in the degree of uncertainty could also lead to shifts in private agents' discount factor, an effect which is ignored here. We thank Morten Ravn for bringing these caveats to our attention.

$$\left(y_t - y^*_t\right) = \phi\left(r_t - r^*\right) + \varepsilon_t^D \tag{2}$$

where $\left(y_t - y^*_t\right)$ is the output gap as defined above, $\left(r_t - r^*\right)$ represents deviations of the ex-ante short-term real interest rate from the natural rate, ϕ<0 denotes real interest rate semi-elasticity, and $\varepsilon^D{}_t$ indicates a random disturbance which is assumed to have a fixed variance $\sigma^2_{\varepsilon^D}$. (Note that this equation can be rewritten substituting nominal for real aggregate demand and the nominal interest rate for the real one).

We will assume that at the beginning of each period *t* individuals form an opinion on the likely state of aggregate demand conditions based on their inference of the real interest rate (the signal), conditional upon available information up to time *t*-1. If recent imperfect information models are a good description of reality, it follows that the coefficient on lagged inflation in the Phillips curve (1'') should be positively related to uncertainty about the evolution of the real interest rate relative to overall uncertainty in aggregate demand (the signal-to-noise ratio or Kalman gain). In this case, the fraction of the overall (conditional) variance that is due to the (conditional) variance of the signal is given by:

$$\kappa_t \quad = \quad \frac{\sigma^2_{r_{t|t-1}}}{\sigma^2_{\tilde{\tilde{y}}_{t|t-1}}} = \frac{\phi^2\sigma^2_{r_{t|t-1}}}{\phi^2\sigma^2_{r_{t|t-1}} + \sigma^2_{\varepsilon^D}}. \tag{3}$$

In order to obtain an estimate of the uncertainty surrounding the evolution of the real interest rate, we need to describe how individuals form their opinion on the likely monetary policy stance. To do that, we model the short-term real interest rate using a monetary reaction function. Such functions generally start from a Taylor rule in which deviations of the desired real interest rate from equilibrium $\left(r - r^*\right)$ depend on the deviations of inflation from its desired level (π_t^*) and the estimated output gap (Taylor, 1989):

$$\left(r - r^*\right) = \delta_1\left(\pi - \pi^*\right) + \delta_2\left(y - y^*\right) \tag{4}$$

To take account of interest rate smoothing, we use the standard approach of incorporating a partial adjustment mechanism in which the change in the interest rate depends upon the gap between the actual and desired rate (see Clarida, Galí, and Gertler, 2000).[14] This generates the following equation:

$$r_t = \rho r_{t-1} + (1-\rho)\left[r^* + \delta_1\left(\pi_t - \pi_t^*\right) + \delta_2\left(y_t - y_t^*\right)\right] + \varepsilon_t^r \tag{5}$$

where $r_t = i_t - \pi_t$, the ex post real interest rate, ρ is the smoothing parameter, and ε_t^r is a monetary policy shock. A feature of equation (5), which is a standard feature of models incorporating learning through signal extraction (see, for example, Erceg and Levin, 2003, and Cogley and Sargent, 2001), is that the conditional expectation of the real interest rate is assumed to depend only on predetermined variables and idiosyncratic disturbances, implying that the conditional variance of the real rate is only a function of uncertainty about the parameters and policy shocks. It should be stressed that we are not proxying here monetary

[14] These models generally focus on the nominal interest rate. We focus on the real rate as it seems the most relevant variable given our theoretical structure.

policy conduct with a myopic reaction function. Rather, we are assuming that, at each period *t*, rational agents revise their inference about the monetary stance in *t*+1, conditional upon Fed's observed behavior up to period *t*.

B. Estimation Strategy

To test the link between the degree of backward-looking behavior in the Phillips curve and uncertainty in monetary policy we adopt a two step procedure. First, we estimate a Phillips curve and monetary reaction function in a manner that incorporates a realistic model of optimal learning over time and produces a time-varying estimate of both variables of interest, namely the degree of persistence in inflation and the conditional variance of the expected real interest rate, that is the uncertainty associated with the prediction of the signal. We do this by using Kalman filter techniques, in which rational individuals update their estimates (as well as the variances associated to these estimates) in a Bayesian fashion. Second, we then examine our key theoretical relationship by testing the long-run relationship between the conditional variance of the real interest rate and the degree of persistence in inflation. This is done using measures of cointegration and Granger causality.

More specifically, at each period *t* we obtain basic filter estimates of two independent measurement equations, conditional upon information available up to time *t*-1:

$$\pi_t = \eta_t + \lambda_t \pi_{t-1} + \gamma_t \left(y_t - {y^*}_t \right) + \varepsilon_t^{\pi} \tag{1'''}$$

$$r_t = \mu_t + \rho_t r_{t-1} + \varphi_{1t} \pi_t + \varphi_{2t} \left(y_t - y_t^* \right) + \varepsilon_t^{r} \tag{5'}$$

where the subscripts t on the parameters $\eta, \mu, \lambda, \gamma, \rho, \varphi_1$, and φ_2 reflect the fact that they are assumed to vary over time, following independent random walk processes.[15] Hence, the transition equation describing the dynamics of the parameters of the Phillips curve ($1'''$) will be given by:

$$\begin{bmatrix} \eta_t \\ \lambda_t \\ \gamma_t \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \eta_{t-1} \\ \lambda_{t-1} \\ \gamma_{t-1} \end{bmatrix} + \begin{bmatrix} \upsilon_{1,t} \\ \upsilon_{2,t} \\ \upsilon_{3,t} \end{bmatrix}, \tag{6}$$

with $\upsilon_{i,t} \sim i.i.d.N(0,\sigma_i^2)$ for i=1,2,3, and $E(\varepsilon_t^\pi \upsilon_{i,s}) = 0$ for all t and s, and for i=1,2,3.

Similarly, the transition equation associated to the parameters of the monetary rule ($5'$) can be represented as follows:

$$\begin{bmatrix} \mu_t \\ \rho_t \\ \varphi_{1,t} \\ \varphi_{2,t} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \mu_{t-1} \\ \rho_{t-1} \\ \varphi_{1,t-1} \\ \varphi_{2,t-1} \end{bmatrix} + \begin{bmatrix} \nu_{1,t} \\ \nu_{2,t} \\ \nu_{3,t} \\ \nu_{4,t} \end{bmatrix}, \tag{7}$$

where $\nu_{i,t} \sim i.i.d.N(0,\sigma_i^2)$ for i=1,...,4, and $E(\varepsilon_t^\pi \nu_{i,s}) = 0$ for all t and s, and for i=1,...,4.

Note that the deep parameters in the original equations ($1''$) and (5) can be recovered from the linear estimating equations ($1'''$) and ($5'$). In the Phillips curve, for example, as η_t =

[15] By construction, the error terms of the two equations are assumed to be orthogonal and uncorrelated with the set of regressors. Allowing for a non-zero error covariance matrix in the presence of ARCH-type disturbance processes (see below) would have rendered Kalman filter estimation overly cumbersome.

$(1-\lambda_t)\ E_t\pi_{t+1}$, and λ_t is estimated directly in the equation, we can infer the value of the time-varying unobserved component $E_t\pi_{t+1}$. Analogously, the long-run coefficients on the monetary reaction function (δ_{1t} and δ_{2t}) can be derived from ρ_t, φ_{1t}, and φ_{2t}. Finally, if it is assumed that the natural rate of interest (r^*) is constant, the unobservable and time-varying desired level of inflation (π_t^*) can be inferred from the intercept, μ_t.[16]

A key parameter in the theoretical model is the conditional variance of next period's real interest rate. Rewriting the monetary reaction function in the generic form $r_t = X\beta_t + \varepsilon_t^r$, the conditional variance of r_t based on information available at t-1 becomes:

$$\sigma^2_{r_{t|t-1}} = X\Sigma_{t|t-1}(\beta_t)X' + \sigma^2_{t,\varepsilon_t^r} \tag{8}$$

where $\Sigma_{t|t-1}(\beta_t)$ is the conditional covariance matrix of the estimated coefficients. The conditional variance of r_t can thus be decomposed into two components: the uncertainty coming from changes in the coefficients of the monetary rule and the uncertainty due to unexpected deviations from such a rule, a partition that will prove informative below.

On the presumption that uncertainty about the state of both real interest rate and inflation arises not only from a changing structure of the economy, but also from shifts in the volatility of external shocks ε_t^r and ε_t^π, we also assume that disturbances on both equations

[16] To keep the lags in the Phillips curve short, the rate of inflation in eqaution (1″) is measured using the annualized rate of quarterly CPI inflation. In the monetary policy rule, where the focus is on underlying infaltionary trends, the annual rate of CPI inflation is used.

follow a GARCH(1,1) process, so that their variances are not constant over time.[17] For ε^{π}_{t}, for example, it follows that:

$$\begin{aligned} &\varepsilon^{\pi}_{t|t-1} \sim N(0, h_t) \\ &h_t = \alpha_0 + \alpha_1 \left(\varepsilon^{\pi}_{t-1}\right)^2 + \alpha_2 h_{t-1}. \end{aligned} \tag{9}$$

The crucial step in testing our model is to examine whether the expected relationship between the coefficient on past inflation in the Phillips curve (λ_t) and the signal-to-noise ratio discussed above (κ_t) holds, at least over the long run. We accomplish this by testing for the existence of a linear cointegrating relationship:

$$\lambda_t \quad = \quad \chi_0 + \chi_1 \kappa_t \tag{10}$$

and examining whether the coefficient χ_1 is positive and statistically significant, and whether changes in monetary policy uncertainty lead to changes in supply-side behavior rather than vice versa.[18,19]

---

[17] For the linearization involved, see Harvey, Ruiz, and Sentana (1992). An alternative approach to account for regime shifts in the variance of random shocks would be to consider a time-varying coefficients monetary reaction function with Markov-switching heteroskedasticity in the disturbance term. See, for instance, Kim (1993), Sims (1999). The major difference between ARCH-type and Markov-switching heteroskedasticity is that whereas the *un*conditional variance of the forecast error is constant in the former, in the latter it is subject to shifts due to endogenous regime breaks.

[18] While we do not estimate an aggregate demand function, for empirical purposes we implicitly assume equation (2) to hold with a unit interest rate semi-elasticity and a standard distribution of demand shocks. As a result, the signal-to-noise ratio in (3) simplifies to:

(continued)

III. Results

The model described by the two independent equations ($1'''$) and ($5'$) was estimated using quarterly data starting in the first quarter of 1961. The Kalman filters are started in early 1967, with initial values for the variances of the coefficients being based on OLS estimates over the full sample and subsequently refined based on the behavior of the estimates.[20]

The output gap cannot be observed directly. This is an important issue in empirical models of monetary policy, as assumptions about the gap can matter, with some authors arguing that conventional measures underplay the role of over-reliance on inaccurate estimates of the output gap in the rise in inflation in the 1970s, and that "real time" data on the output gap (i.e. data available to policy makers at the time) paints a more convincing picture of the evolution of monetary policy.[21] To investigate the role of different approaches

---

$$\kappa_t \quad = \quad \frac{\sigma^2_{r_{t|t-1}}}{\sigma^2_{\tilde{\tilde{y}}_{t|t-1}}} = \frac{\sigma^2_{r_{t|t-1}}}{\sigma^2_{r_{t|t-1}} + 1}. \qquad (3')$$

Note that both the sign and the significance of χ_1 in (10) are unaffected by these assumptions.

[19] While the assumption of non-stationarity of the signal-to-noise ratio is conceptually questionable, the corresponding time series κ_t is found statistically indistinguishable from a I(1) process over the finite sample considered.

[20] Computation has been carried out in Gauss 5.0, by appropriate modification of Kim and Nelson's (1999) routines publicly available on the website http://www.econ.washington.edu/user/cnelson/SSMARKOV.htm.

[21] Orphanides (1998). See also Smets (1999), McCallum (2001), and Walsh (2003).

to the output gap in the analysis, three different output gap measures were used: (a) estimates of the trend and cyclical components of the log of real GDP using Kalman filters, so that permanent shifts in potential output, y^*_t, are decomposed from shocks to the transitory component of real output, the output gap $\left(y_t - y^*_t\right)$, as follows[22]:

$$\begin{aligned} y_t &= y^*_t + \left(y_t - y^*_t\right) \\ y^*_t &= g + y^*_{t-1} + \varepsilon_t^{y^*} \\ \left(y_t - y^*_t\right) &= \theta\left(y_{t-1} - y^*_{t-1}\right) + \varepsilon_t^{\tilde{y}} \end{aligned} \tag{11}$$

(b) a conventional Hodrick-Prescott filter; and (c) "real time" output gap data up to 1995, updated by the Kalman filter estimates subsequently (the two measures are similar by 1995).[23]

Detailed results of Kalman filter estimation of potential output are provided in Table 1.[24] The estimated drift (*g*) indicates that potential output increased at a rate of around ¾ percent per quarter on average over the period (3 percent per annum), whereas inference of the autocorrelation coefficient on the output gap (θ) points to highly persistent deviations from this trend. Not surprisingly, the relative magnitude of equations' standard errors shows

[22] The unobserved components are identified as in Nelson and Plosser (1982) by assuming that their shocks, $\varepsilon_t^{y^*}$ and $\varepsilon_t^{\tilde{y}}$, are mutually independent random errors with zero mean and constant variance.

[23] We are grateful to Athanasios Orphanides, who kindly provided these data to us.

[24] Estimates of trend and cyclical component of real output are obtained separately from the other equations. In this way, all three output gap measures are policy invariant and results are, thus, comparable.

that a significant portion of the quarter-to-quarter innovations in real GDP are cyclical and not permanent.

The path of the three series for the output gap and their autocorrelations are reported in Figure 1. The autocorrelation functions indicate that the alternative measures have somewhat different characteristics, with the Hodrick-Prescott filter exhibiting the lowest degree of correlation over time, followed by the Kalman filter, and the “real time” data exhibiting the most inertia, reflecting the large and highly serially correlated gaps in the 1970s. Given the size of these differences, and the controversy over the correct approach to measuring the output gap, we report results using all three approaches, denoted by KL_ (Kalman filter), HP_ (Hodrick-Prescott Filter), and ORP_ (Orphanides’s real time data).

As discussed further below, a striking feature of our results is that a similar story is obtained using three different series for the output gap. This provides us with some confidence that our results are not an artifact of the specific approach used to measure the data. In addition, the results are similar to those obtained by others, most notably Cogley and Sargent (2001), who examine changes in parameters over time using a similar (but not identical) Bayesian technique and a companion paper of ours that uses rolling regressions (Bayoumi and Sgherri, 2004), suggesting that our estimates of underlying parameters are independent on the particular econometric method employed.

The estimates of the Phillips curve suggest some striking changes in private sector behavior since 1970, particularly as regards the role of past and expected future inflation. Table 2 reports the estimated standard errors on the estimated coefficients of the Phillips curve, while Figure 2 records the corresponding basic filters, with the top panel showing the time-varying coefficient on lagged inflation, λ_t, one of the central parameters in our model.[25] This coefficient rises rapidly over the early 1970s, from under one-quarter to around three-quarters, continues at this value to the late 1970s before gradually declining through 1990 and then stabilizing again below 0.5. Such a pattern fits closely with conventional view that monetary credibility was lost rapidly in the mid-1970s and regained slowly starting in the early 1980s (a story that our own estimated monetary reaction function confirms). It fits less well to other potential explanations of changes in the inertia in the Phillips curve. For example, while deregulation of the U.S. economy over the 1980s and 1990s might help to explain the gradual decrease in nominal inertia, the rapid increase in the early 1970s appears difficult to explain using a slow moving factor such as the macroeconomic impact of structural policies.

Variation over time in the coefficient on the output gap, γ_t, reported in the middle panel, is statistically insignificant. However, it shows a similar pattern to that on past inflation, plausibly reflecting the fact that prices respond more vigorously to activity if monetary policy is more accommodative. Prices are estimated to have become more sensitive to movements in the output gap through the 1970s and less sensitive subsequently, although in the pattern after about 1982 depends somewhat on the data used to measure the output gap.

Three other features of these coefficient estimates are worth emphasizing. First, the estimated coefficient on past inflation has broadly similar values at the start and end of the period. The overall impression is of a process which is knocked out of kilter in the mid-1970s and then gradually regained its initial equilibrium subsequently. In a sense, we have returned

[25] To provide full account of different sources of parameters variation, throughout the paper we only report unrestricted estimates of our stochastic model.

to a similar level of stability as that seen at the end of the Bretton Woods exchange rate period. Second, the Orphanides data tend to imply larger movements in coefficients over the intervening period of monetary policy instability. For example, using Orphanides data, inflation dynamics approximate a unit-root process over the mid-1970s, while in the other two data sets the coefficient on lagged inflation is high but below unity. Third, despite some differences in details, different measures of the output gap tend to show a similar overall picture about changes in inflation behavior over time.

All of these features are evident in the implied series for unobserved expected inflation, $E(\pi_{t+1}|\Omega_t)$, shown in the bottom panel of the Figure.[26] Inflation starts at around 5 percent in 1970s, rise to double digit values in the mid- to late 1970s (and much higher in the Orphanides data), before falling to the low single digits by the late 1980s.

Turning to the estimated GARCH process for supply disturbances, coefficient estimates reported in Table 3 indicate a process with considerable variation over time. The average coefficient on the square of the current residual is around one-quarter and that on the lagged estimate of the variance three-quarters, suggesting a half life of one year, while the standard deviations on both terms are over one-tenth, suggesting that this process has also shifted considerably over time. These features are illustrated in Figure 3, which graphs conditional variance of inflation and decomposes it into the influence from the changing volatility in the residuals (the GARCH process) and uncertainty surrounding the coefficient estimates. The results associated with the three measures of the output gap series are

[26] Variations over time in the unobserved expected inflation are essentially related to shifts in the degree of inflation persistence, given that the estimated product $\eta_t = (1-\lambda_t)\, E_t\pi_{t+1}$ is found to be statistically indistinguishable from a time-invariant process.

extremely similar. The conditional variance lags actual inflation, peaking in the 1980s, during the process of deflation, rather than the inflationary burst of the late 1970s. These trends are dominated by changes in the estimated variance of the residuals, which accounts for over three-quarters of the total variance in inflation. However, the uncertainty in coefficients also plays a role, particularly around the response to the supply shocks in the mid-1970s and the early period of the Volker disinflation in the early 1980s, when it explains about half of uncertainty in inflation.

The estimated coefficients on the Federal Reserve's reaction function also show large changes over time corresponding to the conventional wisdom about the path of monetary policy, as well as providing some additional insights. Table 4 reports the estimated standard errors on the coefficients. In Figure 4, the upper left and right panels graph the long-run coefficients on inflation and the output gap, respectively, whereas the bottom panels report the implied desired level of inflation on the left, and the estimated smoothing parameter on the right. The coefficient on inflation illustrates the perverse monetary practice of the late-'seventies. Using the Kalman filter or the Hodrick-Prescott filter this coefficient falls below 1, implying that, over the period, the Fed accomodated inflationary shocks, by lowering real interest rates as inflation soared. Interestingly, the reaction function using the Orphanides data stays (barely) stable, supporting his view that real time data provide a more plausible picture of the Fed's behavior.[27] This is followed by a clear change in behavior after Paul Volker became chairman, with the coefficient on inflation rising rapidly to about two by the

[27] Orphanides himself emphasizes that it is also important to use real time data on inflationary expectations in the monetary policy reaction function (Orphanides, 1998). Our reaction function uses past information, but we discuss this point in detail in a companion paper (Bayoumi and Sgherri, 2003).

mid-1980s. The coefficient subsequently falls to a value of around one and one-half in the late 1990s, with a blip down in the early 1990s when the rule appears close to unstable. Cogley and Sargent (2001) found a similar result, and plausibly attribute it to the policy of "opportunistic disinflation" adopted at the time. Certainly, such an intrinsically asymmetric policy rule could be difficult to capture in a linear model like ours. Interestingly, the instability is again less pronounced using the real time data for the output gap, suggesting that this may also matter. Finally, there is a decline in the coefficient at the end of the sample period, possibly reflecting a rising concern about the deflation as the U.S. economy fell into recession, leading to a renewed focus on output rather than inflation.

Movements in the coefficient on the output gap tend to mirror those of the coefficient on inflation with an opposite sign. The coefficient is relatively stable through the 1970s before falling (and even becoming negative) early in the early 1980s, reflecting chairman Volker's focus on wringing inflation out of the system. It subsequently returns to the level of the 1970s (except for a temporary dip at the height of the late 1990s/early 2000 boom). The derived estimates of the unobservable steady state inflation are found to be statistically stable around two percent (assuming a natural real interest rate of 4 percent over the whole period), again falling somewhat during the deflation of the 1980s. Finally, estimates of the smoothing parameter show a fall over the 1970s followed by a steady rise over the 1980s and 1990s before falling recently, when the zero bound may have limited the Fed's room to maneuver.

These coefficient estimates tell the conventional story of a loss of monetary control followed by a strong disinflation, but with a number of interesting twists. One is the relatively similarity of the long-term responses to inflation and output between the early 1970s and more recently. This implies significantly more stability in the underlying rule between the end of the Bretton Woods period and the 1990s (both characterized by monetary stability) than has generally been recognized. A second interesting feature is the increase in the smoothing parameter between these two periods. Greater smoothing of interest rates places greater reliance on the expectations channel of policy, wherein individuals respond to expectations of future changes in policy rather than to those that are actually taking place. As the public has regained confidence in policy and become more forward-looking, the Fed appears to have responded by making monetary policy responses more gradual.[28] Hence, changes in private sector behavior appear to have also affected the responses of the Federal Reserve. This is a considerably more subtle interaction between monetary policy and private sector behavior than is generally acknowledged in simple pre- and post-Volker characterization of monetary policy (this issue is discussed further in Bayoumi and Sgherri, 2004).

Turning to the uncertainty associated with unexpected deviations from the described monetary policy rule, a key variable in our analysis, coefficient estimates show no significant evidence of GARCH heteroscedasticity in the disturbances to the real interest rate, even though the results differ somewhat depending on the measure of the output gap being used (Table 5). This implies that essentially all of the uncertainty about policy actions comes from limited information about the underlying parameters in the rule, with virtually no contribution from unexpected deviations from this rule, as clearly visualized in Figure 5. Put differently, uncertainty about the path of interest rates comes from an inability to discern what the underlying monetary rule is, rather than from the belief that the Federal Reserve's

[28] Woodford (1999) demonstrates conditions under which it is optimal to explicitly introduce inertia into interest rate setting. An inertial policy means that any given interest rate change has a larger impact on variables which are driven by forward-looking expectations.

actions are in some manner random—an eminently sensible result, and completely different from the results for the Phillips curve. Consequently, monetary uncertainty is at its highest during periods when the monetary rule appears to have changed rapidly, such as the mid-1970s and early 1980s, and has fallen to extremely low levels in the 1990s as the rule has stabilized and changes in interest rates have become smoother.

The final, and most crucial, step in our analysis is to examine the dynamic relationship between our proxy for inflation persistence (the coefficient on backward-looking expectations in the Phillips curve, λ_t) and our proxy for the portion of uncertainty about aggregate demand due to uncertainty about the monetary policy stance (the signal-to-noise ratio, κ_t, as defined in (3')). For each output gap measure used, Figure 6 illustrates the general co-movement of these estimated time series, with the thick line showing our measure of policy uncertainty and the thin dotted line the coefficient on backward-looking inflation in the Phillips curve. More formally, the analysis was accomplished by testing for a cointegrating relationship since 1970s in a vector autoregression model with two lags.[29] Results are reported in Table 6. As can be seen in the upper panel of the Table, trace-tests indicate the existence of a cointegration using Hodrick-Prescott and Kalman filter measures of the output gap, while no conclusive evidence is provided when using real-time data. The middle panel reports the coefficient estimates on this long-run relationship and the estimated error correction coefficients in the VAR, while providing strong support to the hypothesis of weak exogeneity of monetary policy uncertainty in the relationship describing the evolution of nominal persistence. The coefficient on the cointegrating relationship is correctly signed

[29] Residual analysis shows no evidence of significant mis-specification. Results are available upon request.

and insignificantly different from unity (not shown), while the speed of mean reversion is quite low, with the error correction terms implying an average lag of the order of 5 years. This slow process of adjustment is consistent with the common wisdom that it is not easy to lose credibility but that, once lost, it is difficult to regain the public's confidence. It also helps explain why the disinflation of the 1980s is generally regarded as involving large costs in terms output (Ball, 1994, and Roberts, 1998). Finally, as reported in the lower panel, Granger causality tests indicate that monetary policy uncertainty Granger causes inflation persistence, but there is no feedback in the other direction. These supplemental results strengthen our belief that, over the sample period under analysis, inflation inertia has fallen *because* (not simply "*as*") monetary policy transparency has increased.

IV. **Conclusions**

This paper questions the conventional wisdom that monetary policy has no impact on the supply-side of the economy. Extending recent theoretical insights using models with strong micro-economic foundations, we propose the existence of a link between monetary policy uncertainty and the degree of inflation inertia in the Phillips curve. An empirical model was first estimated using Kalman filters to chart shifts over time in the monetary policy reaction function and the Phillips curve. These results were then used to examine the empirical validity of our hypothesized link between uncertainty about the monetary stance and inflation persistence. We find strong evidence that such a connection exists. More precisely, there is a close, statistically significant, long-term link between changes in the portion of uncertainty about aggregate demand that is due uncertainty in predicting the real interest rate and the coefficient on inflation inertia in the Phillips curve, with no evidence of reverse causation. In other words, a more stable monetary policy appears to gradually make the supply response less sluggish, exactly as predicted by theory.

Linking monetary policy and supply responses has a number of important implications. First, it calls into question the large body of work that assesses monetary rules by assuming that such rules has no impact on underlying private sector behavior. While such analysis may be useful for the short-term impact of changes in monetary rules, the analysis in this paper suggests that it is fraught with difficulty as a guide to the longer-term consequences of a policy shift. Second, the inertia associated with the public first learning about the new policy rule and then incorporating it into their supply-side responses helps explain why the disinflation of the 1980s was so difficult. Third, it implies that there is a direct connection between some of the more recent improvements in the U.S. economy, such as the fall in output volatility, and the conduct of monetary policy. If a more stable monetary policy eventually makes the inflationary response of the economy less backward-looking, this reduction in inflation inertia can make the entire supply side of the economy more efficient, reducing output fluctuations. In short, it appears that some of the seemingly magical improvements in the supply side of the U.S. economy since the early 1970s can be attributed to Federal Reserve behavior. At the same time, increases in monetary stability and reductions in the variance of output are general phenomena across a wide range of countries, suggesting that this analysis has wider implications than the United States.

Our results are also able to generate a more subtle interpretation of recent monetary history than the simple, pre- and post-Volker characterization that is often emphasized. While agreeing that the loss of monetary control in the mid-1970s and focus on reducing inflation attendant with the elevation of Paul Volker to chairman of the Federal Reserve, are key events, we also find a surprising degree of similarity between the rules followed in the early 1970s and late 1990s, both periods of relative monetary stability. The one important difference we find, that there is more interest rate smoothing in the 1990s, is also informative. It suggests that the Federal Reserve has put increasing focus on the expectations channel of monetary policy. We interpret this as illustrating the dynamic relationship between monetary policy and the private sector. The disinflation triggered by Chairman Volker's policies increased public confidence in monetary policy, reducing inertia in the inflation process. This, in turn, made the expectations channel more effective. In response, the Federal Reserve under Chairman Greenspan responded to this opportunity by making its own policies more gradualist and less disruptive. In summary, in addition to the private sector behavior responding to changes in monetary policy, monetary policy makers also appear to have responded to shifts in private sector behavior.

Table 1. Estimates of the Time-Varying Parameters of U.S. Output Gap (1966:1–2002:1)

Parameters	***Model 1 (Kalman Filter)***	
$\sigma_{\tilde{y}}$	0.0085	(0.0005)
σ_{y^*}	0.0009	(0.0020)
θ	0.9229	(0.0212)
g	0.0074	(0.0002)
Log likelihood		484.6667

Figure 1. Comparing Time-Series Properties of Alternative Output Gap Measures

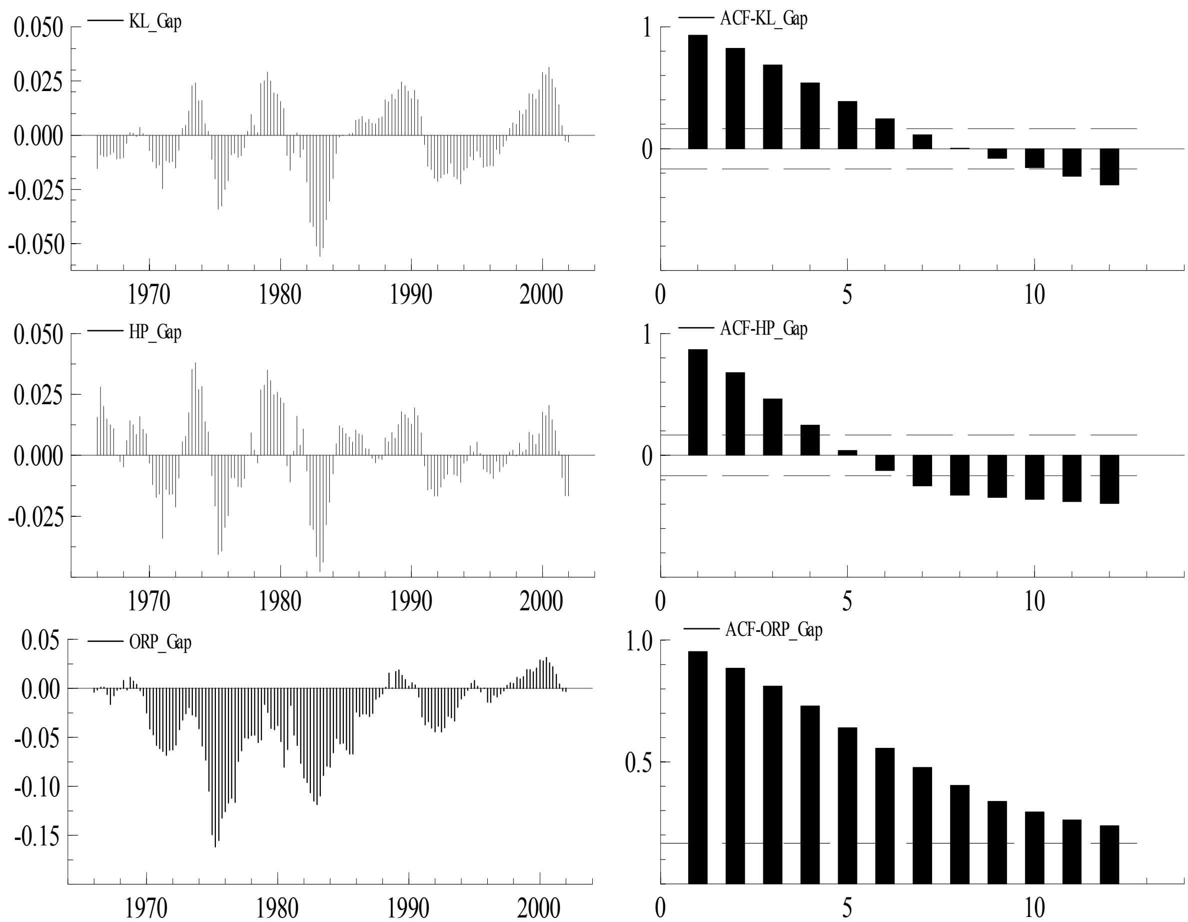

Table 2. Estimates of the Time-Varying Parameters of U.S. Phillips Curve (1967:2–2002:1)

Parameters	***Model 1 (Kalman Filter)***		***Model 2 (HP Filter)***		***Model 3 (Real-time data)***	
σ_η	0.0000	(0.1039)	0.0000	(0.1043)	0.2580	(0.2140)
σ_λ	0.0497	(0.0166)	0.0397	(0.0170)	0.0413	(0.0167)
σ_γ	0.0000	(0.0159)	0.0000	(0.0267)	0.0000	(0.0254)

Figure 2. Kalman-Filter Estimates of the Time-Varying Regression Coefficients of U.S. Phillips Curve

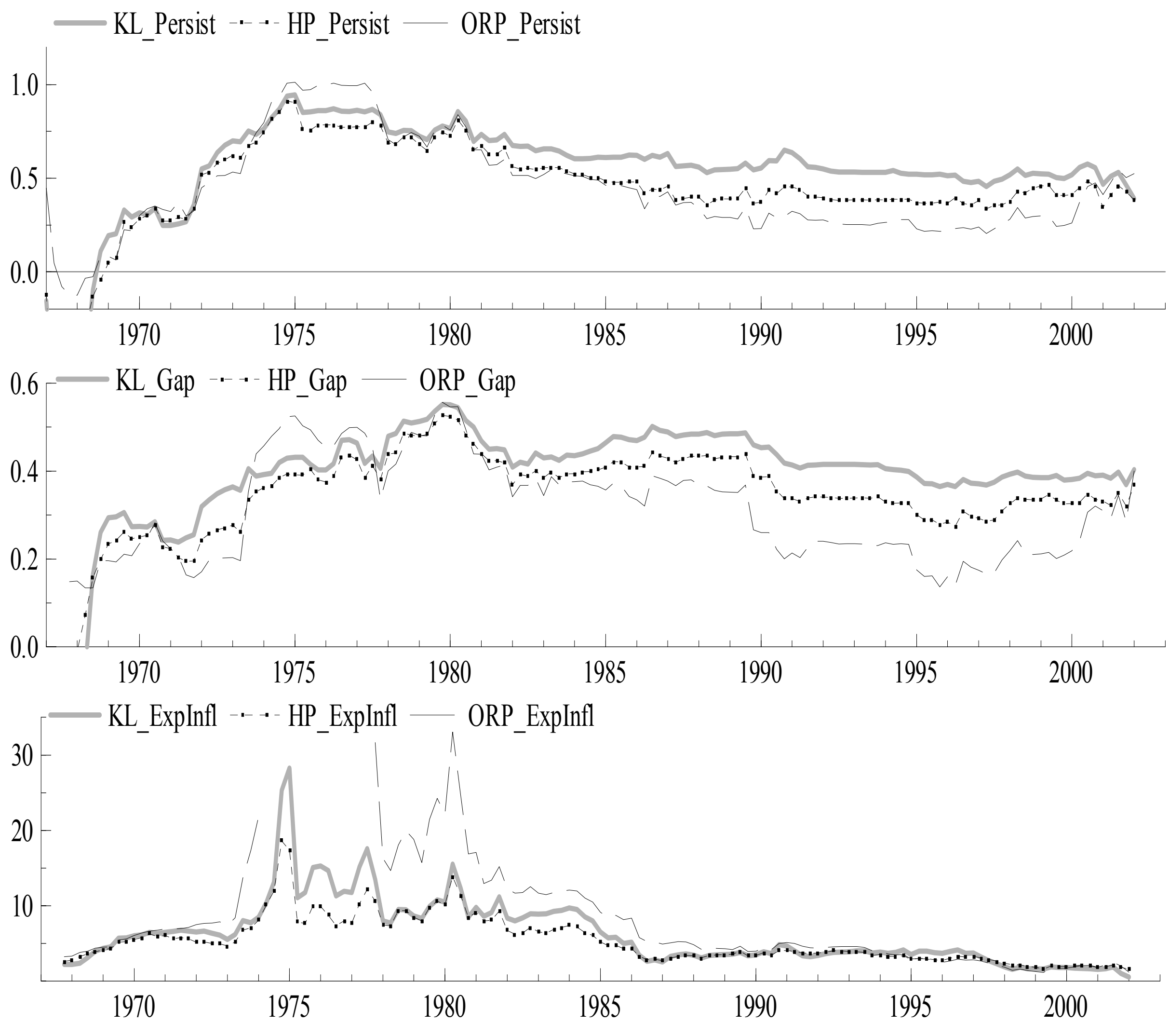

Table 3. Estimates of the GARCH Process Characterizing Disturbances to the Phillips Curve

α_0	0.0881	(0.0633)	0.0709	(0.0594)	0.0600	(0.0502)
α_1	0.1999	(0.0916)	0.2424	(0.1109)	0.3126	(0.1344)
α_2	0.7696	(0.0924)	0.7472	(0.1024)	0.6839	(0.1312)

Figure 3. Decomposition of Inflation Conditional Variance

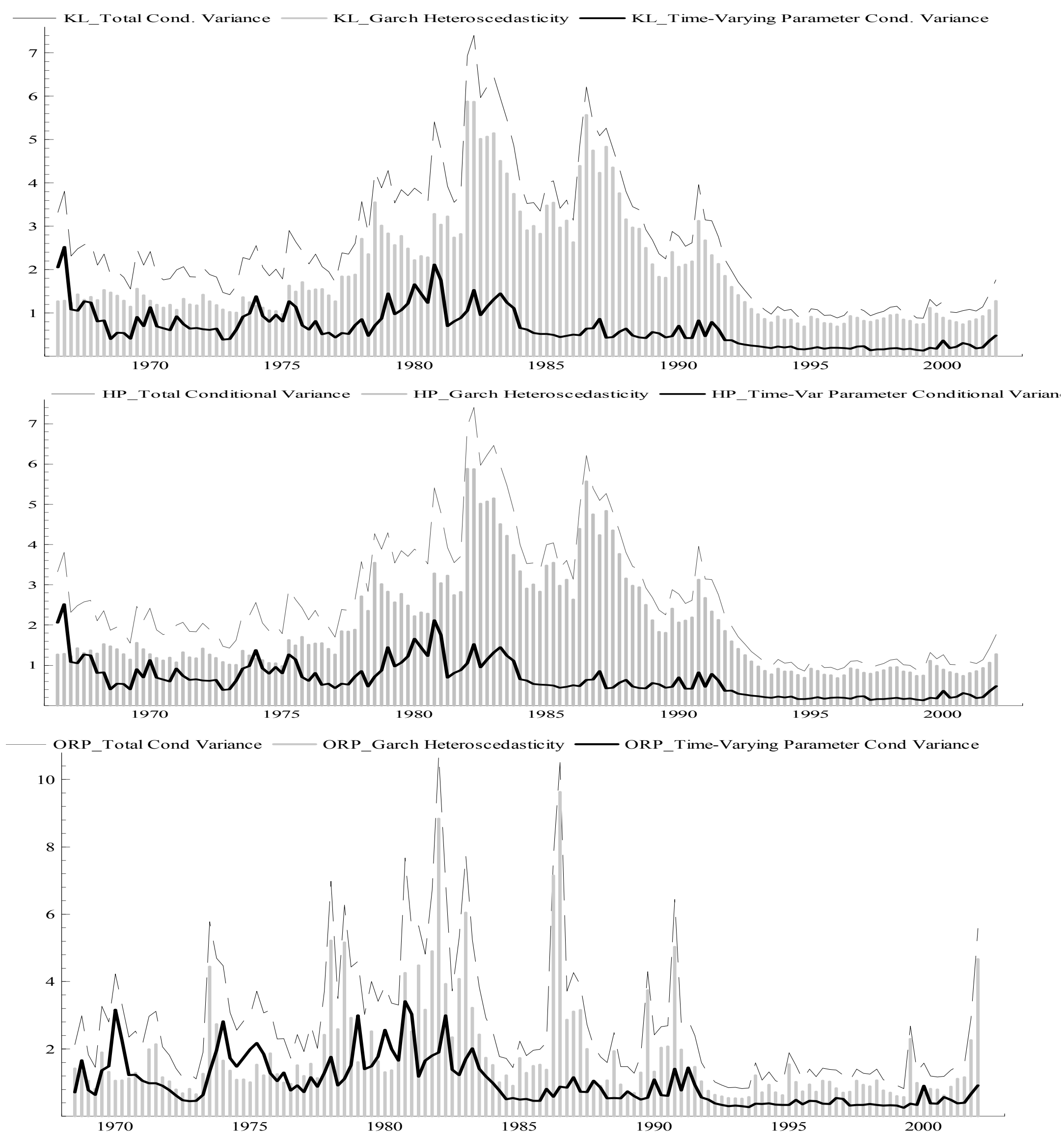

Table 4. Estimates of the Time-Varying Parameters of U.S. Monetary Rule (1967:2–2002:1)

Parameters	***Model 1 (Kalman Filter)***		***Model 2 (HP Filter)***		***Model 3 (Real-time data)***	
σ_μ	0.4321	(0.2510)	0.4199	(0.3598)	0.3708	(0.2456)
σ_ρ	0.0371	(0.0187)	0.0446	(0.0312)	0.0266	(0.0173)
$\sigma_{\varphi 1}$	0.1251	(0.0158)	0.0576	(0.0373)	0.0374	(0.0213)
$\sigma_{\varphi 2}$	0.0653	(0.0353)	0.1216	(0.0211)	0.1325	(0.0140)

Figure 4. Kalman-Filter Estimates of the Time-Varying Regression Coefficients of U.S. Monetary Rule

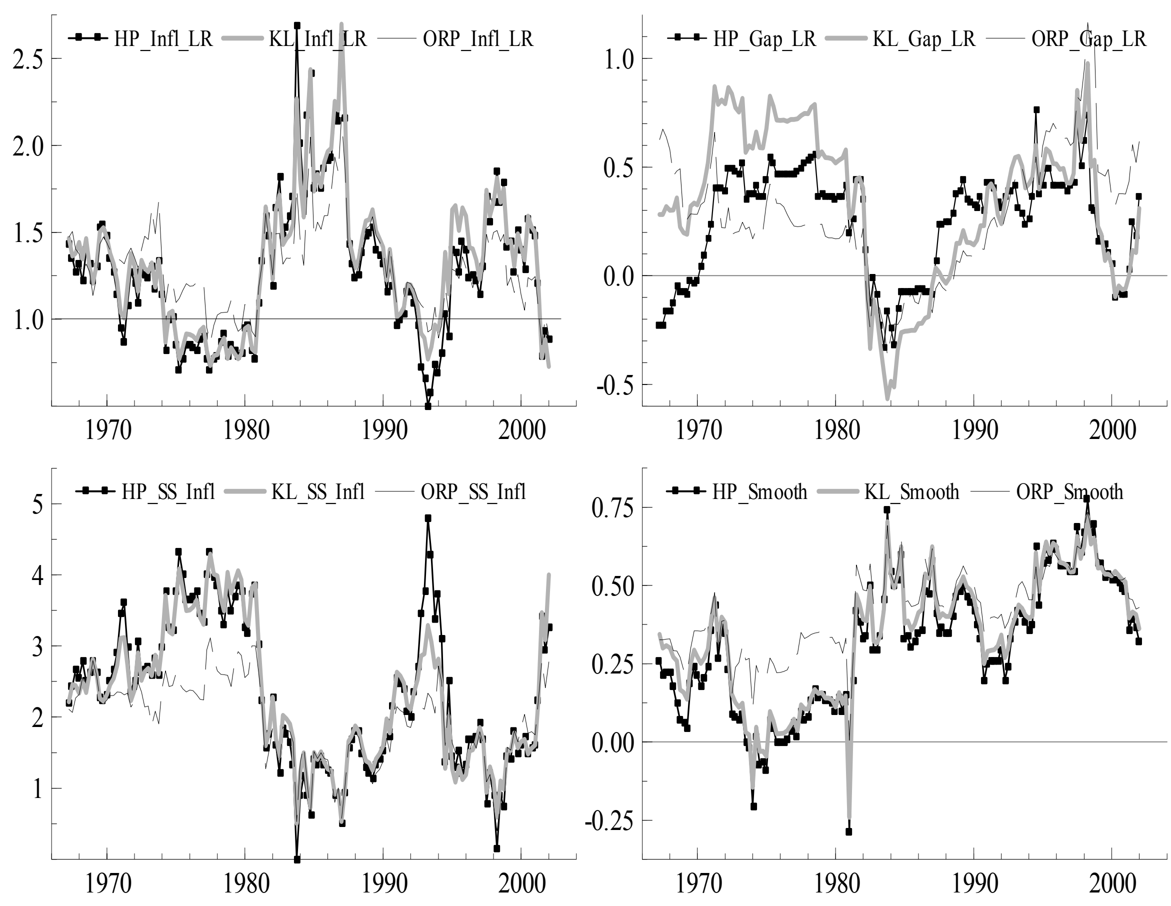

Table 5. Estimates of the GARCH Process Characterizing Disturbances to the Monetary Rule

α_0	0.0016	(0.0092)	0.0008	(0.0006)	0.0024	(0.0332)
α_1	0.1554	(0.7836)	0.2445	(1.6649)	0.0000	(0.0002)
α_2	0.3848	(0.3699)	0.5319	(2.9470)	0.8452	(2.1045)

Figure 5. Decomposition of Real Interest Rate Conditional Variance

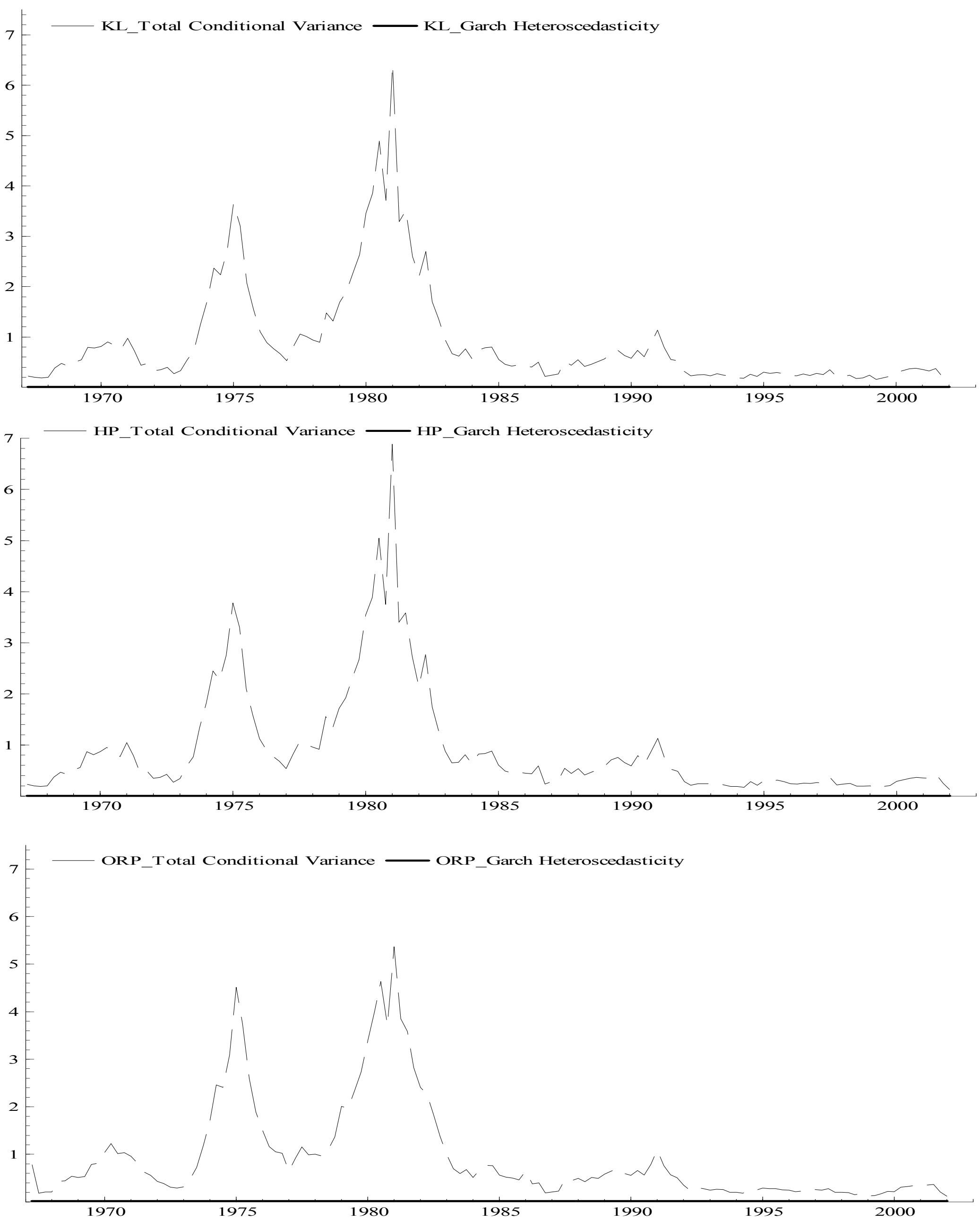

Figure 6. Time-Varying Estimates of Inflation Persistence and Relative Uncertainty About Monetary Policy Based on Three Alternative Models

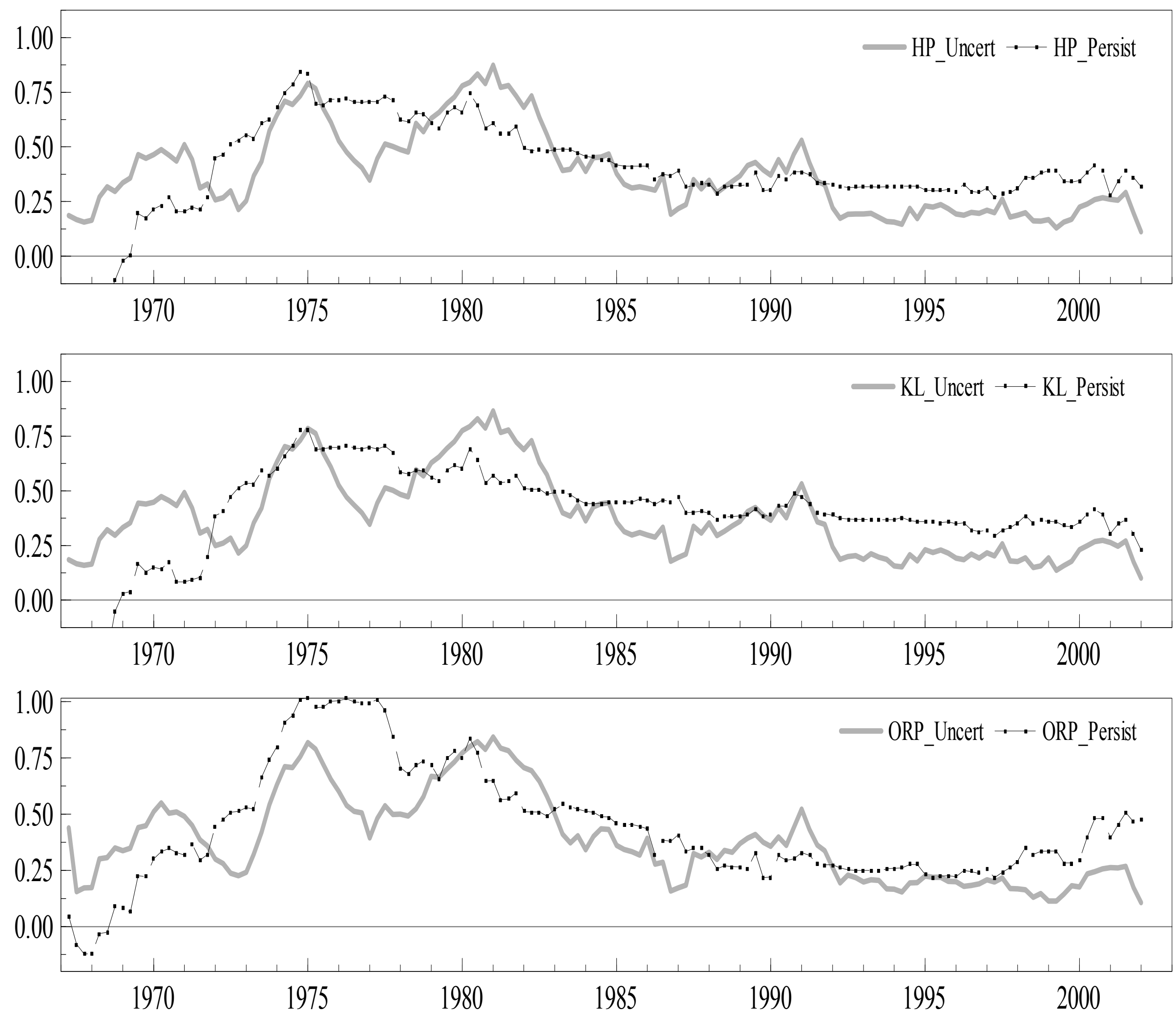

Table 6. Equilibrium Relationship Between Inflation Persistence and Relative Uncertainty About Monetary Policy: Identification and Granger Causality Test (1967:4–2002:1)

Model:

$$\Delta\lambda_t = \sum_{i=1}^{2}\delta_{i\lambda}\Delta\lambda_{t-i} + \sum_{i=1}^{2}\delta_{i\sigma}\Delta\kappa_{t-i} - \alpha_{persistence}(\lambda_{t-1} - \beta_{uncertainty}\kappa_{t-1}) + \varepsilon_t^{\lambda}$$

$$\Delta\kappa_t = \sum_{i=1}^{2}\delta_{i\lambda}\Delta\lambda_{t-i} + \sum_{i=1}^{2}\delta_{i\sigma}\Delta\kappa_{t-i} - \alpha_{uncertainty}(\lambda_{t-1} - \beta_{uncertainty}\kappa_{t-1}) + \varepsilon_t^{\kappa}$$

	Model 1 (Kalman Filter)		***Model 2 (HP Filter)***		***Model 3 (Real-time data)***	
Trace Test (T-nm)						
0	19.15	[0.00]**	16.61	[0.01]**	5.21	[0.54]
1	0.45	[0.57]	0.52	[0.54]	0.23	[0.70]
Test for weak exogeneity of relative uncertainty about monetary policy (identification of the cointegrating vector)						
$\beta_{persistence}$	1	--	1	--	1	--
$\beta_{uncertainty}$	-1.4116	(0.1570)	-1.2032	(0.1158)	-1.4133	(0.2652)
$\alpha_{persistence}$	-0.0657	(0.0161)	-0.0790	(0.0179)	-0.0380	(0.0190)
$\alpha_{uncertainty}$	0	--	0	--	0	--
Identification Restrictions [p-value]		0.6235 [0.43]		0.7682 [0.38]		0.9815 [0.32]
Granger Causality Test						
Persistence does not GC Uncertainty		0.4073 [0.82]		0.5252 [0.77]		2.0979 [0.35]
Uncertainty does not GC Persistence		13.461 [0.00]*		13.193 [0.00]**		8.4475 [0.01]*

References

Amato, Jeffrey D., Stephen Morris, and Hyun Song Shin, 2002, "Communication and Monetary Policy," *Oxford Review of Economic Policy*, Vol. 18, pp. 495-503.

Amato, Jeffrey D., and Hyun Song Shin, 2003, "Public and Private Information in Monetary Policy Models" BIS Working Paper No. 138 (Basel: Bank of International Settlements).

Ball, Laurence, 2000, "Near-Rationality and Inflation in Two Monetary Regimes," NBER Working Paper 7988 (Cambridge, Mass.: National Bureau of Economic Research).

———, 1995, "Disinflation and Imperfect Credibility," *Journal of Monetary Economics*, Vol. 35, pp. 5–24.

Barro, Robert, and David Gordon, 1983, "Rules, Discretion, and Reputation in a Model of Monetary Policy," *Journal of Monetary Economics*, Vol. 12, pp. 101–21

Bayoumi, Tamim, and Silvia Sgherri, 2004, "Deconstructing the Art of Central Banking," forthcoming IMF Working Paper (Washington: International Monetary Fund).

Blanchard, Olivier, and John Simon, 2001, "The Long and Large Decline in U.S. Output Volatility," *Brookings Papers on Economic Activity I*, pp. 135–64. (Washington: Brookings Institution).

Boivin, Jean, and Marc Giannoni, 2002, "Assessing Changes in the Monetary Transmission Mechanism: A VAR Approach", *Federal Reserve Bank of New York Economic Policy Review*, Vol 8, pp. 97-111

Buiter, W., Jewett, I., 1989, "Staggered Wage Setting and Relative Wage Rigidities: Variations on a Theme of Taylor," Reprinted in: *Macroeconomic Theory and Stabilization Policy*, edited by W. Buiter, (University of Michigan Press, Ann Arbor), pp. 183–99.

Calvo, Guillermo A., 1983, "Staggered Prices in a Utility Maximizing Framework," *Journal of Monetary Economics*, Vol. 12, pp. 383–88.

Clarida, Richard, Jordi Galí, and Mark Gertler, 1998, "Monetary Policy Rules In Practice: Some International Evidence," *European Economic Review*, pp. 1033–67

———, 1999, "The Science of Monetary Policy: A New Keynesian Perspective," *Journal of Economic Literature*, pp. 1661–07

———, 2000, "Monetary Policy Rules and Macroeconomic Stability: Evidence and Some Theory," *Quarterly Journal of Economics*, Vol. 115, pp. 147–80

Cogley, Timothy, and Thomas Sargent, 2001, “Evolving Post-World War II U.S. Inflation Dynamics,” in *NBER Macroeconomics Annual 2001*, edited by B. Bernanke and K. Rogoff, (Cambridge, MA: MIT Press), pp. 331–33.

Christiano, Lawrence J., and Cristopher Gust, 2000, “The Expectations Trap Hypothesis,” *Federal Reserve Bank of Chicago Economic Perspectives*, Vol. 24, pp. 21–39.

Clark, Peter, Douglas Laxton, and David Rose, 2001, "An Evaluation of Alternative Monetary Policy Rules in a Model with Capacity Constraints," *Journal of Money Credit and Banking*, Vol. 33, No. 1 (February), pp. 42–64.

Erceg, Christopher J., and Andrew T. Levin, “Imperfect Credibility and Inflation Persistence,” *Journal of Monetary Economics,* Vol. 50, pp. 915–44.

Evans, Martin, 1991, “Discovering the Link Between Inflation Rates and Inflation Uncertainty,” *Journal of Money Credit, and Banking,* Vol. 23(2), pp. 169–84.

Friedman, Milton and Anna Schwartz, 1963, *A Monetary History of the United States 1867-1960*, (NBER, Princeton University Press).

Fuhrer, Jeffrey, and Mark Hooker, 1993, “Learning About Monetary Regime Shifts in an Overlapping Wage Contract Model,” *Journal of Economic Dynamics and Control*, Vol. 17, pp. 531–53.

Fuhrer, Jeffrey, and George Moore, 1995, “Inflation Persistence,” *Quarterly Journal of Economics*, Vol. 110(1), pp. 127–60.

Galí, Jordi, and Mark Gertler, 1999, “Inflation Dynamics: A Structural Econometric Analysis,” *Journal of Monetary Economics,* Vol. 44, pp. 195–222.

Gertler, Mark, 1982. “Imperfect Information and Wage Inertia in the Business Cycle,” *Journal of Political Economy*, Vol. 5, pp. 967–87.

Harvey, Andres, Esther Ruiz, and Enrique Sentana, 1992, “Unobserved Component Time Series Models with ARCH Disturbances,” *Journal of Econometrics*, Vol. 52, pp. 129–57.

IMF, 2002, *World Economic Outlook*, April 2002, *Monetary Policy in a Low Inflation Era*, (Washington: International Monetary Fund).

Ireland, Peter, 2000, “Expectations, Credibility, and Time-Consistent Monetary Policy,” *Macroeconomic Dynamics,* Vol. 4, pp. 448–66.

Kim, Chang-Jin, 1993, “Sources of Monetary Growth Uncertainty and Economic Activity: The Time-Varying-Parameter Model with Heteroskedastic Disturbances,” *Review of Economics and Statistics*, Vol. 75, pp. 483–92.

Kim, Chang-Jin, and Charles R. Nelson, 1999, *State-Space Models with Regime Switching*, (Cambridge, MA: The MIT Press).

———, 1989, "The Time-Varying-Parameter Model for Modeling Changing Conditional Variance: The Case for the Lucas Hypothesis," *Journal of Business and Economic Statistics*, Vol. 74(4), pp. 433–40.

King, Robert, and Alexander Wolman, 1999, "What Should the Monetary Authority Do When Prices Are Sticky?," in *Monetary policy rules*, ed. by John B. Taylor, (Chicago, IL: University of Chicago Press).

Levin, Andrew, Volker Wieland, and John C. Williams, 1999, "Robustness of Simple Monetary Policy Rules under Model Uncertainty," in *Monetary policy rules*, ed. by John B. Taylor, (Chicago, IL: University of Chicago Press).

Lucas, Robert E., 1976, "Econometric Policy Evaluation: a Critique," in *The Phillips Curve and Labor Markets*, ed. by Karl Brunner and Alan Melzer, Carnegie-Rochester Series on Public Policy, Vol. I (North Holland).

———, 1972, "Expectations and the Neutrality of Money," *Journal of Economic Theory*, Vol. 4, pp. 103–24.

Mankiw, Gregory, and Ricardo Reis, 2001, "Sticky Information versus Sticky Prices: A Proposal to Replace the New Keynesian Phillips Curve," NBER Working Paper 8290 (Cambridge, Mass., National Bureau of Economic Research).

McCallum, Bennett, 2001, "Should Monetary Policy Respond Strongly to Output Gaps?," NBER Working Paper 8226 (Cambridge, Mass., National Bureau of Economic Research).

Nelson, Charles and Charles Plosser, 1982, "Trends and Random Walks in Macroeconomic Time Series: Some Evidence and Implications," *Journal of Monetary Economic*s, Vol. 10, pp.139–62.

Orphanides, Athanasios, 1998, "Monetary Policy Evaluation with Noisy Information," Finance and Economics Discussion Series, 1998-50, (Washington: Board of Governors of the Federal Reserve System).

———, 2000, "Activist Stabilization Policy and Inflation: the Taylor Rule in the 1970s," Finance and Economics Discussion Paper No. 2000-13. (Washington, DC: Board of Governors of the Federal Reserve System).

Orphanides, Athanasios, and John Williams, 2002, "Imperfect Knowledge, Inflation Expectations, and Monetary Policy," Finance and Economics Discussion Paper No. 2002-27 (Washington, DC: Board of Governors of the Federal Reserve System).

Phelps, Edmund S., 1983, "The Trouble with 'Rational Expectations' and the Problem of Inflation Stabilization," in *Individual Forecasting and Aggregate Outcomes*, edited by Robert Friedman and Edmund S. Phelps, (Cambridge, MA: Cambridge University Press).

Roberts, John, 2001, "How Well Does the New Keynesian Sticky-Price Model Fit the Data?," Finance and Economics Discussion Paper No. 2001-13 (Washington, DC: Board of Governors of the Federal Reserve System).

———, 1998. "Inflation Expectations and the Transmission of Monetary Policy," Finance and Economics Discussion Paper No. 98-43. (Washington, DC: Board of Governors of the Federal Reserve System).

Rotemberg, Julio, 1996, "Prices, Output, and Hours: An Empirical Analysis Based on a Sticky Price Model," *Journal of Monetary Economics,* Vol. 37, pp. 505–33.

Rotemberg, Julio, and Michael Woodford, 1997, "An Optimization-Based Econometric Framework for the Evaluation of Monetary Policy," in: *NBER Macroeconomics Annual*, ed. by Ben Bernanke and Julio Rotemberg. MIT Press.

Rudebusch, Glenn D., 2001, "Term Structure Evidence on Interest Rate Smoothing and Monetary Policy Inertia," Federal Reserve Bank of San Francisco, Mimeo.

———, 1995, "Federal Reserve Interest Rate, Targeting Rational Expectations, and the Term Structure," J*ournal of Monetary Economics,* Vol. 35, pp.245–74.

Sack, Brian, and Volker Wieland, 1999, "Interest Rate Smoothing and Optimal Monetary Policy: A Review of Recent Empirical Evidence," Finance and Economics Discussion Paper No. 1999-39 (Washington, DC: Board of Governors of the Federal Reserve System).

Sargent, Thomas, 1993, "The End of Four Great Inflations," in *Rational Inflations and Inflation* 2nd Edition (Harper Collins).

Sims, Christopher, 1999, "Monetary Policy, Drifts, (1999): "Drift and Breaks in Monetary Policy," Discussion paper, Princeton University, Mimeo. Paper presented at the Australasian meeting of the Econometric Society. Available at: http://www.princeton.edu/~sims/.

Stock, James, and Mark Watson, 2003, "Has the Business Cycle Changed?," Harvard University, Cambridge, Mimeo. Paper presented at the symposium sponsored by the Federal Reserve Bank of Kansas City. Available at: www.kc.frb.org/publicat/sympos/2003/sym03prg.htm

Svensson, Lars, 1997, "Inflation Forecast Targeting: Implementing and Monitoring Inflation Targets," *European Economic Review*, Vol. 41, pp. 1111-46.

Taylor, John B., 2001, "Expectations, Open Market Operations, and Changes in the Federal Funds Rate," in *Proceedings of the 25th Annual Economic Conference of the Federal Reserve Bank of St. Loui*s, Federal Reserve Bank of St. Louis Review, Vol. 83, pp. 33–48.

———, 1999, "Staggered Price and Wage Settings in Macroeconomics," in *Handbook of Macroeconomics*, edited by John B. Taylor and Michael Woodford, Vol. 1B, pp. 1009–50.

———, 1993, "Discretion Versus Policy Rules in Practice," *Carnegie-Rochester Conference on Public Policy,* Vol. 39, pp. 195-214.

———, 1980, "Aggregate Dynamics and Staggered Contracts," *Journal of Political Economy*, Vol. 88, pp. 1-22.

Viñals, José, 2001, "Monetary Policy Issues in a Low Inflation Environment," CEPR Working Paper No. 2945 (London, UK: Centre for Economic Policy Research).

Walsh, Carl, 2003, "Implications of a Changing Economic Structure for the Strategy of Monetary Policy," University of California, Santa Cruz, Mimeo. Paper presented at the symposium sponsored by the Federal Reserve Bank of Kansas City. Available at: www.kc.frb.org/publicat/sympos/2003/sym03prg.htm

Woodford, Michael, 2001, "Imperfect Common Knowledge and the Effects of Monetary Policy," NBER Working Paper 8673 (Cambridge, Mass., National Bureau of Economic Research).

———, 1999, "Optimal Monetary Policy Inertia," NBER Working Paper 7261 (Cambridge, Mass., National Bureau of Economic Research).