Monetary Policy Inertia or Persistent Shocks: A DSGE Analysis*

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In this paper, we propose a simple econometric framework to disentangle the respective roles of monetary policy inertia and persistent shocks in interest rate rules. We exploit the restrictions of a DSGE model that is confronted with a monetary SVAR. We show that, provided enough informative variables are included in the formal test, the data favor a monetary policy representation with modest inertia and highly serially correlated monetary shocks. To the contrary, when the procedure is based solely on the dynamic behavior of the nominal interest rate, no clear-cut conclusion can be reached about the correct representation of monetary policy.

JEL Codes: C52, E31, E32, E52.

1. Introduction

The purpose of this paper is to investigate whether the dynamics of the nominal interest rate are better described as featuring monetary policy inertia or as characterized by highly persistent factors or shocks. This paper deals with this long-debated issue by reconciling some of the earlier, inconclusive results based on single-equation

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estimates. Essentially, we show that a multivariate system brings useful information to answer this question. Using such an approach, we show that the data favor a representation of monetary policy with serially correlated shocks and modest policy inertia.

Over the recent years, there has been a renewed interest in modeling monetary policymaking in terms of simple rules. In this context, the Taylor rule has become the workhorse description of central bank behavior. Nevertheless, Taylor (1993) pointed out that one should not expect that policymakers "follow policy rules mechanically."[1] Instead, one should consider the Taylor rule as a "hypothetical but representative" description summarizing the complex process of monetary policy.

Applied monetary economists have followed this general guideline by specifying and estimating extended Taylor rules to better approximate the central bank policy. An important result obtained in the literature, as exemplified by Clarida, Galí, and Gertler (2000), is that the lagged interest rate is highly significant in the estimated policy rule, suggesting that the nominal interest rate exhibits a sizable degree of inertia.[2] It has been argued that this apparent inertia might result from deliberate *policy inertia* from the central bank, the latter enforcing a partial-adjustment process on its instrument. However, Rudebusch (2002, 2006) claimed that, if the central bank actually smoothes its policy, then future adjustments in the interest rate should be largely predictable. Unfortunately, Rudebusch (2002) showed that this is not supported by financial data (see also Söderlind, Söderström, and Vredin 2005).

An alternative explanation to the persistence of monetary policy is the presence of serially correlated shocks in the realizations of the interest rate. These shocks represent a set of special factors that cannot be systematically modeled by a simple, parsimonious interest rate rule such as an augmented Taylor rule. If these factors are persistent, the interest rate will display inertia.

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[1] See also Taylor (1999) for different possible interpretations of this monetary policy rule.

[2] This result was also found by Amato and Laubach (2003), Kozicki (1999), Levin, Wieland, and Williams (1999), and Sack and Wieland (2000), among others.

While these two competing views entail very different conclusions about the behavior of central banks, aggregate data have been fairly silent as to which is the correct representation of actual policy. For instance, Rudebusch (2002) cannot distinguish between the two competing specifications in an estimated interest rate rule. English, Nelson, and Sack (2003) find that there is supportive evidence for both representations to be significant components of the Federal Reserve behavior. Castelnuovo (2003) suggests that both views of monetary policy are equally important to describe the central bank decisions.[3] This lack of a clear-cut conclusion may be due to a well-known problem of identification and multiple optima typically arising in models of partial adjustment with serially correlated shocks. The latter calls into question the use of a single equation—i.e., a Taylor rule taken in isolation—as a proper way to discriminate between the two competing views of monetary policy.

To eschew this identification problem, we propose to resort to a dynamic stochastic general equilibrium (DSGE) model to interpret the data and disentangle these two alternative views about monetary policy. An important and celebrated virtue of such models is that they can generate very different aggregate dynamics when subjected to different policy rules, and this holds not only for the nominal interest rate but also for a broader set of macroeconomic variables. We build on this property to assess which of the two views generates aggregate dynamics in accordance with the data.

In order to implement these ideas, we resort to a limited-information approach that allows us to exclusively focus on that portion of aggregate fluctuations due to monetary shocks in U.S. data. We first estimate a structural vector autoregression (SVAR) with short-run restrictions to identify monetary policy shocks and the implied impulse response functions (IRFs) of a set of aggregate variables. Second, we estimate the DSGE parameters that govern policy inertia and the amount of serial correlation in monetary shocks. These parameters are pinned down so that the DSGE model matches as well as possible the IRFs drawn from the SVAR.

[3] Apel and Jansson (2005) and Gerlach-Kristen (2004) find similar results using Kalman filtering to account for omitted unobserved factors in the interest rate rule.

When we consider the IRFs of output, inflation, wage inflation, the federal funds rate, and money growth, we are able to unambiguously discriminate between the two different representations. Our results suggest that the dynamics of these variables are better fitted by a scheme with moderate policy inertia and a high degree of serial correlation. Thus, the smoothness in the interest rate is mainly explained by persistent factors beyond the target level of the policy rule. In contrast, when we consider only the responses of the federal funds rate, we find that there is not enough information to discriminate between the two views. Therefore, we insist that in order to disentangle the relative importance of each regime, one should take into account informative features of the data. In our case, this role is devoted to the responses of inflation and wage inflation. We also investigate two practical differences between these two alternative views about monetary policy: (i) in terms of the responsiveness of the nominal interest rate to inflation and the output gap; (ii) in terms of how the economy responds under the two rules during a specific episode, namely, the Volcker disinflation. In both cases, our findings confirm that the persistent-shocks view provides a more accurate approximation of actual monetary policy.

The remainder of the paper is organized as follows. Section 2 presents the monetary policy rule and discusses identification problems when partial adjustment and serial correlation are included. Section 3 describes the main ingredients of the DSGE model used in our empirical exercise. Section 4 presents the econometric approach employed. Section 5 discusses the main estimation results. Practical differences between policy inertia and persistent shocks are reviewed in section 6. Finally, the last section offers some concluding comments.

2. Identification Problems with Monetary Policy Rules

Following recent studies that have estimated models of central bank behavior, we postulate a monetary policy rule of the form

$$\hat{\imath}_t^\star = a_\pi \hat{\pi}_t + a_y \hat{y}_t, \tag{1}$$

$$\hat{\imath}_t = \rho_i \hat{\imath}_{t-1} + (1 - \rho_i)\hat{\imath}_t^\star + e_t, \tag{2}$$

$$e_t = \rho_e e_{t-1} + \nu_t, \quad \nu_t \sim \text{iid}\big(0, \sigma_\nu^2\big). \tag{3}$$

Equation (1) specifies how the target level $\hat{\imath}_t^\star$ evolves in response to current inflation $\hat{\pi}_t$ and output $\hat{y}_t$. More precisely, a_π and a_y govern the sensitivity of the desired level of the nominal interest rate to the log-deviations of inflation and output, respectively.[4] If the actual nominal interest rate $\hat{\imath}_t$ were equal to $\hat{\imath}_t^\star$, this would correspond to the policy rule proposed by Taylor (1993). Instead, equation (2) allows for a partial adjustment of the nominal interest rate to its target level at rate ρ_i. In addition, the rule is hit by monetary shocks e_t. If the latter were i.i.d., this would correspond to a standard specification for an augmented Taylor rule. Instead, equation (3) specifies a parametric model of serial correlation in e_t that can potentially account for part of the actual persistence found in $\hat{\imath}_t$. These shocks may represent any contingent event the central bank faces when deciding the interest rate, such as credit crunches or financial crises (see Rudebusch 2002 or Taylor 1993). Moreover, the use of real-time data could also reinforce the apparent degree of serial correlation in policy shocks (Orphanides 2004). In addition, a persistent change of the inflation target can be interpreted as a serially correlated shock to monetary policy (see Smets and Wouters 2003, 2005).

The empirical literature on Taylor rules has had trouble reaching a clear-cut conclusion about the correct representation of monetary policy. Although there is no evidence that the partial-adjustment hypothesis is fully responsible for the significance of the lagged interest rate term, there is also no evidence supporting the total rejection of monetary policy inertia.[5] The absence of a clear-cut conclusion is in part due to a well-known problem of identification and multiple optima in the partial-adjustment model with serially correlated shocks (see, e.g., Blinder 1986, Griliches 1967, Harvey 1990, and McManus, Nankervis, and Savin 1994). Rational-expectation econometrics suffer from the same problems, especially when the

[4]Without loss of generality, we omit a constant term. Notice that, here, we assume that the Taylor rule penalizes the log-deviations of output rather than those of the output gap. In the DSGE model presented in the next section, it turns out that this distinction is irrelevant because the implied natural level of output is irresponsive to monetary shocks, which are the only shocks considered in the analysis. Thus, in this framework, the output gap exactly coincides with output. See Woodford (2003, 420).

[5]See Apel and Jansson (2005), Castelnuovo (2003), English, Nelson, and Sack (2003), Gerlach-Kristen (2004), and Rudebusch (2002).

framework conveys little information, as in Keenan (1988) or Sargent (1978).

To see this problem, let us consider our simple representation of monetary policy (1)–(3):

$$\hat{\imath}_t = (\rho_i + \rho_e)\hat{\imath}_{t-1} - \rho_i\rho_e\hat{\imath}_{t-2} + (1 - \rho_i)\left(\hat{\imath}_t^\star - \rho_e\hat{\imath}_{t-1}^\star\right) + \nu_t,$$

where the target $\hat{\imath}_t^\star$ is a linear function of shocks that hit the economy. We assume for simplicity a single shock, namely, the monetary shock ν_t:

$$\hat{\imath}_t^\star = \sum_{k=0}^{\infty} \eta_k \nu_{t-k},$$

where η_k is a complicated nonlinear function of the policy rule parameters, as well as other deep parameters. Suppose that η_k for $k = 0, \ldots, \infty$ are small and not sensitive to ρ_i and ρ_e. In this case, $\hat{\imath}_t^\star$ is essentially zero with a very small amount of variance. The policy function accordingly rewrites

$$\hat{\imath}_t \approx (\rho_i + \rho_e)\hat{\imath}_{t-1} - \rho_i\rho_e\hat{\imath}_{t-2} + \nu_t.$$

In this case, the parameters ρ_i and ρ_e are not identified in general. To see this, consider the reduced form associated with the approximate monetary policy

$$\hat{\imath}_t = \beta_1\hat{\imath}_{t-1} + \beta_2\hat{\imath}_{t-2} + \nu_t.$$

Provided that $\rho_i \neq \rho_e$, there does not exist a unique solution for ρ_i and ρ_e as a function of the reduced-form parameters β_1 and β_2. Indeed, as long as $\beta_2 \neq 0$, the solutions for ρ_i and ρ_e are given by $\rho_i = (\beta_1 \pm (\beta_1^2 + 4\beta_2)^{1/2})/2$ and $\rho_e = \beta_1 - \rho_i$, where $\beta_1^2 + 4\beta_2 = (\rho_i - \rho_e)^2 \geq 0$. This means that two sets of values for ρ_i and ρ_e are observationally equivalent. The first solution is associated with the monetary-policy-inertia view (ρ_i large and ρ_e small), whereas the second is related to the persistent-shocks view (ρ_i small and ρ_e large). When $\hat{\imath}_t^\star \approx 0$, we cannot distinguish between a highly inertial monetary policy with transitory shocks and a monetary policy with small partial adjustment and highly serially correlated shocks. In contrast, if $\rho_i = \rho_e$, the parameters are identified, but this configuration is inconclusive since it assigns the same weights to both views about monetary policy.

When $\hat{\imath}_t^\star$ is responsive to shocks, and thus is more volatile, this multiple-optima problem can potentially disappear, provided that η_k is highly sensitive to perturbations in ρ_i and ρ_e. However, nothing guarantees this in practice, so that estimating ρ_i and ρ_e by focusing on the nominal interest rate only might fail to reveal the correct information about monetary policy.

A way to eschew this problem is to consider additional variables and equations in the estimation stage. We argue that moving from a single-equation setup to a system of equations is likely to aid in discriminating between the two views of policy dynamics (see Rudesbusch and Wu 2004 for a similar approach). Our strategy to identify the policy parameters rests on the restrictions imposed by a DSGE model. When the policy rule parameters have strong effects on aggregate dynamics, this gives us an opportunity to properly identify ρ_i and ρ_e and to deliver clear-cut conclusions. The next section gives a brief overview of the model used in our empirical analysis.

3. The DSGE Model

We consider a standard New Keynesian model with price and wage stickiness,[6] along the lines of Galí and Rabanal (2005) and Giannoni and Woodford (2004). Since, later on, we will seek to compare this model with a monetary SVAR à la Christiano, Eichenbaum, and Evans (1996, 1999), it is important to make sure that both models embed the same timing restrictions. To achieve this, we assume that output, inflation, and wage inflation are decided prior to observing the monetary shock, as in Rotemberg and Woodford (1997, 1999).

The first equation is the New Keynesian Phillips curve:

$$\hat{\pi}_t - \gamma_p \hat{\pi}_{t-1} = \mathrm{E}_{t-1}\left\{ \frac{(1-\alpha_p)(1-\beta\alpha_p)}{\alpha_p((1-\mu_p s_q)^{-1}(1+\theta_p \epsilon_{\mu_p}) + \theta_p \omega_p)}(\hat{w}_t + \omega_p \hat{y}_t) + \beta(\hat{\pi}_{t+1} - \gamma_p \hat{\pi}_t) \right\}, \quad (4)$$

where E_{t-1} is the expectation operator conditional on information available to the firm when reoptimizing its price. $\hat{\pi}_t$, $\hat{y}_t$, and $\hat{w}_t$

[6]See the appendix for more details about the model.

are the log-deviations of inflation, output, and real wage, respectively. The parameter $\beta \in (0,1)$ is the subjective discount factor, $\gamma_p \in [0,1]$ is the degree of indexation of prices to the most recently available inflation measure, $\alpha_p \in [0,1)$ is the degree of nominal rigidity, $s_q \in (0,1)$ represents the share of material goods, $\theta_p > 0$ is the steady-state price elasticity of demand, $\mu_p > 1$ is the steady-state markup factor, ϵ_μ is the steady-state elasticity of the markup factor, and ω_p is the real-marginal-cost elasticity with respect to the level of production.

A second set of equations defines the IS and LM curves:

$$\begin{aligned} \mathrm{E}_{t-1}\{\beta b(\hat{y}_{t+1} - b\hat{y}_t) - (\hat{y}_t - b\hat{y}_{t-1}) \\ + \sigma\chi(\hat{m}_t - \beta b\hat{m}_{t+1}) - \varphi^{-1}\hat{\lambda}_t\} = 0, \end{aligned} \tag{5}$$

$$\hat{\lambda}_t = \hat{\imath}_t + \mathrm{E}_t\{\hat{\lambda}_{t+1} - \hat{\pi}_{t+1}\}, \tag{6}$$

$$\hat{m}_t = \eta_y(\hat{y}_t - b\hat{y}_{t-1}) - \eta_i\hat{\imath}_t, \tag{7}$$

where $\hat{m}_t$, $\hat{\imath}_t$, and $\hat{\lambda}_t$ are the log-deviations of real balances, the nominal interest rate, and the representative household's marginal utility of wealth, respectively. The parameter $b \in [0,1)$ represents the degree of habit formation. The additional parameters σ, χ, φ, η_y, and η_i are deduced from the utility function. Notice that we enforce the implied constraints on these parameters when we calibrate the model. Equation (5) illustrates the role played by habits in consumption, which reinforces the backward dimension of the IS curve. Provided $\sigma\chi > 0$, this equation includes a real balance effect. Equation (6) is the standard Euler equation on bond holdings. Finally, equation (7) is the money-demand function. The difference in the information sets in equations (5) and (6) reflects the timing of decisions. Prior to observing the monetary policy shock, the household decides how much to consume and sets its nominal wage. The shock is then realized, and bond and money holdings decisions are taken.

The wage-setting equation is given by

$$\begin{aligned} \hat{\pi}_t^w - \gamma_w\hat{\pi}_{t-1} = \mathrm{E}_{t-1}\Bigg\{ & \frac{(1-\alpha_w)(1-\beta\alpha_w)}{\alpha_w(1+\omega_w\theta_w)}(\omega_w\phi\hat{y}_t - \hat{\lambda}_t - \hat{w}_t) \\ & + \beta\big(\hat{\pi}_{t+1}^w - \gamma_w\hat{\pi}_t\big)\Bigg\}, \end{aligned} \tag{8}$$

where $\hat{\pi}_t^w$ is the log-deviation of wage inflation. The parameter $\gamma_w \in [0,1]$ is the degree of wage indexation to the most recently available inflation measure, $\alpha_w \in [0,1)$ is the degree of nominal wage rigidity, $\theta_w > 0$ is the wage elasticity of labor demand, $\omega_w > 0$ is the elasticity of the marginal disutility of labor, and $\phi > 1$ is the inverse elasticity of output with respect to the labor input. Finally, $\hat{\pi}_t$ and $\hat{\pi}_t^w$ are linked together through the relation

$$\hat{\pi}_t^w = \hat{w}_t - \hat{w}_{t-1} + \hat{\pi}_t. \tag{9}$$

The model is closed by postulating the monetary policy rule (1)–(3).

4. Econometric Approach

This section details our monetary SVAR and the implied IRFs used to estimate the DSGE model, and presents the minimum distance estimation (MDE) approach.

4.1 The Monetary SVAR

We start our analysis by characterizing the actual economy's response to a monetary policy shock. As is now standard, this is done by estimating a monetary SVAR in the lines of Christiano, Eichenbaum, and Evans (1996, 1999) so as to identify monetary policy shocks.[7] We consider a structural VAR of the form

$$A_0 Z_t = A_1 Z_{t-1} + \cdots + A_\ell Z_{t-\ell} + \eta_t,$$

where the data vector Z_t can be decomposed according to $Z_t = (Z'_{1,t}, \hat{\imath}_t, Z'_{2,t})'$. $Z_{1,t}$ is an $n_1 \times 1$ vector composed of variables whose current and past realizations are included in the information set available to the policymaker at t and that are assumed to be predetermined with respect to the monetary shock ϵ_t. $Z_{2,t}$ is an $n_2 \times 1$ vector containing variables that are allowed to respond contemporaneously to ϵ_t but whose value is unknown to monetary policy authorities at t. The lag length ℓ is determined by minimizing the Hannan-Quinn information criterion. In our empirical analysis, we found that $\ell = 4$.

[7]See also Christiano, Eichenbaum, and Evans (1997, 2005) and Rotemberg and Woodford (1997, 1999) for other examples of this identifying strategy.

4.2 Minimum Distance Estimation

Let ψ denote the whole set of model parameters. Let $\psi_2 = (\rho_i, \rho_e, \sigma_\nu)'$ and let ψ_1 denote the vector collecting all the remaining parameters, so that $\psi = (\psi_1', \psi_2')'$. To implement our approach, it is important that ψ_1 be fixed, so that variations in the empirical performance of the DSGE model result only from changes in ψ_2, thus revealing information about the relevant specification of the monetary policy rule.

The policy parameters ψ_2 are estimated by minimizing a measure of the distance between the empirical responses of key aggregate variables and their model counterparts.[8] More precisely, we focus our attention on the responses of the vector X_t, regrouping the actual data that we are interested in. Here, X_t is a subset of Z_t. We define θ_j as the vector of responses of the variables in X_t to a monetary shock at horizon $j \geq 0$, as implied by the above SVAR.

Then, the object that we seek to match is $\theta = \text{vec}([\theta_0, \theta_1, \ldots, \theta_k])'$, where k is the selected horizon.[9] Then let $h(\cdot)$ denote the mapping from the structural parameters $\psi_2 = (\rho_i, \rho_e\, \sigma_\nu)'$ to the DSGE counterpart of θ. Our estimate of ψ_2 is obtained by minimizing

$$\mathcal{J}_T = (h(\psi_2) - \hat{\theta}_T)V_T(h(\psi_2) - \hat{\theta}_T)',$$

where $\hat{\theta}_T$ is an estimate of θ, T is the sample size, and V_T is a weighting matrix that we assume is the inverse of a matrix containing the asymptotic variances of each element of θ along its diagonal and zeros elsewhere. We make this particular choice for the weighting matrix to avoid singularity problems of the covariance matrix of IRFs. In addition, as suggested by Christiano, Eichenbaum, and Evans (2005), this choice of weighting matrix

[8]See Altig et al. (2004), Amato and Laubach (2003), Boivin and Giannoni (2006), Christiano, Eichenbaum, and Evans (2005), Giannoni and Woodford (2004), and Rotemberg and Woodford (1997). Following these studies, we implicitly assume that the SVAR is able to identify the structural monetary-policy reaction function, which can differ from the reaction function in the DSGE model (see Rudebusch 1998).

[9]Notice that we have to exclude from θ_0 the responses corresponding to the elements in X_t that belong to the information set available to the policymaker at t. It is important to emphasize that the DSGE model previously expounded embeds the same exclusion restrictions as the monetary SVAR.

ensures that the model-based IRFs lie as much as possible inside the confidence interval of the SVAR-based IRFs. Under the null hypothesis that the DSGE model is true, $\mathcal{J}_T$ asymptotically follows a chi-squared distribution with $\dim(\hat{\theta}_T) - \dim(\psi_2)$ degrees of freedom. We will use the statistic $\mathcal{J}_T$ as a discriminating criterion between the two representations of monetary policy. Additionally, we decompose $\mathcal{J}_T$ into components pertaining to each element of X_t. This decomposition provides a simple diagnostic tool that allows us to locate on which dimension the model succeeds or fails to replicate the IRFs implied by the SVAR.

5. Empirical Results

In this section, we first present our data and results drawn from our SVAR analysis. Second, we discuss the calibration of the model's parameters. Third, we present our estimation results. Finally, we provide a sensitivity analysis to calibration.

5.1 Data and SVAR

In addition to the federal funds rate, we use data from the nonfarm business (NFB) sector over the sample period 1960:Q1–2002:Q4.[10] The variables used for estimation are the linearly detrended logarithm of per capita GDP, $\hat{y}_t$; the growth rate of GDP's implicit price deflator, $\hat{\pi}_t$; and the growth rate of nominal hourly

[10] Arguably, this sample period might be characterized by significant changes in monetary policy. As a consequence, the assumption that monetary policy can be represented by a single Taylor rule is rather strong. Unfortunately, the estimated IRFs from the SVAR in the period 1985:Q1–2002:Q4 exhibit a number of pathologies. For example, output persistently rises after a contractionary monetary policy shock. In addition, the estimated IRFs are not precisely estimated, implying that estimating DSGE parameters so as to replicate these responses is meaningless. This is reminiscent of the point raised by Sims (1998) that SVARs estimated on short time series can produce very erratic IRFs. Thus we follow Christiano, Eichenbaum, and Evans (1996, 1999, 2005) and adopt a longer sample. In addition, Sims and Zha (2006) found more evidence in favor of stable dynamics with unstable disturbance variances than of clear changes in model dynamics. See also Leeper and Roush (2003) and Rudebusch and Wu (2004).

compensation, $\hat{\pi}_t^w$.[11] We also include two "information" variables in the SVAR model. First, though not formally justified by the theoretical model, the growth rate of the logarithm of the Commodity Research Bureau price index of sensitive commodities, $\hat{\pi}_t^c$, is included to mitigate the so-called price puzzle (see Christiano, Eichenbaum, and Evans 1996, 1999; Eichenbaum 1992; and Sims 1992). Second, the growth rate of M2, $\hat{\xi}_t$, is included to exploit information included in money growth.[12] To implement the identification strategy outlined above, we set $Z_{1,t} = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}_t^w, \hat{\pi}_t^c)'$ and $Z_{2,t} = (\hat{\xi}_t)$. In addition, the variables of interest are $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}_t^w, \hat{\imath}_t, \hat{\xi}_t)'$. The empirical responses of X_t are reported in figure 1, with $k = 30$. The plain line is our point estimates of the empirical responses of X_t, and the shaded areas indicate the asymptotic 95 percent confidence interval about the point estimates.

Though we focus on a different data set and a different sample period, our findings echo previous results reported by Christiano, Eichenbaum, and Evans (1996, 1997, 1999, 2005).[13] Output initially responds very little and then sharply drops, with an inverted hump pattern. Notice that the latter is precisely estimated. The response of inflation displays a persistent U-shaped profile, with a narrow confidence interval. Inflation's lowest response is reached several quarters (more than three years) after output has reached its trough. Moreover, inflation does not present a significant price puzzle in the very short run. The response of wage inflation is qualitatively similar, with a trough response slightly lagging that of inflation. As discussed in Woodford (2003), the delayed response of inflation is a key stylized fact that any monetary DSGE model should accurately mimic. The federal funds rate instantaneously increases and then gradually declines to the point where it will eventually cross the x axis before reverting back to its steady-state value. Finally, nominal money growth drops sharply and rapidly returns to its steady-state level.

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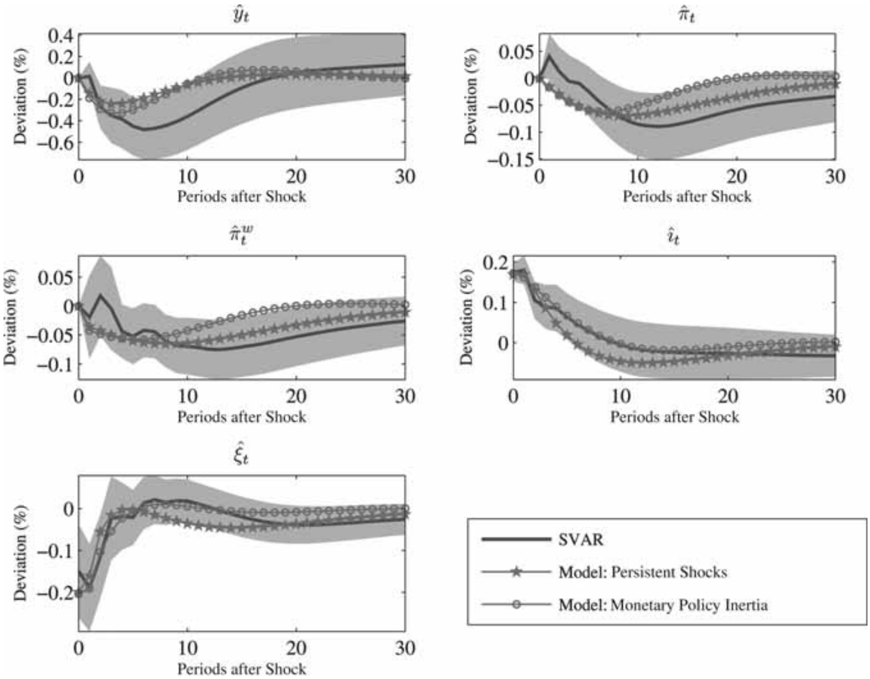
[11] The civilian non-institutional population over age sixteen is used as our measure of population. We also experimented with quadratically detrended or first-differenced output, without quantitatively altering our conclusions.

[12] The data are extracted from the Bureau of Labor Statistics web site, except for the federal funds rate and M2, which are obtained from the FREDII database.

[13] See also Bernanke and Mihov (1998), Leeper and Roush (2003), and Rotemberg and Woodford (1997, 1999) for similar IRF profiles.

Figure 1. IRFs to a Monetary Policy Shock (SVAR and DSGE Models)



5.2 Calibration

As explained above, parameters other than ψ_2 are calibrated prior to estimation. The rationale for doing this is that we want to make sure that the model's IRFs depend only on the particular specification of monetary policy. The calibration is reported in table 1.

Preferences. First, we set $\beta = 0.99$, implying a steady-state annualized real interest rate of 4 percent. The habit persistence parameter b is set to 0.75, lying in the range of available estimates based on aggregate data (see Boivin and Giannoni 2006 and Christiano, Eichenbaum, and Evans 2005). We then set $\sigma = 1 - b$, which implies intertemporal complementarities in consumption decisions (see Rotemberg and Woodford 1997).

As in Altig et al. (2004) and Christiano, Eichenbaum, and Evans (2005), the elasticity of marginal labor disutility, ω_w, is set to 1. The

Table 1. Calibrated Parameters

Parameters	**Interpretation**	**Value**
Preferences		
β	Subjective discount factor	0.99
b	Habit persistence	0.75
σ	Intertemporal elasticity of substitution $(= 1 - b)$	0.25
ω_w	Elasticity of marginal labor disutility	1.00
$\bar{v}$	Steady-state money velocity	1.36
η_y	Money-demand elasticity wrt $\hat{y}_t$	1.00
η_i	Money-demand elasticity wrt $\hat{\imath}_t$	1.18
Technology		
ϕ	Inverse of the elasticity of $\hat{y}_t$ wrt $\hat{n}_t$	1.33
ω_p	$\phi - 1$	0.33
s_q	Share of material goods	0.50
θ_p	Elasticity of demand for goods	6.00
μ_p	Markup $(= \theta_p/(\theta_p - 1))$	1.20
ϵ_μ	Markup elasticity	1.00
θ_w	Elasticity of demand for labor	21.00
μ_w	Markup $(= \theta_w/(\theta_w - 1))$	1.05
Price/Wage Setting		
γ_p	Price indexation	1.00
γ_w	Wage indexation	1.00
α_p	Probability of no price adjustment	0.66
α_w	Probability of no wage adjustment	0.66
Nominal-Interest-Rate Target Level		
a_π	Monetary policy reaction to $\hat{\pi}_t$	1.500
a_y	Monetary policy reaction to $\hat{y}_t$	0.125

money-demand function implied by our model is of the form (7). The parameter η_y governs the elasticity of real money demand to output. Following Woodford (2003), the latter is normalized to 1.

Calibrating the semi-elasticity η_i raises specific issues, especially so if the model has to reproduce the short-run behavior of money demand, as explained by Christiano, Eichenbaum, and Evans (2005). To pin down the value of η_i, we follow a different approach from theirs, yielding very similar results. From the SVAR and identified monetary shocks, we construct data series for real balances ($\tilde{m}_t$), real output ($\tilde{y}_t$), and the nominal interest rate ($\tilde{\imath}_t$) when only monetary shocks hit the SVAR. We then estimate a linear money-demand function using OLS. The estimated money demand takes the form[14]

$$\tilde{m}_t = 0.8571\tilde{m}_{t-1} + 0.1429\tilde{y}_t - 0.1072\tilde{y}_{t-1} - 1.1846\tilde{\imath}_t + \vartheta_t.$$

We use the estimated short-run semi-elasticity of money demand to the nominal interest rate (1.1846) to calibrate η_i. Notice that, in the course of estimation, we imposed $\eta_y = 1$ and took into account the calibrated value of b. The implied long-run semi-elasticity is slightly above 8, which is the value obtained by Chari, Kehoe, and McGrattan (2000), Lucas (1988), and Mankiw and Summers (1986). Consequently, our calibration of η_i must be interpreted as a way to account for the short-run response of money growth, as in Christiano, Eichenbaum, and Evans (2005).

Recall that the parameter χ governs the extent to which a real balance effect is present in our model. Under our calibration, we use the restriction $\chi = (1-\beta b)\eta_y/(\eta_i \bar{v})$. We calibrate the money velocity from actual data and obtain $\bar{v} = 1.36$. From these calibrated values, we obtain $\chi = 0.138$, implying a non-negligible real balance effect.

Technology. Here ϕ is the inverse of the elasticity of value added to labor input. We set $\phi = 1.333$, which corresponds to a steady-state share of labor income of 62.5 percent, after correcting for the markup. Assuming further that the production function is Cobb-Douglas, direct calculations yield $\omega_p = \phi - 1$. The share of material goods in gross output, s_q, is set to 50 percent,

[14]An important limit of our approach is that it assumes that OLS consistently estimates η_i. However, our estimate is not far from previous estimations. Moreover, we conduct a sensitivity analysis of our results to η_i (see section 5.4). Our findings are not qualitatively affected.

as in Basu (1995) and Rotemberg and Woodford (1995). Following Christiano, Eichenbaum, and Evans (2005) and Rotemberg and Woodford (1997), we set the markup on prices to 20 percent, i.e., $\mu_p = 1.20$. This implies an elasticity of demand for goods $\theta_p = 6$. The markup elasticity to relative demand, ϵ_μ, is set to 1, as in Bergin and Feenstra (2000) and Woodford (2003). Finally, we set θ_w to 21, as in Christiano, Eichenbaum, and Evans (2005), implying a wage markup of 5 percent.

Price/Wage Setting. Following Rotemberg and Woodford (1997), we set α_p to 0.66, implying an average spell of no price reoptimization of 2.5 quarters. This value is consistent with microeconomic evidence, e.g., Bils and Klenow (2004). We set $\gamma_p = 1$, as in Christiano, Eichenbaum, and Evans (2005). This value allows us to reinforce the backward dimension of inflation. Following Amato and Laubach (2003), we symmetrically set $\alpha_w = 0.66$. As in Christiano, Eichenbaum, and Evans (1995), we also set $\gamma_w = 1$.

Nominal-Interest-Rate Target Level. Following Taylor (1993), we set $a_\pi = 1.5$ and $a_y = 0.125$, since we focus on quarterly measures of y_t, π_t, and i_t. These values are approximately the same as those considered by Christiano, Eichenbaum, and Evans (2005) in their sensitivity analysis.

5.3 Estimation Results

The estimation results are reported in table 2 for different X_t and different restrictions on the policy rule parameters. In each case, we set the IRF horizon k to 30. The table is organized as follows: the left panel reports parameter estimates when $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}_t^w, \hat{\imath}_t, \hat{\xi}_t)'$, i.e., when ψ_2 is selected so as to reproduce the responses of output, inflation, wage inflation, the federal funds rate, and money growth to a monetary policy shock; the right panel corresponds to the case where $X_t = \hat{\imath}_t$, i.e., when we exclusively focus on the federal funds rate's behavior. In each panel, we consider five cases, depending on the minimum value of $\mathcal{J}_T$ reached at convergence and on restrictions on ρ_i or ρ_e. More precisely, column 1 corresponds to the minimum value of $\mathcal{J}_T$ reached when using as an initial condition a large ρ_e and a small ρ_i. Conversely, column 2 corresponds to the case with a large ρ_i and a small ρ_e. Column 3 corresponds to the restriction $\rho_i = 0$, i.e., to a model with only serially correlated shocks and

Table 2. Estimation Results

Parameter	Based on $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}_t^w, \hat{\imath}_t, \hat{\xi}_t)'$					Based on $X_t = (\hat{\imath}_t)$				
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**
ρ_i	0.2976	0.8986	0.0000	0.9026	0.7535	0.3018	0.8951	0.0000	0.9091	0.5606
	(0.061)	(0.009)	(*)	(0.007)	(0.001)	(0.079)	(0.049)	(*)	(0.035)	(0.002)
ρ_e	0.8740	0.0676	0.8900	0.0000		0.8207	0.0440	0.8383	0.0000	
	(0.006)	(0.092)	(0.007)	(*)		(0.012)	(0.135)	(0.016)	(*)	
σ_ν	0.1691	0.1720	0.1882	0.1760	0.1232	0.1731	0.1742	0.1898	0.1754	0.1679
	(0.010)	(0.010)	(0.009)	(0.009)	(0.005)	(0.011)	(0.011)	(0.010)	(0.010)	(0.006)
$\mathcal{J}$	145.66	245.54	162.64	246.15	246.52	15.78	14.36	26.97	14.48	21.63
	[62.94]	[0.00]	[28.17]	[0.00]	[0.00]	[96.86]	[98.44]	[57.32]	[98.86]	[83.51]
$\mathcal{J}_y$	40.22	40.76	43.80	40.25	43.39	—	—	—	—	—
$\mathcal{J}_\pi$	34.22	86.25	28.34	86.43	72.05	—	—	—	—	—
$\mathcal{J}_{\pi^w}$	17.03	75.55	12.74	75.73	56.64	—	—	—	—	—
$\mathcal{J}_i$	23.50	14.48	37.63	14.56	50.65	15.78	14.36	26.97	14.48	21.63
$\mathcal{J}_\xi$	30.70	28.48	40.12	29.22	23.79	—	—	—	—	—

Notes: Standard errors in parentheses, P-value in brackets. (1) Initialization with ρ_i small and ρ_e large; (2) initialization with ρ_e small and ρ_i large; (3) constrained case $\rho_i = 0$; (4) constrained case $\rho_e = 0$; (5) constrained case $\rho_e = \rho_i$. In columns (3) and (4), * denotes a standard error not available.

no policy inertia. Column 4 corresponds to the restriction $\rho_e = 0$, i.e., to a model with nominal-interest-rate inertia and i.i.d. shocks to monetary policy. Finally, column 5 reports the estimation outcome when imposing the constraint $\rho_i = \rho_e$, thus granting the same weight to both alternative views about monetary policy. The point estimates of ψ_2 are reported, with their standard errors in parentheses. The table also reports the value of $\mathcal{J}_T$ at convergence, with the associated P-value in brackets. Finally, with our choice of weighting matrix, we can further decompose the $\mathcal{J}_T$ statistic into various components pertaining to each element of X_t.

Let us first consider the case with $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}_t^w, \hat{\imath}_t, \hat{\xi}_t)'$. In this context, we obtain a global minimum associated with the model with serially correlated monetary shocks. Additionally, the model successfully passes the overidentification test (see column 1). In contrast, the local minimum associated with a model with monetary policy inertia is blatantly rejected by the data (see column 2). The global-minimum-distance estimator yields $\rho_i = 0.30$ and $\rho_e = 0.87$. This suggests that the correct representation of monetary policy is a mix of serially correlated shocks and a modest degree of policy inertia, in the line of Rudebusch (2002, 2006). Notice that these two parameters are found to be significant. In addition, the data do not reject a model version imposing $\rho_i = 0$, while they reject the restriction $\rho_e = 0$ (see columns 3 and 4 in table 2). Notice that a quasi-likelihood ratio test would, however, reject the restriction of no monetary policy inertia (see columns 1 and 3).

To understand why the data reject the model with high monetary policy inertia, it is instructive to consider the decomposition of $\mathcal{J}_T$ according to the components of X_t. When comparing columns 1 and 2, we see that the two representations of monetary policy deliver very similar results when it comes to output, the nominal interest rate, and money growth. In other words, these three variables are weakly informative about the relevant form of monetary policy. What turns out to be really discriminating is the behavior of inflation and wage inflation. In this case, the DSGE model with policy inertia proves unable to mimic the delayed and persistent responses of these variables.

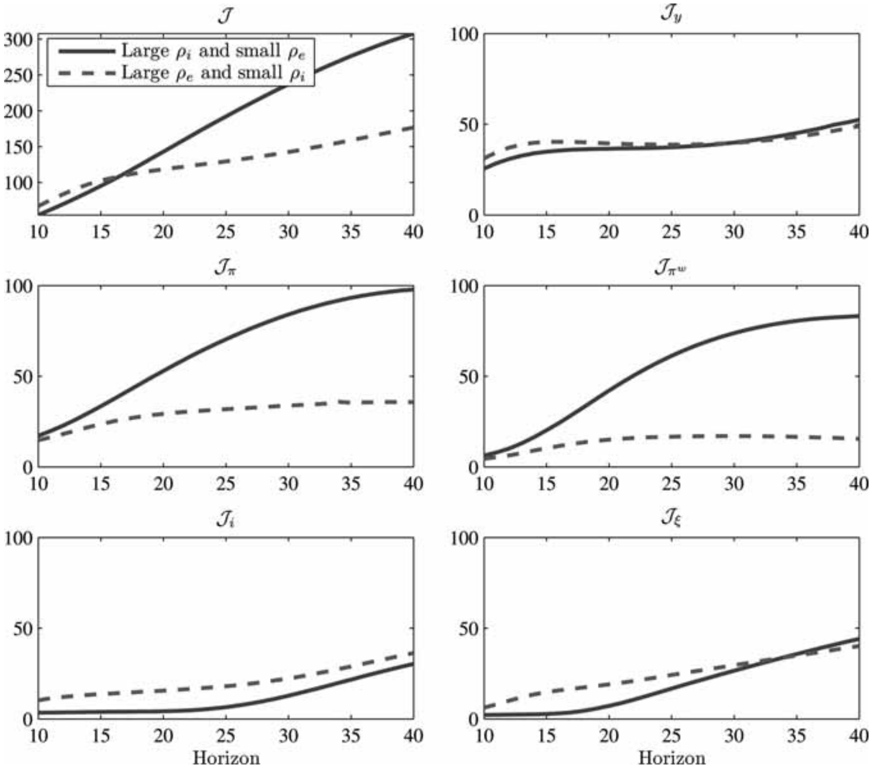
This failure is illustrated by comparing IRFs in figure 1. The lines marked with circles correspond to the DSGE point estimates with monetary policy inertia, whereas the lines marked with stars correspond to the model with persistent shocks. The dynamic responses of output, the federal funds rate, and money growth do not appear to be qualitatively affected by the specification of monetary policy. To the contrary, the model's IRFs of inflation and wage inflation sharply differ. The model with persistent shocks and moderate interest rate inertia successfully matches the essential features of the data. This is no longer the case when we consider a model with a large degree of interest rate inertia, especially so when it comes to inflation and wage inflation.

Column 5 in table 2 shows that the restriction $\rho_i = \rho_e$ is not supported by the data. Indeed, such a restriction deteriorates the model fit on virtually all dimensions, except maybe for money growth. Thus, a specification of monetary policy that grants the same weights to policy inertia and persistent shocks provides a fit that is substantially worse than the one with highly persistent shocks and moderate policy inertia.

Second, let us consider the case with $X_t = \hat{\imath}_t$. The latter is investigated as a simple way of illustrating the lack of information resulting from a quantitative assessment of our model based on a single variable. In some sense, this problem is reminiscent of the absence of clear-cut conclusions obtained in the literature focusing on a single policy rule equation; see Rudebusch (2002, 2006). Now we face the "multiple-optima" problem, since the two representations of monetary policy deliver very close objective functions at convergence. In addition, none are rejected by the data, so that they appear to be "observationally equivalent" in terms of the $\mathcal{J}_T$ statistic (see columns 1–5 in the right panel of table 2). This experiment illustrates that focusing only on the nominal interest rate does not yield a clear conclusion as to the relevant representation of monetary policy. What really matters is the aggregate dynamics (especially the dynamics of inflation and wage inflation) implied by the alternative specifications of monetary policy.

The previous results are obtained for a horizon $k = 30$. Under this assumption, we were able to discriminate between the two competing representations of monetary policy, because a model with large interest rate inertia fails to mimic the delayed U-shaped responses of inflation and wage inflation. To further illustrate the information contained in these hump-shaped patterns, we now vary the horizon k between 10 and 40. Figure 2 reports the $\mathcal{J}_T$ statistic as well as its

Figure 2. Decomposition of the $\mathcal{J}_T$ Statistic as a Function of the Time Horizon



decomposition according to the elements of X_t. In this exercise, we select $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}_t^w, \hat{\imath}_t, \hat{\xi}_t)'$ and reestimate the policy parameters for each selected horizon. In each panel, the solid lines correspond to the value of the objective function $\mathcal{J}_T$ as well as its decomposition in the case of monetary policy inertia, while the dashed lines correspond to the case with persistent shocks. Let us first focus on the global test—i.e., the $\mathcal{J}_T$ statistic—in the upper-left panel. We see that for relatively short horizons ($k = 10, \ldots, 15$), the two representations of monetary policy yield comparable results. Clearly, focusing only on short-run responses does not allow us to discriminate between the two specifications. However, as soon as k is sufficiently large to include the delayed hump patterns of inflation and wage inflation (see the two middle graphs), the performances of the two competing versions start to dramatically diverge. In particular, the monetary-policy-inertia specification faces more and more troubles reproducing the data.

5.4 Sensitivity to Calibration

We check whether the previous findings crucially depend on our particular calibration. A simple way to assess the importance of our calibration is to redo our analysis, perturbing some key model parameters. Table 3 reports the outcome of this sensitivity analysis. We identify key parameters governing the dynamic behavior of our model relating to preferences, technology, price/wage setting, and the nominal-interest-rate target level. For each alternative parameter value, we reestimate the model and recompute the $\mathcal{J}_T$ statistic at convergence with $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}^w_t, \hat{\imath}_t, \hat{\xi}_t)'$ or $X_t = (\hat{\imath}_t)$.

Preferences. The "Preferences" panel of table 3 reports the effect of shutting habit formation down (i.e., $b = 0$). When $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}^w_t, \hat{\imath}_t, \hat{\xi}_t)'$, this has the obvious effect of dramatically worsening the model's performance. Notice that, in this case, the two representations are unambiguously rejected by the data. Following Giannoni and Woodford (2004), we drastically decrease the elasticity of labor supply, setting $\omega_w = 10$. In this case, the model's performances are always improved, but the model with policy inertia is still rejected. Finally, we increase the sensitivity of money demand to the nominal interest rate, i.e., $\eta_i = 3$. The model's performances with persistent shocks are affected, but the model still passes the overidentification test. In contrast, when we focus exclusively on the nominal interest rate ($X_t = (\hat{\imath}_t)$), neither of the alternative representations can be rejected. More importantly, we cannot discriminate between these two policies based on the $\mathcal{J}_T$ statistic. This means that while the estimated models cannot generically mimic the dynamic responses of inflation and wage inflation, focusing exclusively on $\hat{\imath}_t$ would lead us to incorrectly fail to reject any model versions. This is a further illustration of the need for considering the dynamic behavior of alternative variables to properly discriminate between the competing monetary policies.

Technology. In the "Technology" panel of table 3, we investigate the sensitivity of our results to perturbations on technology parameters. Following Galí and Rabanal (2004), we assume constant

Table 3. Sensitivity to Calibration

Initialization		Based on $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}_t^w, \hat{\imath}_t, \hat{\xi}_t)'$		Based on $X_t = (\hat{\imath}_t)$	
		$\rho_e \gg \rho_i$	$\rho_i \gg \rho_e$	$\rho_e \gg \rho_i$	$\rho_e \gg \rho_i$
Parameters	**Value**	$\mathcal{J}_T$			
Preferences					
b	0.00	396 [0.00]	528 [0.00]	25 [62.69]	#
ω_w	10.00	120 [97.00]	218 [0.04]	13 [99.25]	14 [99.02]
η_i	3.00	181 [5.56]	275 [0.00]	16 [97.20]	14 [98.48]
Technology					
ϕ	1.00	103 [99.99]	199 [0.65]	11 [99.89]	13 [99.25]
θ_p	11.00	110 [99.50]	211 [0.11]	25 [62.69]	13 [99.21]
θ_w	11.00	183 [4.33]	281 [0.00]	19 [88.62]	16 [96.47]
Price/Wage Setting					
γ_p	0.00	235 [0.00]	300 [0.00]	10 [99.93]	16 [96.47]
γ_w	0.00	266 [0.00]	308 [0.00]	22 [76.44]	19 [90.21]
α_p	0.00	356 [0.00]	#	20 [87.69]	#
α_w	0.00	481 [0.00]	499 [0.00]	24 [70.36]	#
Target Level					
a_π	3.00	133 [86.45]	264 [0.00]	15 [98.29]	16 [96.77]
a_y	0.50	147 [60.38]	216 [0.05]	18 [92.15]	14 [98.88]

Notes: The label $\rho_e \gg \rho_i$ refers to an initialization of the estimation with ρ_e larger than ρ_i. Symmetrically, the label $\rho_i \gg \rho_e$ refers to an initialization of the estimation with ρ_i larger than ρ_e. P-value in brackets. A # in the "$\rho_i \gg \rho_e$" panel refers to the corresponding figure in the "$\rho_e \gg \rho_i$" panel.

returns to scale in labor input, thus imposing $\phi = 1$. The model's performances are improved for both specifications of monetary policy. However, the policy inertia is again rejected. We also modify the markups on prices without affecting our results. To the contrary, when we increase the degree of market power on the labor market, we substantially reduce the model's ability to reproduce the IRFs of X_t. Under this assumption, both versions are rejected by the data. Once again, when we focus on $X_t = (\hat{\imath}_t)$, we fail to reject either of the two competing representations of monetary policy.

Price/Wage Setting. In the "Price/Wage Setting" panel of table 3, we experiment with altering the details of the price- and wage-setting side of the model. We first shut down the indexation to past inflation in either the price or wage equations ($\gamma_p = 0$ or $\gamma_w = 0$). In both cases, this dramatically worsens the model's fit, especially so when it comes to inflation and wage inflation. Recall that these two variables were crucial in helping us sort out which specification of monetary policy was supported by the data. Not surprisingly, in the present case, both versions are rejected. Second, we assume perfect flexibility of either prices or wages ($\alpha_p = 0$ or $\alpha_w = 0$). In both cases, the model is rejected. Contrary to the previous experiment, when we focus on $X_t = (\hat{\imath}_t)$, we cannot reject either of the two competing representations of monetary policy, which prove almost completely insensitive to such parameter perturbations. This illustrates once more the need for further information.

Nominal-Interest-Rate Target Level. Finally, in the "Target Level" panel of table 3, we experiment with the parameters governing the target level of the nominal interest rate, namely a_π and a_y. We set a_π to a larger value than considered by Taylor (1993), $a_\pi = 3$. When $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}_t^w, \hat{\imath}_t, \hat{\xi}_t)'$, the discrepancy between the two alternative specifications of monetary policy widens, especially so when it comes to inflation and wage inflation. This results from the fact that increasing a_π increases the amount of information in the target level of the nominal interest rate. When it comes to a_y, the quantitative findings are left unaffected. Conversely, when we focus on $X_t = (\hat{\imath}_t)$, we fail to reject either of the two competing representations of monetary policy. This is more troubling than one would have expected. Indeed, increasing the volatility of the target can potentially eliminate the identification problem. This is not the case in practice. When we focus only on $\hat{\imath}_t$, the discriminating power of inflation and wage inflation is shut down, which keeps us from reaching a clear-cut conclusion.

6. Practical Differences between Policy Inertia and Persistent Shocks

This section presents two illustrations of the practical differences between policy inertia and persistent shocks. We first investigate their quantitative implications for policy rule estimation. Second, we perform forecasting exercises using the Volcker disinflation as a case study that can potentially reveal striking differences between the two views.

6.1 Implications for Monetary Policy

The previous exercise has allowed us to discriminate between two alternative representations of monetary policy. However, since all the parameters were calibrated, including the responsiveness of monetary policy to inflation and output, this exercise is necessarily silent on the consequences of a monetary policy misspecification. The question we ask now is the following: Would we get different estimates of the responsiveness of monetary policy to inflation and output in the case of monetary policy inertia and in the case of persistent shocks?

So as to answer this question, we reestimate our model under the two alternative representations and allow a_π and a_y to be freely estimated. Table 4 reports the estimation results. The first column reports results obtained with persistent shocks, while the second column corresponds to policy inertia. As before, the global minimum is obtained with highly serially correlated monetary shocks and a small degree of interest rate smoothing. Once again, the responses of inflation and wage inflation allow us to discriminate between the two competing views.

Our results also suggest that the two alternative views yield very contrasted findings relative to the reaction of monetary authorities to inflation and output. In the case of policy inertia, the latter is almost passive regarding inflation but highly reactive to output fluctuations. Under persistent shocks, we obtain a reverse configuration, suggesting a very aggressive monetary policy in response to inflation

Table 4. Estimation Results for the Complete Taylor Rule

Parameter	$\rho_e \gg \rho_i$	$\rho_i \gg \rho_e$
ρ_i	0.2773 (0.187)	0.8832 (0.233)
ρ_e	0.9501 (0.013)	0.4169 (0.223)
a_π	3.0815 (0.568)	1.0509 (0.809)
a_y	0.0000 (−)	0.7065 (2.152)
σ_ν	0.1670 (0.010)	0.1594 (0.010)
$\mathcal{J}$	129.43 [88.64]	173.64 [9.06]
$\mathcal{J}_y$	46.28	42.32
$\mathcal{J}_\pi$	26.56	51.13
$\mathcal{J}_{\pi^w}$	10.90	34.73
$\mathcal{J}_i$	20.87	24.43
$\mathcal{J}_\xi$	24.82	21.03

Notes: The label $\rho_e \gg \rho_i$ refers to an initialization of the estimation with ρ_e larger than ρ_i. Symmetrically, the label $\rho_i \gg \rho_e$ refers to an initialization of the estimation with ρ_i larger than ρ_e. Standard errors in parentheses, P-value in brackets.

and a zero concern for output fluctuations. The global minimum thus corresponds to a policy rule enforcing the Taylor principle.

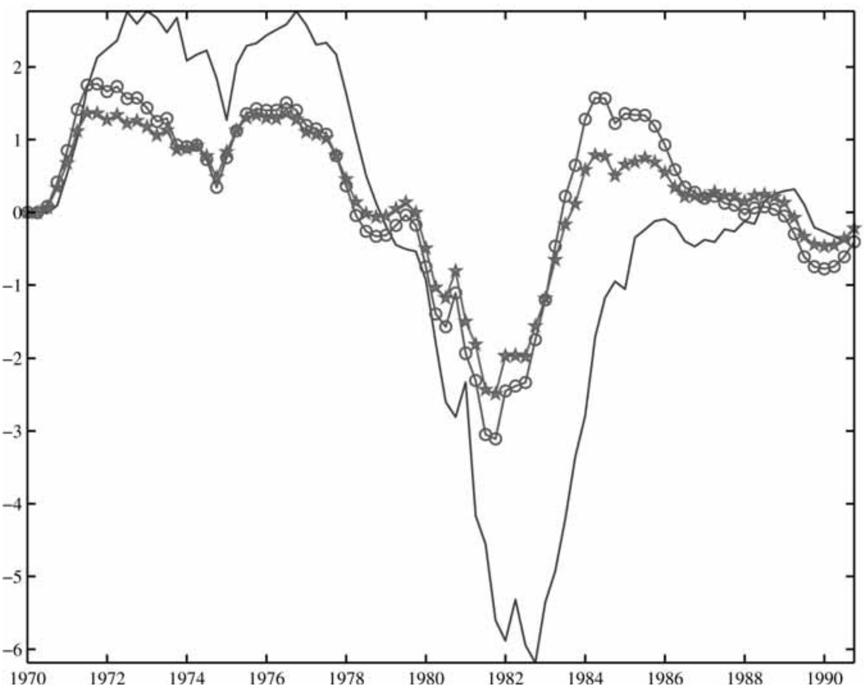
6.2 Inspecting the Monetary Policy Rules through the Lenses of the Volcker Disinflation

In this section, we compare the performances of the two alternative representations of monetary policy using the Volcker disinflation as an episode that can potentially reveal striking differences between these policy rules. This episode corresponds to what can be a priori viewed as a period of large contractionary monetary policy shocks.

Thus, comparing the model under the two rules with what actually happened during this episode constitutes a legitimate experiment.

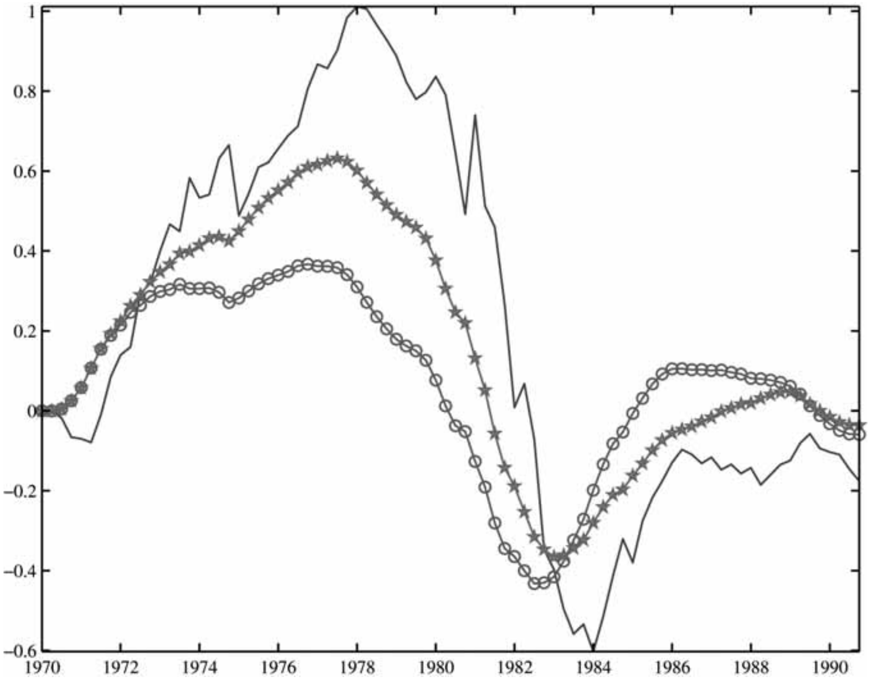
Since our limited-information approach exclusively relies on monetary policy shocks, we start by reconstructing historical data from the SVAR after having shut down all other shocks than monetary shocks. We then feed the identified monetary shocks in our DSGE model using either of the rules and compute artificial data. To compare the performances of the two rules on output and inflation, we focus on the sample period preceeding and succeeding the Volcker disinflation, 1970:Q1–1990:Q4. To make things comparable, we use the same initial conditions in 1970 either for the SVAR or for the model. The outcome of these comparisons is reported in figures 3

Figure 3. Historical Simulation of Output



Note: Plain line: SVAR. Line marked with circles: DSGE model with policy inertia. Line marked with stars: DSGE model with persistent shocks. All the data are demeaned prior to simulation.

Figure 4. Historical Simulation of Inflation



Note: Plain line: SVAR. Line marked with circles: DSGE model with policy inertia. Line marked with stars: DSGE model with persistent shocks. All the data are demeaned prior to simulation.

and 4 for output and inflation, respectively. In each case, the plain line corresponds to the SVAR-based historical data, while the lines marked with circles and stars correspond to the DSGE model with policy inertia and serially correlated shocks, respectively.[15]

Figure 3 confirms our previous findings: the two alternative policy rules have similar implications when it comes to output dynamics. As is clear from the picture, in both cases, the simulated samples are very similar, either in terms of persistence or volatility. To the contrary, as shown in figure 4, the two alternative rules have very

[15]Notice that this exercise contains the same information as our IRF-based assessment of the model's performance. What is interesting here is that it allows us to focus on a specific episode within our sample.

different implications in terms of inflation dynamics. Under the rule with policy inertia, inflation dynamics exhibit a smaller variance when compared to the SVAR-based counterpart. In addition, inflation drops too slowly during the Volcker disinflation and rises much too fast after 1983. Overall, this gives an inaccurate description of actual inflation dynamics during this particular episode. In contrast, when the rule features serially correlated shocks and a modest degree of policy inertia, the model is better suited to capture the large inflation peak of 1978 as well as the sharp decline following the disinflation. Additionally, the model does not predict a rapid rise in inflation after 1983, consistent with what the SVAR-based path suggests.

This exercise provides a confirmation that inflation dynamics contain more useful pieces of information than the dynamics of output for the purpose of disentangling the two alternative representations of monetary policy. The Volcker disinflation, taken as a case study, favors the persistent-shocks view as a practical approximation of actual monetary policy, as was to be expected from our previous quantitative investigation.

7. Conclusion

In this paper, we proposed a simple econometric framework to discriminate between two alternative representations of monetary policy. This approach draws heavily from the restrictions contained in the monetary DSGE model used in our empirical analysis. More precisely, thanks to these restrictions, different monetary policies can have radically different implications in terms of aggregate dynamics. Building on this well-known property of DSGE models, we are able to identify which policy rule best fits the data.

Our results are twofold. First, when the framework contains enough information, a policy rule with modest interest rate inertia and highly serially correlated shocks—which contrasts with most current implementations of monetary policy rules—satisfactorily matches the data. In particular, we found that the dynamics of inflation and wage inflation are particularly helpful for inferring the correct specification of monetary policy. However, output, the nominal interest rate, and the money growth rate do not contain very discriminating information. In addition, the inverted hump patterns displayed by the impulse responses of inflation and wage inflation are found to be particularly relevant for this purpose. Second, when the framework is not informative enough—i.e., when we focus on the sole dynamics of the federal funds rate—we are unable to discriminate between the two alternative monetary policy rules. These results highlight the low discriminating power of single-equation approaches. Overall, our results suggest that using extra macroeconomic information can help reach clear-cut conclusions as to the correct empirical representation of monetary policy rules. These two main findings are confirmed when we investigate practical differences between these two alternative monetary policy representations.

Appendix. Model Details

Production Side

A large number of competitive firms produce a homogeneous good that can be either consumed (y_t) or used as a material input in production (q_t). The overall aggregate demand is $d_t \equiv y_t + q_t$, and P_t is the associated nominal price. Following Kimball (1995) and Woodford (2003), the production function is of the form

$$\int_0^1 G\left(\frac{d_t(\varsigma)}{d_t}\right) \mathrm{d}\varsigma = 1, \tag{10}$$

where $d_t(\varsigma)$ denotes the input of intermediate goods $\varsigma \in [0, 1]$, and the function G is increasing, strictly concave, and satisfies the normalization $G(1) = 1$. The representative final goods producer chooses $\{d_t(\varsigma), \varsigma \in [0, 1]\}$ and d_t in order to maximize profits

$$\max_{\{d_t(\varsigma)\}} P_t d_t - \int_0^1 P_t(\varsigma) d_t(\varsigma) \mathrm{d}\varsigma,$$

subject to (10), where $P_t(\varsigma)$ is the nominal price of intermediate good ς. Monopolistic firms produce the intermediate goods $\varsigma \in [0, 1]$. Each firm ς is the sole producer of intermediate good ς. Following Woodford (2003), we assume that monopolist ς produces good ς

with the inputs of aggregate labor $n_t(\varsigma)$ and material goods $q_t(\varsigma)$ according to the following production possibilities:

$$\min\left\{\frac{F(n_t(\varsigma))}{1-s_q}, \frac{q_t(\varsigma)}{s_q}\right\} \geq d_t(\varsigma),$$

where $F(\cdot)$ is an increasing and concave production function and s_q is the share of material goods in gross output. Let $\theta_p(z)$ denote the elasticity of demand for a producer of intermediate goods facing the relative demand $z = d_t(\varsigma)/d_t$. According to our specification, $\theta_p(z) \equiv -G'(z)/(zG''(z))$. This illustrates that intermediate-goods firms face a varying elasticity of demand for their output, implying a varying markup, which is denoted by $\mu_p(z) \equiv \theta_p(z)/(\theta_p(z)-1)$.

Following Calvo (1983), we assume that in each period of time and prior to observing the monetary policy shock, a monopolistic firm can reoptimize its price with probability $1-\alpha_p$, irrespective of the elapsed time since it last revised its price. As in Woodford (2003), if the firm cannot reoptimize its price, the latter is rescaled according to the simple revision rule $P_T(\varsigma) = (1+\delta^p_{t,T})P_t(\varsigma)$, where

$$1+\delta^p_{t,T} = \begin{cases} \prod_{j=t}^{T-1}(1+\pi)^{1-\gamma_p}(1+\pi_j)^{\gamma_p} & \text{if } T > t \\ 1 & \text{otherwise} \end{cases},$$

where $\pi_t = P_t/P_{t-1} - 1$ represents the inflation rate, π is the steady-state inflation rate, and $\gamma_p \in [0,1]$ measures the degree of indexation to the most recently available inflation measure. Let $P^\star_t(\varsigma)$ denote the price chosen in period t by monopolist ς if drawn to reoptimize. Then, firm ς chooses $P^\star_t(\varsigma)$ in order to maximize

$$\mathrm{E}_{t-1}\sum_{T=t}^{\infty}(\beta\alpha_p)^{T-t}\lambda_T\left\{\frac{\left(1+\delta^p_{t,T}\right)P^\star_t(\varsigma)}{P_T}d^\star_{t,T}(\varsigma) - S\left(d^\star_{t,T}(\varsigma)\right)\right\},$$

where λ_T is the representative household's marginal utility of wealth in period T; $\mathrm{E}_{t-1}\{\cdot\}$ is the expectation operator conditional on information available when the firm sets its price; $S(d_t(\varsigma))$ is the real cost of producing $d_t(\varsigma)$ units of good ς; and $d^\star_{t,T}(\varsigma)$, the demand for good ς at T if firm ς last reoptimized its price at t, obeys

$$G'\left(\frac{d^\star_{t,T}(\varsigma)}{d_T}\right) = \left(\frac{\left(1+\delta^p_{t,T}\right)P^\star_t(\varsigma)}{P_T}\int_0^1\frac{d_t(u)}{d_t}G'\left(\frac{d_t(u)}{d_t}\right)\mathrm{d}u\right).$$

Standard manipulations yield the log-linear New Keynesian Phillips curve

$$\hat{\pi}_t - \gamma_p \hat{\pi}_{t-1} = \mathrm{E}_{t-1}\{\kappa_p(\hat{w}_t + \omega_p \hat{y}_t) + \beta(\hat{\pi}_{t+1} - \gamma_p \hat{\pi}_t)\}, \qquad (11)$$

with

$$\kappa_p \equiv \varkappa \frac{(1-\alpha_p)(1-\beta\alpha_p)}{\alpha_p}, \quad \varkappa \equiv \frac{1}{(1-\mu_p s_q)^{-1}(1+\theta_p \epsilon_\mu) + \theta_p \omega_p}.$$

In equation (11), $\hat{\pi}_t$ is the log-deviation of $1+\pi_t$; $\hat{y}_t$ and $\hat{w}_t$ are the log-deviations of y_t and w_t (real wage), respectively; $\theta_p \equiv \theta_p(1)$ is the steady-state elasticity of demand for a producer of intermediate goods; $\mu_p \equiv \mu_p(1)$ is the steady-state markup factor; and

$$\omega_p \equiv -\frac{F''(n)n}{F'(n)} \frac{F(n)}{F'(n)n}.$$

Here, $F(n)$, $F'(n)$, and $F''(n)$ denote the value of F and its first and second derivatives, evaluated at the steady-state value of n. Following Woodford (2003), we let ϵ_μ denote the elasticity of $\mu_p(z)$ in the neighborhood of $z=1$, i.e., $\epsilon_\mu = \mu_p'(1)/\mu_p(1)$.

Aggregate Labor Index and Households

Following Erceg, Henderson, and Levin (2000), we assume for convenience that a set of differentiated labor inputs, indexed by $\upsilon \in [0,1]$, are aggregated into a single labor index h_t by competitive firms, which will be referred to as labor intermediaries. They produce the aggregate labor input according to the following constant elasticity of substitution technology:

$$h_t = \left(\int_0^1 h_t(\upsilon)^{(\theta_w - 1)/\theta_w} \mathrm{d}\upsilon \right)^{\theta_w/(\theta_w - 1)},$$

where $\theta_w > 1$ is the elasticity of substitution between any two labor types. The associated aggregate nominal wage obeys

$$W_t = \left(\int_0^1 W_t(\upsilon)^{1-\theta_w} \mathrm{d}\upsilon \right)^{1/(1-\theta_w)},$$

where $W_t(\upsilon)$ denotes the nominal wage rate paid to type υ labor. The economy is inhabited by a continuum of differentiated households, indexed by $\upsilon \in [0,1]$. A typical household, say household υ, must select a sequence of consumptions and nominal money and bond holdings, as well as a nominal wage. The timing of events is as follows. Prior to observing the monetary policy shock, the household decides how much to consume and sets its nominal wage. The shock is then realized, and bond and money holdings decisions are taken. Household υ's goal in life is to maximize

$$\mathrm{E}_{\Phi_t} \sum_{T=t}^{\infty} \beta^{T-t} [U(c_T - bc_{T-1}, m_T) - V(h_T(\upsilon))],$$

where $\beta \in (0,1)$ is the subjective discount factor; $b \in (0,1)$ is the habit parameter; c_t is consumption; $m_t \equiv M_t/P_t$ denotes real cash balances at the end of the period, where M_t denotes nominal cash balances; and $h_t(\upsilon)$ denotes household υ's labor supply at period t. Here, E_{Φ_t} is a conditional expectation operator reflecting the particular information sets at the household's disposal when taking its decisions. Household υ maximizes its intertemporal utility subject to the sequence of constraints

$$P_t \mathrm{tax}_t + P_t c_t + M_t + \frac{B_t}{1+i_t} \leq W_t(\upsilon) h_t(\upsilon) + B_{t-1} + M_{t-1} + P_t \mathrm{div}_t,$$

where div_t denotes real profits redistributed by monopolistic firms; B_t denotes the nominal bonds acquired in period t and maturing in period $t+1$; i_t denotes the gross nominal interest rate; and tax_t is a lump-sum tax levied by the government. As in Woodford (2003), we assume that there is a satiation level $m^\star$ for real balances such that $U_m = 0$ for $m \geq m^\star$. Thus, when m_t reaches $m^\star$ from below, the transaction services of real cash balances yield lower and lower marginal utility. Let λ_t denote the Lagrange multiplier associated with the household's budget constraint. According to the timing of decisions embedded in Φ_t, the log-linearization of the first-order conditions associated with c_t, B_t, and M_t yields

$$\begin{aligned} \mathrm{E}_{t-1}\{&\beta b(\hat{c}_{t+1} - b\hat{c}_t) - (\hat{c}_t - b\hat{c}_{t-1}) \\ &+ \sigma\chi(\hat{m}_t - \beta b\hat{m}_{t+1}) - \varphi^{-1}\hat{\lambda}_t\} = 0, \end{aligned} \tag{12}$$

$$\hat{\lambda}_t = \hat{\imath}_t + \mathrm{E}_t\{\hat{\lambda}_{t+1} - \hat{\pi}_{t+1}\}, \tag{13}$$

$$\hat{m}_t = \eta_y(\hat{c}_t - b\hat{c}_{t-1}) - \eta_i\hat{\imath}_t, \tag{14}$$

where $\hat{c}_t$, $\hat{m}_t$, $\hat{\imath}_t$, and $\hat{\lambda}_t$ are the log-deviations of c_t, m_t, $1+i_t$, and λ_t, respectively, and where we defined the auxiliary parameters $\sigma^{-1} = -U_{cc}c/U_c$, $\chi = U_{cm}m/U_c$, $\varphi^{-1} = (1-\beta b)\sigma$, $\eta_y = -U_{mc}c/(U_{mm}m)$, and $\eta_i = -(1-\beta b)U_c/(U_{mm}m)$. Notice that $\chi = (1-\beta b)\eta_y/(\bar{v}\eta_i)$, where $\bar{v}$ is the steady-state value of c_t/m_t.

A typical household v acts as a monopolistic supplier of type v labor. It is assumed that at each point in time, and prior to observing the monetary policy shock, only a fraction $1-\alpha_w$ of the households can set a new wage, which will remain fixed until the next time period the household is drawn to reset its wage. The remaining households simply revise their wages according to the simple rule $W_T(v) = (1+\delta^w_{t,T})W_t(v)$, where

$$1+\delta^w_{t,T} = \begin{cases} \prod_{j=t}^{T-1}(1+\pi)^{1-\gamma_w}(1+\pi_j)^{\gamma_w} & \text{if } T > t \\ 1 & \text{otherwise} \end{cases},$$

where $\gamma_w \in [0,1]$ measures the degree of indexation to the most recently available inflation measure.

Let us now consider the wage-setting decision confronting a household drawn to reoptimize its nominal wage rate in period t, say household v. Let us define wage inflation $\pi^w_t \equiv W_t/W_{t-1} - 1$. Now, let $W^\star_t(v)$ denote the wage rate chosen in date t and $h^\star_{t,T}(v)$ denote hours worked in period T if household v last reoptimized its wage in period t, which obey the relationship

$$h^\star_{t,T}(v) = \left(\frac{(1+\delta^w_{t,T})W^\star_t(v)}{W_T}\right)^{-\theta_w} h_T.$$

$W^*_t(v)$ is then selected so as to maximize

$$\mathrm{E}_{t-1}\sum_{T=t}^{\infty}(\beta\alpha_w)^{T-t}\left\{\lambda_T\frac{(1+\delta^w_{t,T})W^\star_t(v)}{P_T}h^\star_{t,T}(v) - V\big(h^\star_{t,T}(v)\big)\right\}.$$

Log-linearizing the associated first-order condition yields

$$\hat{\pi}^w_t - \gamma_w\hat{\pi}_{t-1} = \mathrm{E}_{t-1}\big\{\kappa_w(\omega_w\hat{h}_t - \hat{\lambda}_t - \hat{w}_t) + \beta\big(\hat{\pi}^w_{t+1} - \gamma_w\hat{\pi}_t\big)\big\}, \tag{15}$$

where $\hat{\pi}_t^w$ is the log-deviation of $1 + \pi_t^w$ and where we defined the composite parameters

$$\kappa_w = \frac{(1-\alpha_w)(1-\beta\alpha_w)}{\alpha_w(1+\omega_w\theta_w)}, \quad \omega_w = \frac{V_{hh}h}{V_h}.$$

Finally, $\hat{\pi}_t$ and $\hat{\pi}_t^w$ are linked together through the relation

$$\hat{\pi}_t^w = \hat{w}_t - \hat{w}_{t-1} + \hat{\pi}_t. \tag{16}$$

The model is closed by specifying the policy rule (1)–(3).

In equilibrium, it must be the case that $y_t = c_t$ and $h_t = n_t$. Furthermore, from the aggregate production function, it must also be the case that $\hat{n}_t = \phi\hat{y}_t$, where $\phi^{-1} = F'(n)n/F(n)$. Substituting these relations in the system composed of (11)–(16), augmented with equations (1)–(3), we obtain a rational-expectations system of linear equations, which we solve using standard methods.

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