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Monetary Policy Inertia or Persistent Shocks?[1]

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Résumé :

Ce papier propose un cadre économétrique simple permettant d'évaluer les rôles respectifs de la politique monétaire inertielle et des chocs monétaires autocorrélés. La méthode exploite les restrictions inter-équations issues d'un modèle DSGE qu'on estime de façon à reproduire les réponses des variables d'intérêt à des chocs monétaires tirées d'un modèle VAR structurel. Nous montrons que lorsque les variables incluses dans l'estimation fournissent suffisamment d'information, la méthode permet de discriminer entre les deux représentations de la politique monétaire. En revanche, lorsque seul le taux d'intérêt est retenu dans l'estimation, il n'est pas possible de départager clairement les deux configurations.

Mots-clés : Règle de Taylor, identification de la politique monétaire, inertie de la politique

Abstract:

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. The procedure exploits the cross-equation restrictions provided by a DSGE model which is confronted to a monetary SVAR. We show that, provided enough informative variables are included in the formal test, the data favor a monetary policy representation with low 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 as to the correct representation of monetary policy.

Keywords: Taylor rule, Monetary policy identification, Policy Inertia.

JEL Codes: C52, E31, E32, E52.

Résumé non technique :

Au cours des dernières années, de nombreuses études empiriques ont consacré d'importants efforts afin d'illustrer le pouvoir descriptif de règles monétaires de type Taylor. Un résultat récurrent dans cette littérature est que le taux d'intérêt retardé dans ce type de règles est très significatif et prend des valeurs élevées dans l'intervalle [0,7 0,9], sur données trimestrielles. D'aucuns considèrent ces résultats comme révélateurs d'un comportement de lissage du taux d'intérêt mené par la banque centrale.

L'objectif de ce papier est d'étudier l'importance relative de l'hypothèse d'ajustement partiel contre une interprétation alternative du comportement de la banque centrale qui pourrait créer l'illusion d'une inertie de la politique monétaire. Plus précisément, nous étudions dans quelle mesure la banque centrale peut répondre à des facteurs persistants, omis dans la règle de Taylor, créant ainsi l'illusion d'une certaine volonté de lissage. En particulier, une explication alternative possible de la persistance observée des taux d'intérêt est la présence de chocs autocorrélés dans la règle de taux. Ces chocs peuvent représenter des événements contingents auxquels la banque centrale doit faire face lorsqu'elle décide sa politique.

Bien que ces deux visions concurrentes de la politique monétaire induisent des conclusions contrastées quant au comportement effectif de la banque centrale, les données agrégées sont plutôt silencieuses sur cette question. Cette absence de conclusions claires peut s'expliquer par le problème d'identification bien connu dans les modèles combinant ajustement partiel et chocs autocorrélés. De telles difficultés ont été abondamment abordées dans la littérature économétrique et dans différentes applications empiriques. Un fait robuste dans cette littérature est que le problème d'identification apparaît plus fréquemment lorsque les régresseurs sont faiblement informatifs. Dans le cadre de la règle de Taylor, le problème se présente lorsque la cible ne présente pas assez de volatilité. Ce problème remet en cause l'utilisation d'une équation unique (la règle de Taylor uniquement) comme une façon adéquate de discriminer entre ces deux visions concurrentes de la politique monétaire.

Dans ce papier, nous exploitons les restrictions inter-équations issues d'un modèle DSGE avec agents optimisateurs afin d'évaluer la pertinence empirique de ces deux représentations de la politique. Si la critique de Lucas s'applique dans ce contexte, différentes règles monétaires peuvent avoir

des implications dynamiques significativement différentes non seulement sur le taux d'intérêt, mais aussi sur d'autres variables pertinentes pour l'analyse économique.

Nous procédons de la manière suivante. Nous estimons dans un premier temps, sur données américaines, un modèle VAR structurel avec des restrictions de court terme afin d'identifier les chocs sur la politique monétaire. Dans un second temps, les paramètres de la règle dans le modèle DSGE sont estimés de façon à reproduire les réponses de l'économie à une innovation monétaire.

Lorsque nous considérons les réponses du produit, de l'inflation, de l'inflation salariale, du taux d'intérêt nominal et du taux de croissance de la monnaie, nous pouvons sans ambiguïté discriminer entre les deux représentations de la politique monétaire. Nos résultats mettent en évidence qu'une configuration avec un ajustement rapide et un fort degré d'autocorrélaion des chocs fournit le meilleur ajustement aux données.

En revanche, lorsque nous considérons uniquement la réponse du taux d'intérêt nominal, il n'est pas possible de discriminer entre ces deux représentations. C'est pourquoi nous insistons sur la nécessité d'inclure des éléments informatifs des données. Dans notre cadre, cette information est essentiellement contenue dans les réponses persistantes et en cloche de l'inflation et de l'inflation salariale.

Non-technical summary:

Over the recent years, empirical studies devoted important efforts to illustrating the descriptive power of Taylor-like monetary rules. A recurrent finding of this literature is that the lagged interest rate in the Taylor-like rules is highly significant, taking on large values in the [0.70,0.90] range on quarterly data. It has been argued that this may represent a process of interest–rate smoothing or policy inertia by the central bank, which would operate a partial adjustment process in the announced or realized interest rate.

The aim of this paper is to study the relative importance of the partial adjustment hypothesis versus different interpretations of central bank behavior that can make the illusion of policy inertia. More precisely, we ask whether it could be the case that the central bank were mainly responding to persistent factors not included in the Taylor rule equation and that would make the illusion of a certain desire for smoothing. In particular, a possible alternative explanation of actual persistence in monetary policy is the presence of serially correlated shocks in the realizations of the interest rate rule. These shocks may represent any contingent event the central bank faces when deciding the interest rate.

While these two competing views of monetary policy entail very different conclusions about the effective behavior of central banks, aggregate data have typically been fairly silent as to which is the correct representation of actual monetary policy. It might be the case that the lack of clear-cut conclusions as to the relative importance of these two competing views of monetary policy is a figment of well known identification and multiple optima issues typically arising in models of partial adjustment with serially correlated shocks. Such difficulties have been documented in the econometric literature as well as in various empirical applications of the partial adjustment model. A robust finding of this literature is that the identification problem may arise more frequently with regressors that are of minor empirical importance. Within the framework of the Taylor rule with partial adjustment and serially correlated shocks, the problem would appear when the target does not display enough variability. This problem 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 discussed above.

In this paper, we argue that we can exploit the cross-equation restrictions arising from a Dynamic Stochastic General Equilibrium (DSGE) model with optimizing agents. More precisely, we assess the relevance of these two competing views of monetary policy in terms of aggregate effects on the dynamics of other variables than solely the nominal interest rate. If the *Lucas critique* holds in this context, we should expect that these alternative rules will have significantly different impacts on the equilibrium dynamics.

Using US data, we follow the methodology of the Minimum Distance Estimator (MDE) popularized by Rotemberg and Woodford (1997) and Christiano, Eichembaum and Evans (2005). We use the predictions of a monetary DSGE model with real and nominal frictions on different sets of endogenous variables in order to estimate the degree of inertia and the degree of serial correlation of monetary shocks in the interest rate rule. We proceed as follows. We first estimate a structural vector autoregression (SVAR) with short-run restrictions so as to identify monetary policy shocks. Second, the monetary policy rule parameters in the DSGE model are pinned down so as to reproduce as well as possible the impulse response functions of key aggregate variables, as implied by the SVAR. Importantly, so as to fully control the aggregate implications of alternative policy rules, all the remaining model parameters are calibrated prior to estimation. Accordingly, the model predicted impulse response functions (IRFs) can only be affected by the unknown parameters associated with the rival policymaking hypotheses. We thus use the DSGE model as a instrument to help us discriminate between the competing representations of monetary policy.

When we consider the IRFs for output, inflation, wage inflation, the Fed funds rate, and money growth, the general result is that we still find the multiple optima phenomenon. However, using the model as a instrument, we are able to unambiguously discriminate between the two different schemes (partial adjustment versus serially correlated shocks). Our results point out that a configuration with fast partial adjustment and a high degree of serial correlation of monetary shocks minimizes the distance between the actual and predicted IRFs. This configuration also delivers a significantly better fit than the reverse configuration of high inertia and mildly serially correlated shocks.

In contrast, when we consider only the actual and predicted responses for the Fed funds rate and estimate the parameters values which accomplish the MDE criterion, we find that there is not enough evidence to discriminate between the two competing views about monetary policy. Therefore, we insist that in order to disentangle these two views, one should take into account informative features of the data. In our case, this role is devoted to the hump-shaped responses of inflation and wage inflation.

Finally, we study the robustness of our findings to the timing of decisions in the DSGE model as well as different calibrations. While the timing proves relatively innocuous, we find that changing the calibration can have dramatic effects when the estimation criterion is informative. In contrast, when the latter conveys no information (i.e. when we exclusively focus on the Fed funds rate), we are no longer able to discriminate between the two competing views. This is a further confirmation of the need to exploit the information contained in the cross-equation restrictions created by our DSGE model.

1 Introduction

Over the recent years, there has been a renewed interest in modelling monetary policymaking in terms of simple rules. A voluminous empirical literature has devoted substantial efforts to illustrating the descriptive power of this approach. In this literature, the Taylor rule has become the workhorse description of central bank behavior. Although Taylor (1993) recognized that a single equation is far from encompassing all specificities of how monetary policy is conducted, he shows that a relation stating the responses of the federal funds rate reacting to inflation and the output gap describes quite well the interest rate sequence.

Importantly, Taylor (1993) pointed out that one should not expect that policymakers "follow policy rules mechanically."[1] In addition to some technical reasons, as information availability or real–time data for forecasting purposes, there could also be some special factors for which monetary policy will need to be adjusted, such as credit crunches or financial crises, that cannot be summarized in a single equation. Therefore, one should consider the Taylor rule as a "hypothetical but representative policy rule" that could serve as a simple guide to understand the central bank performance and/or a contingency plan for policymakers.

These words of caution had echo in applied monetary economics where some researchers have specified extended Taylor rules in a parsimonious way so as to better describe the central bank policy. For example, Clarida et al. (2000) showed that the lagged interest rate is highly significant in the estimated policy rule, and concluded that nominal interest rates exhibit a sizable degree of inertia. The coefficient of the lagged interest rate in the Taylor rule has typically been found in the [0.70,0.90] range on quarterly frequency. Such a representation has now become a landmark.[2] It has been argued that this may represent a process of interest–rate smoothing or policy inertia by the central bank, which would operate a partial adjustment process in the announced or realized interest rate.

The aim of this paper is to study the relative importance of the partial adjustment hypothesis versus

[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) or Sack and Wieland (2000), among others.

different interpretations of central bank behavior that can make the illusion of policy inertia. More precisely, we ask whether it could be the case that the central bank were mainly responding to persistent factors not included in the Taylor rule equation and that would make the illusion of a certain desire for smoothing.

The interest rate smoothing hypothesis finds theoretical support in the optimal monetary policy inertia literature. The usual rationales for partial adjustment in the interest rate are the following. First, monetary policy gradualism reduces the short-run volatility of the interest rate and asset prices (see Goodfriend 1987). The second rationale corresponds to lever on expectations, so that forward–looking agents trust the monetary policymaker is committed to a gradual policy rule and thus engaged in controlling macroeconomic fluctuations (see Rotemberg and Woodford, 1999 and Woodford 2003). Last but not least, uncertainty may also motivate optimal monetary policy inertia, since caution about the actual effects of policy suggests a gradual adjustment of the nominal interest rate.

Rudebusch (2002, 2005) claims that, overall, the policy rate partial adjustment implies that future changes in the interest rate should have an important degree of forecastability. Thus, actual changes in the interest rate term structure given by the yield curve should be largely predictable. However, using data from financial markets, with the change of future interest rates measured as the rate of eurodollar deposits, Rudebusch (2002) shows that the predictive power of the expected changes is unambiguously low. Therefore, using this additional information, Rudebusch concludes that the high degree of partial adjustment may be a misinterpretation of the true nature of monetary policy.

An alternative explanation of actual persistence in monetary policy is the presence of serially correlated shocks in the realizations of the interest rate rule. These shocks may represent any contingent event the central bank faces when deciding the interest rate. Examples are credit crunches or financial crises, as noted by Taylor (1993), or the effect of real-time data when estimating the inflation and output gaps. Thus these serially correlated shocks represent a set of special factors that cannot be systematically modeled by a simple interest rate rule. As a consequence, these shocks might be interpreted as a measure of our ignorance about the monetary policymaking. The implications of this alternative representation are completely opposite to those of the partial adjustment view. Indeed, under serially correlated shocks, the central bank does not especially smooth the interest rate but rather reacts to the arrival of new information about the state of the economy.

While these two competing views of monetary policy entail very different conclusions about the effective behavior of central banks, aggregate data have typically been fairly silent as to which is the correct representation of actual monetary policy. For example Rudebusch (2002) is not able to distinguish the partial adjustment parameter from the serially correlated parameter in the interest rate equation. English et al. (2002) study the different implications of partial adjustment and serially correlated errors over the first difference of the interest rate realizations and find that there is supportive evidence that both partial adjustment and serially correlated errors are significant components of the Federal Reserve behavior since the late 1980's. Castelnuovo (2003) extends their approach, including additional variables to the usual specification of the Taylor rule. His results suggest that partial adjustment and serially correlated errors are equally important to describe the central bank behavior. Gerlach–Kristen (2004) and Apel and Jansson (2005) find similar results using Kalman filtering to account for omitted unobserved factors in the interest rate rule.

However, it might be the case that the lack of clear-cut conclusions as to the relative importance of these two competing views of monetary policy is a figment of well known identification and multiple optima issues typically arising in models of partial adjustment with serially correlated shocks. Such difficulties have been documented in the econometric literature (Griliches, 1967, Blinder, 1986, Harvey, 1990, McManus et al. 1994) as well as in various empirical applications of the partial adjustment model (Maccini and Rossana, 1984, Blinder, 1986, Goldfeld and Sichel, 1990). Importantly, for our purpose, rational expectation econometrics have also been subject to this problem, as exemplified by Sargent (1978), Eichenbaum (1983), and Kennan (1988). A robust finding of this literature is that the identification problem may arise more frequently with regressors that are of minor empirical importance. Within the framework of the Taylor rule with partial adjustment and serially correlated shocks, the problem would appear when the target does not display enough variability. This problem 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 discussed above.

In this paper, we argue that we can exploit the cross-equation restrictions arising from a Dynamic Stochastic General Equilibrium (DSGE) model with optimizing agents. Indeed, as noted by Rotemberg and Woodford (1997) "Ultimately, it is the only way in which the 'observational equivalence' of a multitude of alternative possible structural interpretations of the co-movements of aggregate series can be resolved" (page 298). More precisely, we assess the relevance of these two competing views of monetary policy in terms of aggregate effects on the dynamics of other variables than solely the nominal interest rate. If the *Lucas critique* holds in this context, we should expect that these alternative rules will have significantly different impacts on the equilibrium dynamics.

Using US data, we follow the methodology of the Minimum Distance Estimator (MDE) popularized by Rotemberg and Woodford (1997) and Christiano, Eichembaum and Evans (2005). We use the predictions of a monetary DSGE model with real and nominal frictions on different sets of endogenous variables in order to estimate the degree of inertia and the degree of serial correlation of monetary shocks in the interest rate rule. We proceed as follows. We first estimate a structural vector autoregression (SVAR) with short-run restrictions so as to identify monetary policy shocks. Second, the monetary policy rule parameters in the DSGE model are pinned down so as to reproduce as well as possible the impulse response functions of key aggregate variables, as implied by the SVAR. Importantly, so as to fully control the aggregate implications of alternative policy rules, all the remaining model parameters are calibrated prior to estimation. Accordingly, the model predicted impulse response functions (IRFs) can only be affected by the unknown parameters associated with the rival policymaking hypotheses. We thus use the DSGE model as an instrument to help us discriminate between the competing representations of monetary policy.

When we consider the IRFs for output, inflation, wage inflation, the Fed funds rate, and money growth, the general result is that we still find the multiple optima phenomenon. However, using the model as a instrument, we are able to unambiguously discriminate between the two different schemes (partial adjustment versus serially correlated shocks). Our results point out that a configuration with fast partial adjustment and a high degree of serial correlation of monetary shocks minimizes the distance between the actual and predicted IRFs. This configuration also delivers a significantly better fit than the reverse configuration of high inertia and mildly serially correlated shocks.

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Finally, we study the robustness of our findings to the timing of decisions in the DSGE model as well as different calibrations. While the timing proves relatively innocuous, we find that changing the calibration can have dramatic effects when the estimation criterion is informative. In contrast, when the latter conveys no information (i.e. when we exclusively focus on the Fed funds rate), we are no longer able to discriminate between the two competing views. This is a further confirmation of the need to exploit the information contained in the cross-equation restrictions created by our DSGE model.

The remainder is as follows. Section 2 describes the DSGE monetary model that we use for this exercise. Section 3 explains in deeper details the econometric approach employed. Section 4 discusses the main results from the estimation. Finally, the last section offers some concluding comments.

2 Model

We consider a simple New Keynesian model with price and wage stickiness, along the lines of Giannoni and Woodford (2005) and Galí and Rabanal (2005). The latter can be viewed as the benchmark DSGE model typically in use in the literature. Its empirical performances have been assessed through a variety of empirical techniques which confirm the model's goodness-of-fit along several dimensions.

Since our analysis focuses on the empirical performances of alternative representations of monetary policy, we consider a model hit by monetary policy shocks only. Since, later on, we will seek to compare this model with a monetary SVAR in the lines of Christiano et al. (1996, 1999), it is important to make sure that they both 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) and many others. The precise timing of events is described in figure 1.

2.1 Production Side

A large number of competitive firms produce a homogeneous good that can be either consumed (y_t) or used as material goods 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

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

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

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

subject to (1). Solving this program yields the demand addressed to the producer of intermediate good ς

$$G'\left(\frac{d_t(\varsigma)}{d_t}\right) = \frac{P_t(\varsigma)}{P_t} \int_0^1 \frac{d_t(u)}{d_t} G'\left(\frac{d_t(u)}{d_t}\right) \mathrm{d}u. \tag{2}$$

Monopolistic firms produce the intermediate goods $\varsigma \in [0,1]$. Each firm ς is the sole producer of intermediate good ς. Following Rotemberg and Woodford (1995), we assume that monopolist ς produces good ς with the inputs of labor $n_t(\varsigma)$ and material goods $x_t(\varsigma)$ according to the following production possibilities

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

where $F(\cdot)$ is an increasing and concave production function, $n_t(\varsigma)$ denotes the input of aggregate labor (to be defined later), $x_t(\varsigma)$ denotes the input of material goods, and s_x is the share of material goods in gross output. The associated real cost function is

$$S(d_t(\varsigma)) = w_t F^{-1}\left((1-s_x)\, d_t(\varsigma)\right) + s_x d_t(\varsigma),$$

where w_t is the real wage rate.

Let $\theta_p(z)$ denote the elasticity of demand for a producer of intermediate good 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 good firms face a varying elasticity of demand for their output, implying a varying markup, which is denoted by $\mu_p(z) \equiv \theta_p(z)/(\theta_p(z)-1)$. This turns out to be a powerful source of strategic complementarity between price setters, as shown in Woodford (2003).

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. Denote $d^\star_{t,T}(\varsigma)$ the production of good ς in period T if firm ς last reoptimized its price in period t. Then, firm ς chooses $P^\star_t(\varsigma)$ in order to maximize

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

where λ_t is the representative household's marginal utility of wealth, $\mathrm{E}_t\{\cdot\}$ is the expectation operator conditional on information available as of time t, $S(d_t(\varsigma))$ is the real cost of producing $d_t(\varsigma)$ units good of ς, and $d^\star_{t,T}(\varsigma)$ obeys

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

Standard manipulations yield the approximate New Keynesian Phillips curve

$$\hat{\pi}_t - \gamma_p \hat{\pi}_{t-1} = \mathrm{E}_{t-1}\{\kappa_p(\hat{w}_t + \omega_p \hat{y}_t) + \beta(\hat{\pi}_{t+1} - \gamma_p \hat{\pi}_t)\}, \tag{3}$$

with

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

In equation (3), $\hat{\pi}_t$ is the logdeviation of $1+\pi_t$, $\hat{y}_t$ and $\hat{w}_t$ are the logdeviations of y_t and w_t, respectively, $\theta_p \equiv \theta(1)$ is the steady state elasticity of demand for a producer of intermediate good, and the composite parameter ω_p obeys

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

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 $\mu(\xi)$ in the neighborhood of $\xi = 1$, i.e. $\epsilon_\mu = \mu'(1)/\mu(1)$. Conventionally, γ_p is interpreted as a measure of intrinsic inflation persistence, i.e. the backward dimension of inflation.

The composite parameter $\varkappa$ is linked to the degree of strategic complementarity between the price-setting firms. More precisely, the smaller $\varkappa$, the higher the degree of strategic complementarity. When the latter is high, inflation and output turn out to be persistent, i.e. adjust gradually to shocks. As is clear from the above definition of $\varkappa$, we can observe that a positive share of material goods s_x reduces the responsiveness of inflation. Furthermore, ϵ_μ and θ_p play a similar role to that of s_x.

2.2 Aggregate Labor Index and Households

Following Erceg et al. (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 in the sequel. 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

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

where $W_t\left(\upsilon\right)$ denotes the nominal wage rate associated 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 households 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

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

where $\beta \in (0, 1)$ is the subjective discount factor, $b \in (0, 1)$ is the habit parameter, c_t is consumption, $m_t = M_t/P_t$ denotes real cash balances at the end of the period, where M_t denotes nominal cash balances; $h_t\left(\upsilon\right)$ 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 their decisions – see the timing of events described in figure1.

Household υ maximizes his intertemporal utility subject to the sequence of constraints

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

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; Υ_t is a nominal transfer from the government. As in Woodford (2003), we assume that there is a satiation level m^* for real balances such that $U_m = 0$ for $m \geq m^*$. Thus, when m_t reaches m^* 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 loglinearization of the first order conditions associated with c_t, B_t, and M_t yields

$$\mathrm{E}_{t-1}\{\beta b(\hat{c}_{t+1} - b\hat{c}_t) - (\hat{c}_t - b\hat{c}_{t-1}) + \sigma\chi(\hat{m}_t - \beta b\hat{m}_{t+1}) - \varphi^{-1}\hat{\lambda}_t\} = 0, \quad (4)$$

$$\hat{\lambda}_t = \hat{\imath}_t + \mathrm{E}_t\{\hat{\lambda}_{t+1} - \hat{\pi}_{t+1}\}. \tag{5}$$

$$\hat{m}_t = \eta_y(\hat{c}_t - b\hat{c}_{t-1}) - \eta_i \hat{\imath}_t. \tag{6}$$

where $\hat{c}_t$, $\hat{m}_t$, $\hat{\imath}_t$, and $\hat{\lambda}_t$ are the logdeviations of c_t, m_t, $1+i_t$, and λ_t, respectively, and where we defined the auxiliary parameters

$$\sigma^{-1} = -\frac{U_{cc}c}{U_c}, \quad \chi = \frac{U_{cm}m}{U_c}, \quad \varphi^{-1} = (1-\beta b)\,\sigma,$$
$$\eta_y = -\frac{U_{mc}c}{U_{mm}m}, \quad \eta_i = -\frac{(1-\beta b)\,U_c}{U_{mm}m}.$$

Notice that

$$\chi = \frac{1-\beta b}{\bar{v}}\frac{\eta_y}{\eta_i},$$

where $\bar{v}$ is the steady state value of c_t/m_t. Throughout the paper, we enforce this constraint and calibrate $\bar{v}$ to its observed value. Equation (4) illustrates the role played by habits in consumption, which reinforces the backward dimension of the IS curve. Equation (5) is the standard Euler equation on bond holdings. Finally, equation (6) is a standard money demand function.

A typical household υ acts as a monopoly supplier of type-υ 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 simply rule $W_T(\upsilon) = (1+\delta^w_{t,T})W_t(\upsilon)$, 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 υ. In the sequel, it will be convenient to define wage inflation $\pi^w_t \equiv W_t/W_{t-1} - 1$. Now, let $W^\star_t(\upsilon)$ denote the wage rate chosen in date t, and $h^\star_{t,T}(\upsilon)$ denote hours worked in period T if household υ last reoptimized its wage in period t, which obeys the relationship

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

$W_t^*(\upsilon)$ is selected so as to maximize

$$\mathrm{E}_{t-1}\sum_{T=t}^{\infty}(\beta\alpha_w)^{T-t}\left\{\lambda_T\frac{(1+\delta_{t,T}^w)W_t^\star(\upsilon)}{P_T}h_{t,T}^\star(\upsilon)-V(h_{t,T}^\star(\upsilon))\right\}.$$

Loglinearizing the associated first order condition yields

$$\hat{\pi}_t^w-\gamma_w\hat{\pi}_{t-1}=\mathrm{E}_{t-1}\{\kappa_w(\omega_w\hat{h}_t-\hat{\lambda}_t-\hat{w}_t)+\beta(\hat{\pi}_{t+1}^w-\gamma_w\hat{\pi}_t)\} \tag{7}$$

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

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

Here, γ_w governs the sensitivity of nominal wages to inflation, and thus reinforces the persistence of wage inflation; κ_w is a crucial parameter which measures the effect of labor wedges on the nominal wage. Here, we define labor wedges as the discrepancy between the marginal disutility of labor and the marginal gain in consumption units of working an extra hour. Finally, notice that $\hat{\pi}_t$ and $\hat{\pi}_t^w$ are linked together through the relation

$$\hat{\pi}_t^w=\hat{w}_t-\hat{w}_{t-1}+\hat{\pi}_t, \tag{8}$$

which is a simple "accounting" identity.

2.3 Monetary Policy

The model is closed by postulating a monetary policy rule of the form

$$\hat{\imath}_t^*=a_\pi\hat{\pi}_t+a_y\hat{y}_t, \tag{9}$$

$$\hat{\imath}_t=\rho_1\hat{\imath}_{t-1}+(1-\rho_1)\,\hat{\imath}_t^*+e_t, \tag{10}$$

$$e_t=\rho_2e_{t-1}+\nu_t,\quad \nu_t\sim\mathrm{iid}(0,\sigma_\nu^2). \tag{11}$$

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 (3)-(11), we obtain a rational expectations system of linear equations which we solve using standard methods.

Let us now comment on the monetary policy specification. Equation (9) is similar to that proposed by Taylor (1993). Here $\hat{\imath}_t^*$ is the target interest rate that depends on current inflation and (model–based) output gaps. 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 gaps.

To complete the description of monetary policy, we combine two very different views about its behavior. The usual interpretation commonly attributed to equation (10) is as follows: it represents a process of interest–rate smoothing or policy inertia by the central bank, which would operate a partial adjustment process in the announced or realized interest rate. A high level of ρ_1 implies a slow speed of adjustment of the nominal interest rate. Thus, if $\rho_1 = 0.8$ for a 1 % change in the in target rate $\hat{\imath}_t^*$, *ceteris paribus*, the actual interest rate would adjust by 20 basis points in the first quarter, and by around 60 points at the end of the first year. Kozicki (1999), Amato and Laubach (2003), Levin, Wieland and Williams (1999), Sack and Wieland (2000) and Clarida, Galí and Gertler (2000) find that ρ_1 lies in the [0.70,0.90] interval.

The interest rate smoothing hypothesis finds theoretical support in the optimal monetary policy literature. There are at least two views according to which smoothing interest rates may be important for the central bank.

The first is that the central bank preferences may feature a concern for interest rate volatility. Goodfriend (1987) provides an argument supporting this idea by stating that central banks appear to prefer smooth interest rates in order to minimize unexpected asset price movements that raise the risk of bankruptcies and banking crises. In the same tenor, Woodford (2003) justifies the inclusion of the interest rate volatility term into the central bank loss function by arguing that the existence of non negligible transactions frictions or a concern to reduce the frequency with which a linear policy rule would require violation of the zero lower bound on nominal interest rates may affect expected utility maximization.

The second approach supporting interest rate partial adjustment is that observed smoothing of the funds rate may indeed be optimal, even if the central bank is not explicitly concerned with interest rate volatility.[3] Using different structural macroeconometric models with forward-looking

[3] See Sack and Wieland (1999) for further details.

expectations, Levin, Wieland, and Williams (1999) show that the variance of inflation and output gap are reduced in all cases when a certain degree of partial adjustment in the funds rate is supposed.

Another reason for which smoothing the interest rate may be appreciated by policymakers is a concern for avoiding the effects of parameter uncertainty over the variance of inflation and output gap. Sack and Wieland (1999) show that adjusting the policy instrument in a sluggish way may help the central bank to learn about the effects of such policy actions. This might contribute to reducing the scope of the effects of parameter uncertainty.[4]

Rudebusch (2002 and 2005) argues that, overall, the policy rate partial adjustment implies that future changes in the interest rate should be forecastable by agents. As a consequence, actual changes in the term structure given by the yield curve should also be largely predictable. However, using data from financial markets, with changes of future interest rates proxied by the rate of eurodollar deposits, Rudebusch (2002) shows that the predictive power of the expected interest rate changes is unambiguously low compared to what suggests the strong interest rate smoothing typically found in empirical studies. Therefore, Rudebusch concludes, the apparently high degree serial correlation in nominal interest rates could be spuriously attributed to a partial adjustment mechanism toward the targeted level of the Fed Funds rate.[5]

Equation (11) represents an alternative hypothesis to the apparent smoothing of the interest rate as given by (10), and emphasizes the effects of serially correlated policy shocks, e_t, in the realizations of the policy rule. These shocks may represent any contingent event the central bank faces when deciding the interest rate, such as credit crunches or financial crises, see Taylor (1993) or Rudebusch (2002). For example, the former pointed out that the Federal Reserve had to adjust monetary policy as to provide additional reserves to the banking system after the stock-market break of October 19, 1987 and helped to prevent a contraction of liquidity and restore confidence. Rudebusch (2002) describes the effects of the persistent 1992-1993 credit crunch in the U.S. and the response of the Fed by holding unusual low levels in the interest rates. In the same flavor, Rudebusch suggests how the 1998 financial crisis started in Russia played a large role in lowering the interest rates

[4] Notice that the empirical results of parameter uncertainty over the optimal policy choice of the interest rate have had mixed results. See Sack (1998) and Rudebusch (2000).

[5] See also Söderlind et al. (2003).

in 1998 and 1999, not in accordance with the target rate given by the Taylor rule. Finally, the use of real-time data could also reinforce the apparent degree of serial correlation in policy shocks. Indeed, the incomplete information used when estimating these parameters is modified by revisions over time, thus affecting the policy rate level in a persistent way, as argued by Orphanides (2004).

Importantly, this alternative view of monetary policy yields implications that are completely opposite to those of the partial adjustment hypothesis. Under serially correlated shocks, the central bank does not effectively smooth the interest rate, even though the latter might prove persistent. Therefore, the predictive power of the term structure about the future variations of the funds rate need not be strong in this case, since the latter are driven by the innovation ν_t. Provided that these shocks are not the main sources of aggregate fluctuations (as is a generally accepted view), shocks to monetary policy will not play a large role in the predictability of the term structure.

In the sequel, we will refer to the configuration with a large ρ_1 and a small ρ_2 as the *smoothing dominant policy*. Symmetrically, the configuration with a large ρ_2 and a small ρ_1 will be referred to as the *contingent dominant policy*.

3 Econometric Approach

This section describes our methodological approach. More precisely, we detail our monetary SVAR so as to compute a set of IRFs used to estimate the DSGE model. We then expound the MDE approach and discuss some practical issues relating to the multiple optima problem.

3.1 Using the DSGE Model as an Instrument

The purpose of our paper is to provide tools to help us discriminate between the two competing views of monetary policy: interest rate inertia versus persistent shocks. To do so, we use the DSGE model previously expounded as an instrument. More precisely, we use the cross-equation restrictions embedded in the model as a way to quantitatively assess the relative performances of these two alternative representations of monetary policy.

Let ψ denote the whole set of model parameters. Let $\psi_2 = (\rho_1, \rho_2, \sigma_\nu)'$ 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 model result only from changes in ψ_2. To see this, let us consider the reduced form model solution. The latter can generically write as

$$Y_t = A(\psi_1, \psi_2) Y_{t-1} + B(\psi_1, \psi_2) \nu_t,$$

where Y_t contains all the endogenous and exogenous variables as well as any lags thereof necessary. According to this equation, the model dynamics depends on both ψ_1 and ψ_2. In the experiment which we propose, it is necessary to control for the impact of ψ_1 on the model dynamics, so that the model can be interpreted as an instrument. As a consequence, only variations in ψ_2 affect the quantitative performances of the DSGE model, thus revealing information about the relevant specification of the monetary policy rule.

To make this controlled experiment implementable, we now need a plausible measure of monetary policy shock from actual data. To do so, we follow a large literature on the identification of monetary policy shocks.

3.2 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 et al. (1996, 1999) so as to identify monetary policy shocks.[6] We first assume that monetary authorities set their instrument, $\hat{\imath}_t$ (here the Federal Funds rate), according to the policy rule

$$\hat{\imath}_t = f(\Omega_t) + \sigma_i \epsilon_t,$$

where Ω_t is the information set available at the time monetary authorities take their decisions, σ_i is a positive scalar, and ϵ_t is a white noise monetary shock orthogonal to the elements generating Ω_t. Formally, let us consider the data vector of dimension m

$$Z_t = \left(Z_{1,t}', \hat{\imath}_t, Z_{2,t}'\right)'.$$

[6] See also Christiano et al. (1997, 2005), and Rotemberg and Woodford (1997, 1999) for other examples of this identifying strategy.

The vector $Z_{1,t}$ is a $n_1 \times 1$ vector composed of variables whose current and past realizations are included in Ω_t and that are assumed to be predetermined with respect to ϵ_t. $Z_{2,t}$ is a $n_2 \times 1$ vector containing variables that are allowed to respond contemporaneously to ϵ_t but whose value is unknown to monetary policy authorities at t. Thus only lagged values of $Z_{2,t}$ appear in Ω_t. Accordingly, $m = n_1 + n_2 + 1$.

So as to implement this identification strategy, we first estimate an unconstrained Vector Autoregression (VAR) of the form

$$Z_t = B_1 Z_{t-1} + \cdots + B_\ell Z_{t-\ell} + u_t, \quad \mathrm{E}\{u_t u_t'\} = \Sigma,$$

where ℓ is the maximal lag, which we determine by minimizing the Hannan-Quinn information criterion. In our empirical analysis, we found that $\ell = 4$. Now in order to recover the structural shock to monetary policy ϵ_t, we assume that the canonical innovations u_t are linear combinations of the structural shocks η_t, i.e.

$$u_t = S\eta_t,$$

for some non singular matrix S. As usual, we impose an orthogonality assumption on the structural shocks, which combined with a scale normalization implies $\mathrm{E}\{\eta_t \eta_t'\} = I_m$, where I_m is the identity matrix and m is the number of variables in Z_t.

With the recursiveness assumptions above, a monetary policy shock can be recovered as follows. Set S to be the Cholesky factor of Σ, so that $SS' = \Sigma$. Then, σ_i is the $(n_1 + 1, n_1 + 1)$ element of S, and ϵ_t is the shock appearing in the $(n_1 + 1)$th equation of the system

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

where $A_0 = S^{-1}$ and $A_i = S^{-1} B_i$, $i = 1, \ldots, \ell$.

3.3 Minimum Distance Estimation

As in a large strand of the literature that follows the original work by Rotemberg and Woodford (1997), we estimate the policy parameters ψ_2 by minimizing a measure of the distance between the empirical responses of key aggregate variables and their model counterparts.[7]

More precisely, we focus our attention on the responses of the vector X_t regrouping the actual data which we are interested in. Here, X_t is a subset of Z_t, which we define with the relation $X_t = \Gamma_{XZ} Z_t$, where Γ_{XZ} is a selection matrix. Then, we define θ_j the vector of responses to a monetary shock at horizon $j \geq 0$, as implied by the above SVAR estimated on actual data, i.e.

$$\theta_j \equiv \frac{\partial X_{t+j}}{\partial \epsilon_t} = \Gamma_{XZ} \frac{\partial Z_{t+j}}{\partial \epsilon_t}, \quad j \geq 0,$$

where ϵ_t is the monetary policy shock previously identified.

Then, the object which we seek to match is $\theta = \text{vec}([\theta_0, \theta_1, \ldots, \theta_k])'$ where k is the selected horizon. Notice that we have to exclude from θ_0 the responses corresponding to the elements in X_t that belong to Ω_t. It is important to emphasize that the DSGE model previously expounded embeds the same exclusion restrictions as the monetary SVAR.[8] Then let $h(\cdot)$ denote the mapping from the structural parameters $\psi_2 = (\rho_1, \rho_2\, \sigma_\nu)'$ to the DSGE counterpart of θ. Our estimate of ψ_2 is solution to the problem

$$\hat{\psi}_2 = \arg\min_{\psi_2 \in \Psi} (h(\psi_2) - \hat{\theta}_T) V_T (h(\psi_2) - \hat{\theta}_T)',$$

where $\hat{\theta}_T$ is an estimate of θ, T is the sample size, Ψ is the set of admissible values of ψ_2, and V_T is a weighting matrix which we assume is the inverse of a matrix containing the variances of each element of θ along its diagonal and zeros elsewhere. These variances are obtained from the SVAR parameters.

For further references, let us define the objective function at convergence

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

Under the null hypothesis, as shown in Hansen (1982), $\mathcal{J}_T \sim \chi^2(\dim(\theta) - \dim(\psi_2))$. Given our choice of weighting matrix, we can further decompose $\mathcal{J}_T$ into components pertaining to each element of X_t, according to

$$\mathcal{J}_T = \sum_{i=1}^{\dim(X)} \mathcal{J}_{T,i}.$$

[7] See also Altig et al. (2004), Amato and Laubach (2003), Boivin and Giannoni (2005), Christiano et al. (2005), and Giannoni and Woodford (2004).

[8] We assess the sensitivity of our results to the timing restrictions in the next section.

The latter decomposition provides a simple diagnostic tool that allows us to locate on which dimension the model succeeds or fails to replicate the impulse response functions implied by the SVAR.

3.4 Practical Issues

The empirical literature on Taylor rules has had trouble reaching clear-cut conclusion as to 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. English et al. (2002) study the different implications of partial adjustment and serially correlated errors over the first difference of the interest rate realizations. In particular, they found that there is supportive evidence that both partial adjustment and serially correlated errors are important elements for the understanding of Federal Reserve behavior since the late 1980's, and that both play an important role in explaining deviation of the Federal funds rate from the Taylor–rule.[9]

Rudebusch (2002) estimates the two competing representations of monetary policy within a single-equation framework. He concludes that it is not possible to distinguish the relative importance of the partial adjustment parameter from the serially correlated parameter. This is the reason why he develops an indirect test, relying on extra financial market information as instrument, about the importance of interest rate smoothing.

This absence of 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. Griliches, 1967, Blinder, 1986, McManus et al. 1994). Our approach, based on rational expectation econometrics, suffers from the same problems, especially when the framework conveys little information, as in Sargent (1978) or Kennan (1988).

To see this problem most clearly, let us consider our simple representation of the monetary policy combining partial adjustment and serially correlated shocks

$$\hat{\imath}_t = (\rho_1 + \rho_2)\hat{\imath}_{t-1} - \rho_1\rho_2\hat{\imath}_{t-2} + (1-\rho_1)\left(\hat{\imath}_t^\star - \rho_2\hat{\imath}_{t-1}^\star\right) + \nu_t,$$

[9] See also Castelnuovo (2003), Gerlach–Kristen (2004), Apel and Jansson (2005) for further developments.

where the target $\hat{\imath}_t^\star$ is a linear function of lagged shocks to monetary policy. Indeed, according to the timing of decisions in our model, output and inflation cannot respond to a monetary shock ν_t on impact, so that

$$\hat{\imath}_t^\star = \sum_{k=1}^{\infty} \eta_k(\psi_2)\nu_{t-k},$$

where $\eta_k(\psi_2)$ is a complicated nonlinear function of the policy rule parameters ψ_2, as well as the remaining deep parameters ψ_1, though this latter point is not directly reflected in our notations. Suppose that $\eta_k(\psi_2)$ for $(k = 1, ..., \infty)$ are small and not sensitive to any change in ρ_1 and ρ_2. In this case, $\hat{\imath}_t^\star$ displays only a very small amount of variance. In our estimation setup, which is based on IRFs, this implies that $\hat{\imath}_t^\star \approx 0$, $\forall t$ and the policy function accordingly rewrites

$$\hat{\imath}_t \approx (\rho_1 + \rho_2)\hat{\imath}_{t-1} - \rho_1\rho_2\hat{\imath}_{t-2} + \nu_t.$$

In this case, the parameters ρ_1 and ρ_2 are not identified. To see this, consider the reduced form associated to the monetary policy

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

Since the IRFs of the Fed fund rate $\hat{\imath}_t$ can be directly deduced from β_1 and β_2, this reduced form completely determines the objects to be matched using our MDE approach. The parameters ρ_1 and ρ_2 are only identified when $\rho_1 = \rho_2$. However, this case is inconclusive as it puts the same weights on the two competing representations of the monetary policy. Except for this very special case, there does not exist a unique solution for ρ_1 and ρ_2 as a function of the reduced form parameters β_1 and β_2. Indeed, provided that $\beta_2 \neq 0$, the solutions for ρ_2 are given by

$$\rho_2 = \frac{\beta_1 \pm \sqrt{\beta_1^2 + 4\beta_2}}{2}$$

where $\beta_1^2 + 4\beta_2 = (\rho_1 - \rho_2)^2 \geq 0$. This means that two sets of values for ρ_1 and ρ_2 are observationally equivalent. The first solution is a *smoothing dominant* optimum (ρ_1 large and ρ_2 small). The second one is a *contingent dominant* optimum (ρ_1 small and ρ_2 large). Thus when $\hat{\imath}_t^\star \approx 0$, we cannot distinguish between an inertial monetary policy with transitory shocks and a monetary policy with small partial adjustment and highly serially correlated shocks.

When $\hat{\imath}_t^\star$ is responsive to shocks and thus more volatile, this problem of multiple optima can potentially disappears, provided that $\eta_k(\psi_2)$ is highly sensitive to perturbations in ρ_1 and ρ_2. However, nothing guarantees this in practice, so that estimating ρ_1 and ρ_2 by focusing on the nominal interest rate only might fail to reveal the correct information about monetary policy.

Another way to escape this problem is to consider additional variables in our MDE estimation. Rather than using a single variable, which in our case would be akin to considering a single equation framework, we introduce a complementary set of variables. When the policy rule parameters have strong effects on aggregate dynamics, due to the cross-equation restrictions embedded in our model, this give us an opportunity to properly identify ρ_1 and ρ_2. This has the potential to deliver clear-cut conclusions. This is what is to be investigated in the sequel.

4 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 expound our estimation results and explain how to use them in order to discriminate between the two competing views about monetary policy. Finally, we provide some sensitivity analyses.

4.1 Data and SVAR

In addition to the Fed Funds rate, we use data from the Non Farm Business (NFB) sector over the sample period 1960(1)-2002(4).[10] The variables used for estimation are the linearly detrended

[10] Arguably, this sample period might be characterized by significant changes in monetary policy. As a consequence, assuming 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(1)-2002(4) exhibit a number of pathologies. For example, output persistently rises after a contractionary monetary policy shock. In addition, the estimated IRF 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 et al. (1996, 1999, 2005) and adopt a longer sample. In addition, Sims and Zha (2002) 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).

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 compensation, $\hat{\pi}^w_t$.[11] We also include two "information" variables in the model. First, though not formally justified by the theoretical model, the growth rate of the logarithm of the CRB price index of sensitive commodities, $\hat{\pi}^c_t$, is included to mitigate the so-called price puzzle.[12] Second, the growth rate of M2, $\hat{\xi}_t$, is included to exploit some potential information included in money growth.[13] The data are reported on figure 2.[14] So as to implement the identification strategy outlined above, we set

$$Z_{1,t} = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}^w_t, \hat{\pi}^c_t)', \quad Z_{2,t} = (\hat{\xi}_t).$$

In addition, the variables of interest, X_t are defined as

$$X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}^w_t, \hat{\imath}_t, \hat{\xi}_t)'.$$

The empirical responses of X_t are reported on figure 3, with $k = 31$. The plain line is our point estimates of the empirical responses of X_t and the shaded areas indicate the 95% confidence interval about the point estimates. Following Hamilton (1994), the latter are obtained by direct numerical integration.

Though we focus on a different dataset and a different sample period, our findings echo previous results reported by Christiano et al. (1996, 1997, 1999, 2005).[15] In particular, figure 3 exhibits all the usual features of the responses to a monetary policy shock. Output initially responds very little, and then sharply drops, with a hump pattern. Notice that the latter is precisely estimated. When it comes to inflation, we obtain a small and not significant price puzzle in the very short-run, as has been previously documented in the literature, followed by a protracted decline displaying

[11] The civilian non-institutional population over 16 is used as our measure of population. We also experimented with quadratically detrended or first-differenced output, without quantitatively altering our conclusions.

[12] This is a fairly standard practice in the literature. See Sims (1992), Eichenbaum (1992), Christiano et al. (1996, 1999).

[13] We also experimented with M1 instead of M2, without altering our findings.

[14] The data are extracted from the Bureau of Labor Statistics website, except for the Fed Funds rate and M2 which are obtained from the FREDII database.

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

a persistent hump-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. The response of wage inflation is qualitatively similar, with a trough response slightly lagging that of inflation. As discussed at length 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, up to the point where it will eventually cross the x axis. Finally, nominal money growth drops sharply and rapidly reaches back its steady state level.

As in Christiano et al. (2005), our empirical strategy for discriminating between alternative representations of monetary policy relies on a single source of aggregate fluctuations, namely a shock to monetary policy. Though our primary purpose is not to provide a general model of business cycle fluctuations, it is legitimate to wonder what portion of aggregate fluctuations are accounted for by this monetary shock. To answer this question, we report in table 1 the percentage of variance of the k-step-ahead forecast errors of X_t. The number in brackets are the boundary of the 95% confidence regions obtained by standard bootstrap methods, using 1000 simulations.

As is clear, monetary policy shocks account only for a tiny portion of the variance of output, inflation, and wage inflation when the forecast horizons are short ($k = 4$), and a substantially bigger portion at longer horizons ($k = 20, 30$), comprised between 10% and 20%. This result confirms a host of previous studies, suggesting that monetary policy shocks might not be the dominant source of business cycle fluctuations.

Abstracting from this variance decomposition, what turns out to be important and informative for our purpose is to obtain precisely estimated responses of aggregate variables, especially so when it comes to the typical hump and persistence patterns previously discussed.

4.2 Parameters 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 impulse response functions depend only on the particular specification of monetary policy.

Of course, the model features a large number of parameters, accounting for preferences, technology, and market frictions. Calibrating those can be source of arbitrariness. As a first requirement, it is thus important to make consensual and conservative choices for these parameters. A second requirement is that we must make sure that the chosen calibration does not imply implausible responses of the model's counterparts of X_t.[16] The calibration is reported in table 2.

Preferences. First, we set $\beta = 0.99$, implying a steady state annualized real interest rate of 4%. The habit persistence parameter b is set to 0.75, lying in the range of available estimates based on aggregate data (see Boldrin et al., 2001, Jermann, 1998, Boivin and Giannoni, 2005, Christiano et al., 2005). We then set $\sigma = 1 - b$, which implies intertemporal complementarities in consumption decisions. Notice that this value is close to that retained by Rotemberg and Woodford (1997).

As in Christiano et al. (2005) and Altig et al. (2004), the elasticity of marginal labor disutility, ω_w, is set to 1. The money demand function implied by our model is of the form

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

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 et al. (2005). For example, their counterpart of η_i is estimated to a very small value compared to previous estimates in Lucas (1988), which leads them to interpret this parameter as a short-run semi-elasticity. To pin down the value of η_i, we follow a different approach, yielding very similar results.

From the SVAR and identified monetary shocks, we generate a sample path of X_t conditional on ϵ_t only. More precisely, we construct data series for real balances ($\tilde{m}_t$), real output ($\tilde{y}_t$), and the nominal interest rate ($\tilde{\imath}_t$) when the SVAR is subject to monetary shocks only. We then estimate a linear money demand function using OLS. Because the SVAR needs historical initial conditions, we loose four points of data in our sample, which accordingly ranges from 1961(1)-2002(4). The

[16] The sensitivity of our results to the calibration is assessed in section 4.4.

estimated money demand takes the form

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

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 account of the calibrated value of b, thus partly taking into account the money demand equation (6) from the model. The implied long-run semi elasticity is slightly above 8, which is the value obtained by Lucas (1988), Chari et al. (2000), 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 et al. (2005).

Recall that $\chi \equiv U_{cm}m/U_c$ 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%, after correcting for the markup. Assuming further that F is Cobb-Douglas, direct calculations yield $\omega_p = \phi - 1$. The share of material goods in gross output, s_x, is set to 50%, as in Rotemberg and Woodford (1995) and Basu (1995). Following Rotemberg and Woodford (1997) and Christiano et al. (2005), we set the markup on prices to 20%, 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). Interestingly, this value implies that a 1% increase in relative prices translates into a 6.7% decrease in demand – the corresponding constant elasticity of substitution production function would imply a 5.8% decrease in demand. Finally, we set θ_w to 21 as in Christiano et al. (2005), implying a wage markup of 5%.

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 et al. (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 et al. (1995), we also set $\gamma_w = 1$. Once again, this enhances the model's ability to generate a large amount of persistence for nominal wage inflation.

Nominal Interest Rate Target Level. Following Taylor (1993), we set $a_\pi = 1.5$ and $a_y = 0.5/4$, since we focus on a quarterly measure of the output gap. These values are approximately the same as those considered by Christiano et al. (2005) in their sensitivity analysis.

4.3 Estimation Results

The estimation results are reported in table 3, for different X_t and different restrictions on the policy rule parameters. In each case, we set the impulse response functions horizon k to 31.[17] The table is organized as follows. The left panel reports parameters 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 Fed Funds rate, and money growth to a monetary policy shock as identified in the SVAR. The right panel corresponds to the case where $X_t = \hat{\imath}_t$, i.e. when we exclusively focus on the Fed Funds rate's behavior. In each panel, we consider four cases, depending on the minimum value of $\mathcal{J}_T$ reached at convergence and on zero restrictions on ρ_1 or ρ_2. More precisely, column (1) corresponds to the minimum value of $\mathcal{J}_T$ reached when using as an initial condition a large ρ_2 and a small ρ_1. Conversely, column (2) corresponds to the case with a large ρ_1 and a small ρ_2. Indeed, as explained above, the estimation of a partial adjustment model with serially correlated shock raises well-known multiple optima issues. Column (3) corresponds to the restriction $\rho_1 = 0$, i.e. to a model with only serially correlated shocks and no partial adjustment. Column (4) corresponds to the restriction $\rho_2 = 0$, i.e. to a model with nominal interest rate inertia and *iid* shocks to monetary policy. Finally, column (5) reports the estimation outcome when imposing the constraint $\rho_1 = \rho_2$, thus granting the same weight on both alternative views about monetary policy.

The point estimates of ψ_2 are reported together with their standard errors, in parentheses. The

[17] Later, we investigate the sensitivity of our results to k.

table also reports the value of $\mathcal{J}_T$ at convergence, together 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. This allows us to assess on which particular dimension the model fails or succeeds to replicate the data.

Let us first consider the case with $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}^w_t, \hat{\imath}_t, \hat{\xi}_t)'$. In practice, we encounter the "multiple local optima" problem. This is revealed in two different ways. First, we can change the initial condition of the estimation algorithm, which leads us to converge to different optima. Second, we can reveal the problem by a direct search on a fine grid of values for ρ_1 and ρ_2. Here, we take advantage of our parsimonious econometric approach which allows us to evaluate the objective function in a small dimensional space.

When we set ρ_1 large and ρ_2 small, we converge toward a local *smoothing dominant* minimum. Conversely, when we set ρ_2 large and ρ_1 small, we converge toward a local *contingent dominant* minimum. However, comparing the $\mathcal{J}_T$'s from columns (1) and (2) clearly shows that the *contingent dominant* case is the global minimum of $\mathcal{J}_T$. Notice also that there is a huge difference between the two objective functions at convergence (145 versus 245) and that one version is blatantly rejected by the data (column 2) while the other successfully passes the over-identification test (column 1). Thus, our results suggest that the data favor a model with serially correlated monetary shocks and a modest degree of interest rate inertia.

As explained above, another way to reveal this problem of "multiple local optima", is to compute the objective function $\mathcal{J}_T$ for different values of (ρ_1, ρ_2). The outcome of this exercise is reported on figure 4. This figure provides a contour plot of $\mathcal{J}_T$ in the (ρ_1, ρ_2) plane. The point denoted by a star represents the global minimum (ρ_2 large, ρ_1 small) while that denoted by a triangle corresponds to the local minimum (ρ_1 large, ρ_2 small). This figure clearly illustrates that initial conditions matter for inference about the true model.

The truly global minimum distance estimator yields $\rho_1 = 0.30$ and $\rho_2 = 0.87$. This suggests that the correct representation of monetary policy is a mix of persistent, serially correlated shocks and a modest degree of partial adjustment, in accordance with Rudebusch (2002, 2005). Notice that these two parameters are found to be significant. However, as emphasized by McManus et al.

(1994), in presence of multiple optima, the Wald test statistics can be poorly approximated by their asymptotic distribution. Conversely, the likelihood ratio test is less subject to this critique. When comparing columns (1) and (3), where we impose the restriction $\rho_1 = 0$, we see that the partial adjustment, though moderate, is essential. Indeed, a likelihood ratio test would overwhelmingly reject the latter restriction. Notice that ρ_2 is not significantly different from zero in the case of the *smoothing dominant* local minimum.

These results are once again illustrated in figure 4. The point denoted by a pentagram corresponds to the restriction $\rho_1 = 0$ (i.e. columns 3) while that denoted by a square corresponds to the restriction $\rho_2 = 0$ (i.e. column 4). Clearly, the restriction $\rho_1 = 0$ substantially worsens the model fit, without implying a rejection since the associated P-value is equal to 28%.

To understand why the model with large partial adjustment and almost *iid* shocks is rejected by the data, 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, nominal interest rate, and money growth. In other words, these three variables are weakly informative about the relevant form of monetary policy.[18] What turns out to be really discriminating is the behavior of inflation and wage inflation. In this case, the partial adjustment model proves unable to mimic the delayed and persistent responses of these variables.

This failure is illustrated by comparing figures 5 and 6. In each figure, the solid line represents the SVAR-based responses of X_t to a monetary policy shock, the grey area is the corresponding asymptotic 95% confidence region, and the solid lines marked with a circle correspond to the DSGE point estimates. The dynamic responses of output, the Fed Funds rate, and money growth do not appear to be qualitatively affected by the specification of monetary policy. To the contrary, the model's impulse response functions of inflation and wage inflation sharply differ. The model with persistent shocks and moderate interest rate inertia successfully matches the essential features of

[18] When we estimate the model with $X_t = (\hat{y}_t, \hat{\imath}_t, \hat{\xi}_t)'$, we cannot discriminate between the two local minima. To the contrary, when we set $X_t = (\hat{\pi}_t, \hat{\pi}^w_t, \hat{\imath}_t)'$, the local minimum with ρ_2 small and ρ_1 large is unambiguously rejected while the global minimum with ρ_2 large and ρ_1 small successfully passes the overidentification test (P-value of 92%).

the data. This is no longer the case when we consider a model with a large degree of interest rate inertia, especially when it comes to inflation and wage inflation.

Column (5) shows that the restriction $\rho_1 = \rho_2$ 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 which grants the same weights to partial adjustment and serially correlated shocks provides a fit which is substantially worse than that of the global maximum. A formal likelihood ratio test would unambiguously reject this restriction.

Second, let us now consider the case with $X_t = \hat{\imath}_t$. We investigate this case 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 focussing on a single policy rule equation, see Rudebusch (2002, 2005).

Now, we face a more severe "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 $\mathscr{J}_T$ statistic. See columns (1)–(4) in the right panel of table 3. These findings are further confirmed by the contour plot reported in figure 7.

This exercise illustrates that focussing 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 when it comes to inflation and wage inflation) implied by the alternative specifications of monetary policy.

Finally, the last column reports the estimation outcome when we impose $\rho_1 = \rho_2$. In contrast with what previously obtained, this representation of monetary policy is not rejected by the data, though this restriction might not be the one favored by the data when we use a larger set of variables.

The previous results are obtained for an horizon $k = 31$. Under this assumption, we were able to discriminate between two competing representations of monetary policy because a model with large interest rate inertia fails to mimic the delayed hump-shaped response of inflation and wage inflation. To further illustrate the information contained in the hump-shape pattern, we now vary the horizon k between 10 and 40. Figure 8 reports the $\mathcal{J}_T$ statistic as well as its decomposition according to the elements of X_t. In this exercise, we select $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}_t^w, \hat{\imath}_t, \hat{\xi}_t)'$. In each panel, the plain line corresponds to the value of the objective function $\mathcal{J}_T$ as well as its decomposition in the case of a *smoothing dominant* optimum while the dashed line corresponds to the case with a *contingent dominant* optimum. 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 ($k = 10, ..., 15$), the two representations of monetary policy yield comparable results. Clearly, focussing 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 pattern of inflation and wage inflation (see the third and fourth panels), the performances of the two competing versions start to dramatically diverge. In particular, the *smoothing dominant* specification faces more and more troubles reproducing the data.

4.4 Sensitivity Analysis

4.4.1 The Role of Timing

The model with the two alternative representations of monetary policy is estimated under the timing assumption outlined in figure 1. The latter ensures the model consistency with the assumptions needed to identify a shock to monetary policy in our SVAR. Here, we want to assess to what extent our results hinge upon this particular timing of events.

Using the estimated parameters values $\hat{\psi}_2$, we compute the impulse response functions of $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}_t^w, \hat{\imath}_t, \hat{\xi}_t)'$ when output, inflation, and wage inflation are allowed to contemporaneously respond to a monetary shock. We report in figures 9 and 10 the responses obtained under the alternative two timing assumptions, for the *contingent dominant* and the *smoothing dominant* cases, respectively. The line with a circle corresponds to the model with timing restrictions on output, inflation, and wage inflation while the simple line corresponds to the case without timing restrictions.

As is clear from these pictures, the role of timing is qualitatively inessential when it comes to output, inflation, wage inflation, and the Fed Funds responses. While the first three variables are now allowed to respond on impact, their persistence and hump pattern are barely modified. Notice that in figure 9, the response of the Fed Funds rate is slightly less pronounced, due to the impact response of output, according to the Taylor rule. To the contrary, in figure 9, because of the large degree of inertia in the rule, the discrepancy between the responses of $\hat{\imath}_t$ under the two alternative timing assumptions is less marked. Finally, the role of timing has a more pronounced effect on money growth, due to the fact that output can respond on impact, thus driving money demand to an even lower value than before. Notice that without timing restrictions, the large surge of money growth in the second period is a direct consequence of habit persistence in consumption. This analysis thus suggests that our previous findings are relatively insensitive to the timing assumptions, essentially because the hump patterns of inflation and wage inflation still obtain under this alternative timing of decisions. Thus the discriminating power of our model is left unaffected.

4.4.2 Sensitivity to Calibration

The empirical assessment of the two alternative representations of monetary policy previously obtained relies on a particular calibration of the structural parameters collected in ψ_1. We now want to check whether the previous findings crucially depend on this particular calibration. In other words, we want to assess whether this calibration biased our results in favor of the serially correlated shock representation of monetary policy.

One simple way to assess the importance of our calibration is to redo our analysis perturbing some key model parameters. We can claim that calibration does not matter if qualitative inference is robust to such parameter perturbations.

At the same time, if we always obtain that at least one monetary policy representation is never rejected by the data, regardless of which calibration is used in the analysis, then we would conclude that our procedure is not very informative. Thus great caution should be taken in interpreting our empirical findings. Fortunately, this is not the case when the procedure features sufficiently demanding over-identification tests.

Tables 4 and 5 report the outcome of this sensitivity analysis. We identify key parameters governing the dynamic behavior of our model, which we partition according to preferences, technology, price/wage setting, and the nominal interest rate target level.

For each alternative parameter value, we recompute the $\mathcal{J}_T$ statistic at convergence, as well as its decomposition according to $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}^w_t, \hat{\imath}_t, \hat{\xi}_t)'$ (see table 4) or $X_t = (\hat{\imath}_t)$ (see table 5). As has been previously shown, the first set of variable is sufficiently informative to allow us to discriminate between the two competing representations of monetary policy. To the contrary, we showed that using only the Fed Funds rate did not help to reach a clear-cut conclusion. In addition, if the procedure with $X_t = (\hat{\imath}_t)$ turns out to be insensitive to our calibration choices, contrary to $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}^w_t, \hat{\imath}_t, \hat{\xi}_t)'$, this constitutes an important warning against the use of a single-equation approach.

Except for some particular cases, we always encounter the "multiple local optima" problem discussed above. This is why tables 4 and 5 report two sets of results for the *contingent dominant* and the *smoothing dominant* cases, respectively.

Preferences. The "Preferences" panel of table 4 reports the effect of shutting habit formation down (i.e. $b = 0$). This has the obvious effect of dramatically worsening the model performance, especially when it comes to $\hat{y}_t$ and $\hat{\imath}_t$. Notice that in this case, the two representations are unambiguously rejected by the data.

Following Giannoni and Woodford (2005), we drastically decrease the elasticity of labor supply, setting $\omega_w = 10$. In such a case, the model performances are always improved, but the *smoothing dominant* case is still rejected.

Finally, we increase the sensitivity of money demand to the nominal interest rate, i.e. $\eta_i = 3$. The model performances are severely affected since now it barely passes the over-identification test, due to the bad behavior of money growth in the short-run.

In contrast, as shown in the "preferences" panel of table 5, when we focus exclusively on the nominal interest rate ($X_t = (\hat{\imath}_t)$), none of the alternative two representations of monetary policy 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 both 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 4, we investigate the sensitivity of our results to perturbations on technology parameters. Following Galí and Rabanal (2005), we assume constant returns to scale in labor input, thus imposing $\phi = 1$ (and $\omega_p = 0$ as a matter of consequence). The model's performances are improved for both specifications of monetary policy. However, the *smoothing dominant* case is again rejected. We also modify the markups on prices without affecting much our results. To the contrary, when we increase the degree of market power on the labor market, we substantially reduce the ability of the model to reproduce the impulse response functions of X_t. Under this assumption, both versions are rejected by the data.

Once again, when we focus on $X_t = (\hat{\imath}_t)$ ("technology" panel of table 5), we fail to reject any of the two competing representations of monetary policy.

Price/Wage Setting. In the "Price/Wage Setting" panel of table 4, we experiment with altering the details of the price and wage setting side of the model. Two exercises are considered. First, we 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. Notice that in the case of perfect price flexibility, the "multiple optima" problem completely disappears. More precisely, when $\alpha_p = 0$, the unique minimum implies a large ρ_2 (0.80) and a small ρ_1 (0.28). In these cases, the dynamics of inflation are less smooth, so that the target level is more volatile, thus conveys more information. Notice however that in these cases, the model is strongly rejected by the data.

Contrary to the above analysis, when we focus on $X_t = (\hat{\imath}_t)$ ("Price/Wage Setting" panel of table 5), we cannot reject any of the two competing representations of monetary policy, which prove almost completely insensitive to such parameters perturbations. Once again, this illustrates the need for further information. Notice also that when $\alpha_p = 0$ or $\alpha_w = 0$, the multiple optima problem also disappears with $X_t = (\hat{\imath}_t)$, for the same reasons as those outlined above.

Nominal Interest Rate Target Level. Finally, in the "Target Level" panel of table 4, 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 an extreme value compared to standard estimates available in the literature, $a_\pi = 3$. In this case, the discrepancy between the two alternative specifications of monetary policy widens, especially 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, which limits the extent of the "multiple optima" problem. When it comes to a_y, the quantitative findings are left unaffected.

Finally, when we focus on $X_t = (\hat{\imath}_t)$ ("Target Level" panel of table 5), we fail to reject any of the two competing representations of monetary policy. This is all the more troubling as one would have expected that our quantitative results might have proved sensitive to large perturbations on the target level parameters. When we focus only on $\hat{\imath}_t$, the discriminating power of inflation and wage inflation is shut down, which keeps us away from reaching a clear-cut conclusion.

4.4.3 Alternative Taylor Rules

To conclude our sensitivity analysis, we now consider an alternative specification of the target level in the Taylor rule.[19] We follow Batini and Haldane (1999) and Clarida et al. (2000) and specify a forward-looking Taylor rule. The new target level is given by

$$\hat{\imath}_t^* = a_\pi \mathrm{E}_t\{\hat{\pi}_{t+4}\} + a_y \mathrm{E}_t\{\hat{y}_{t+4}\}.$$

Using a similar econometric methodology and a closely related DSGE model, Boivin and Giannoni (2005) have shown that forward-looking Taylor rule deliver a better fit to US data. In our context, this type of rules is particularly interesting. Indeed, since it incorporates expectations of future

[19] We have also investigated Taylor rules featuring lagged values of inflation and output in addition to current values, in the line of Smets and Wouters (2003, 2005). However, the additional parameters implied by these extended Taylor rules only marginally improved the fit without altering our previous conclusions.

inflation and output gap, the target level is allowed to react contemporaneously to the monetary innovation. This creates additional variability in the target which has the potential to eliminate the multiple optima problem, thus yielding a better identification of monetary policy.

Results are reported in table 6, the structure of which is similar to that of table 3. An important exception is that under this alternative monetary rule, the multiple optima problem completely disappears when we include sufficiently informative variables, i.e. $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}_t^w, \hat{\imath}_t, \hat{\xi}_t)'$. Accordingly, in the first panel, columns 1 and 2 are identical. This table corroborates our prevous findings. Additionally, it confirms that identification problems in our partial adjustment model with serially correlated shocks totally vanish when sufficiently volatile regressors (here the target level) are included. Once again, our results favor the *contigent-dominant* view of monetary policy. However, as previously obtained, a moderate but significant smoothing of nominal interest rate is necessary to match the monetary SVAR.

In addition, the second panel of table 6 reports estimation results when $X_t = (\hat{\imath}_t)$. In spite of the additional volatility provided by the forward rule, this framework fails to deliver a clearcut conclusion about the proper way to model monetary policy. This is again a consequence of a single-equation approach, which is insufficiently informative.

5 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 cross-equation restrictions contained in our monetary DSGE model. 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 featuring a high degree of inertia is rejected by the data while a rule hit by highly serially correlated shocks satisfactorily matches the data. In particular, we found that the dynamics of inflation and wage inflation are particularly informative about the correct specification of monetary policy. However, output, the nominal interest rate, and the money growth rates do not contain very discriminating information. In addition, the hump patterns displayed by the impulse responses of most variables is 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 Fed 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 echo Rudebusch's (2002, 2005) findings which suggest the use of extra information in order to reach clear-cut conclusions as to the correct empirical representation of monetary policy rules.

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Figure 1: Timing of Events

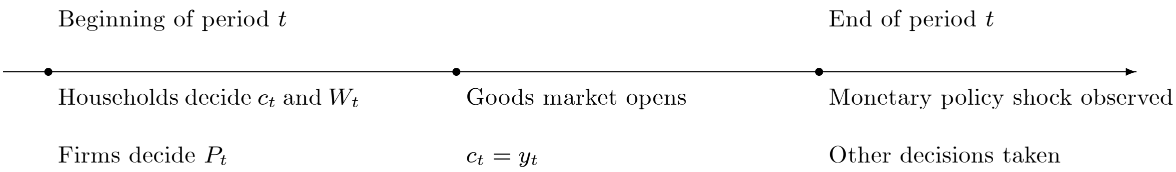

Figure 2: Data Used for Estimation

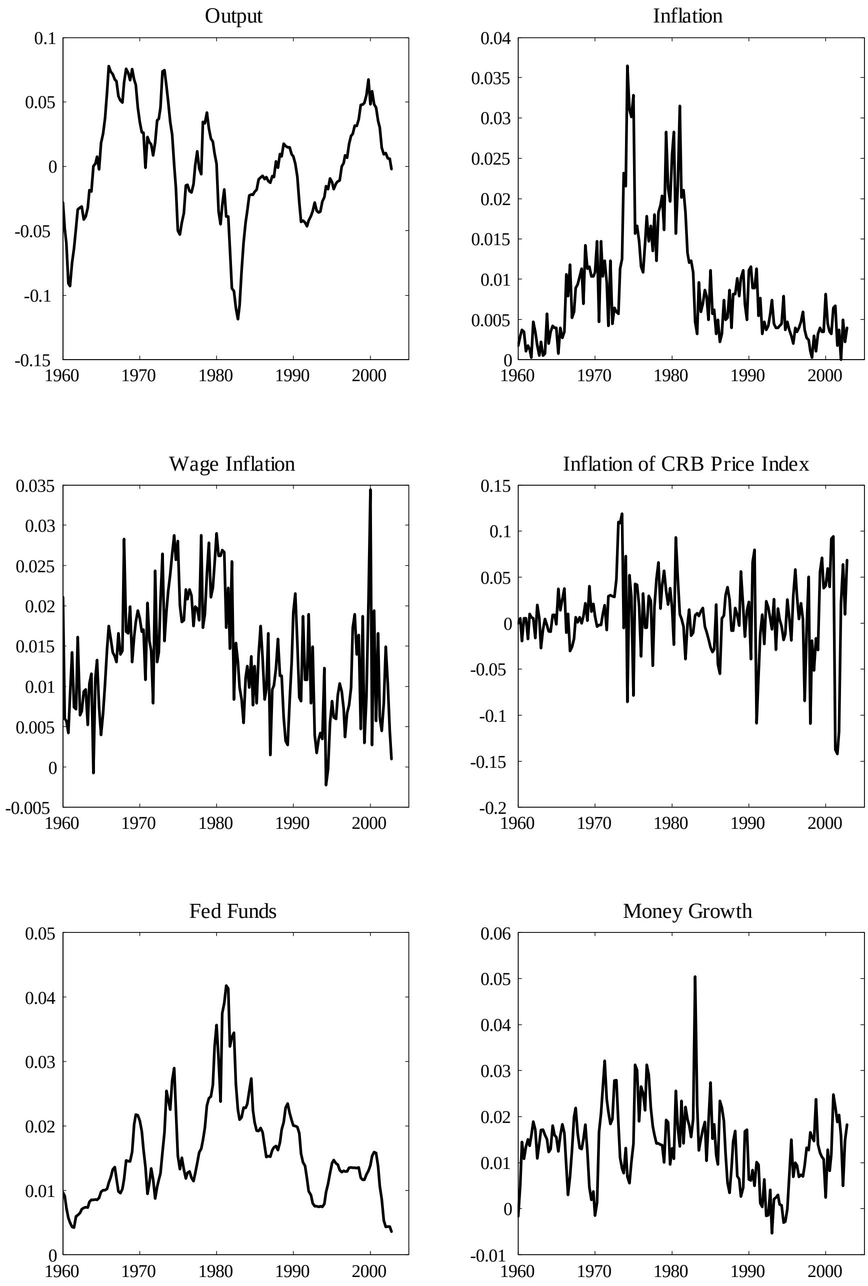

Figure 3: IRF from the Monetary SVAR

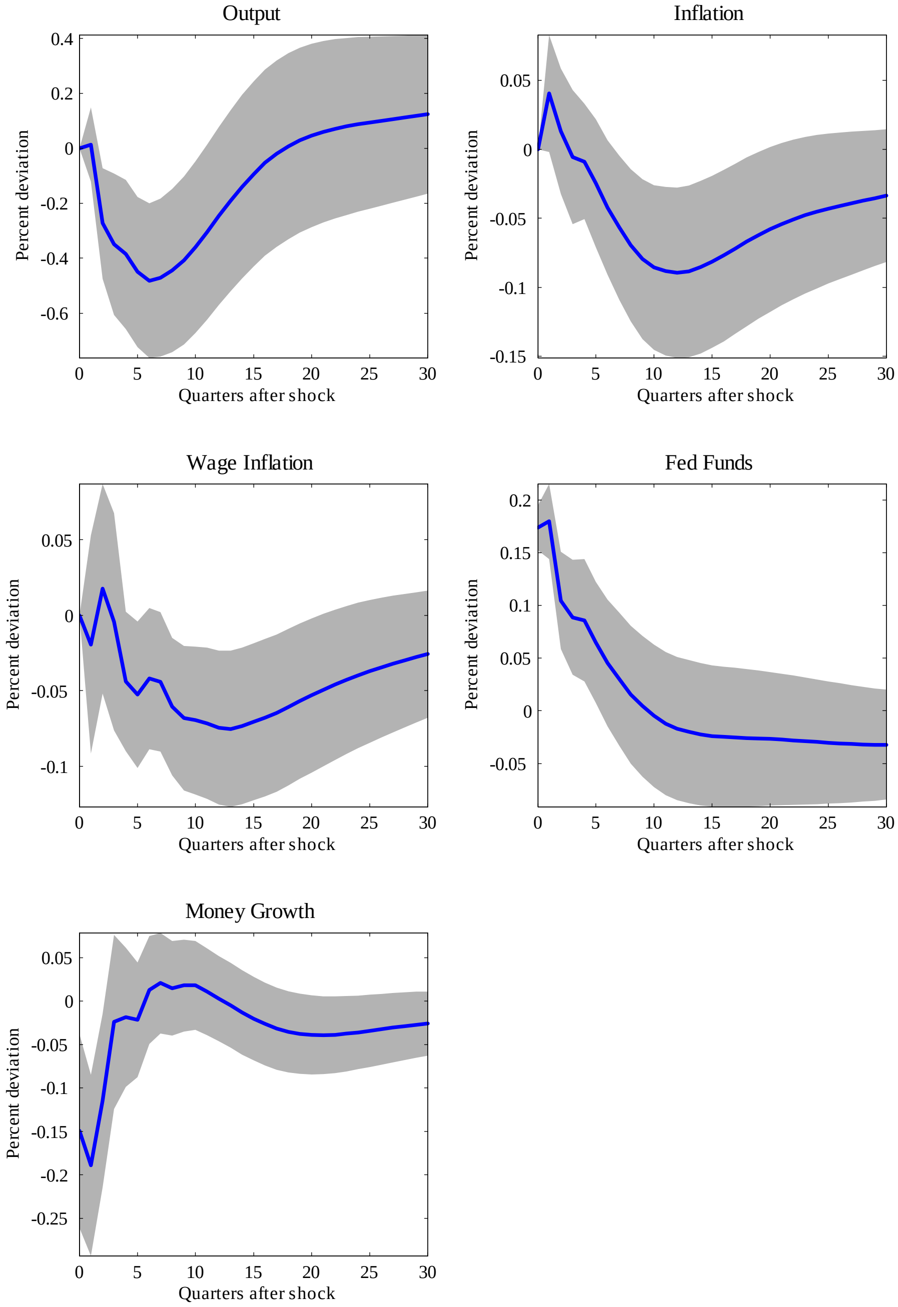

Figure 4: Value of $\mathcal{J}_T$ as a function of ρ_2 and ρ_1.

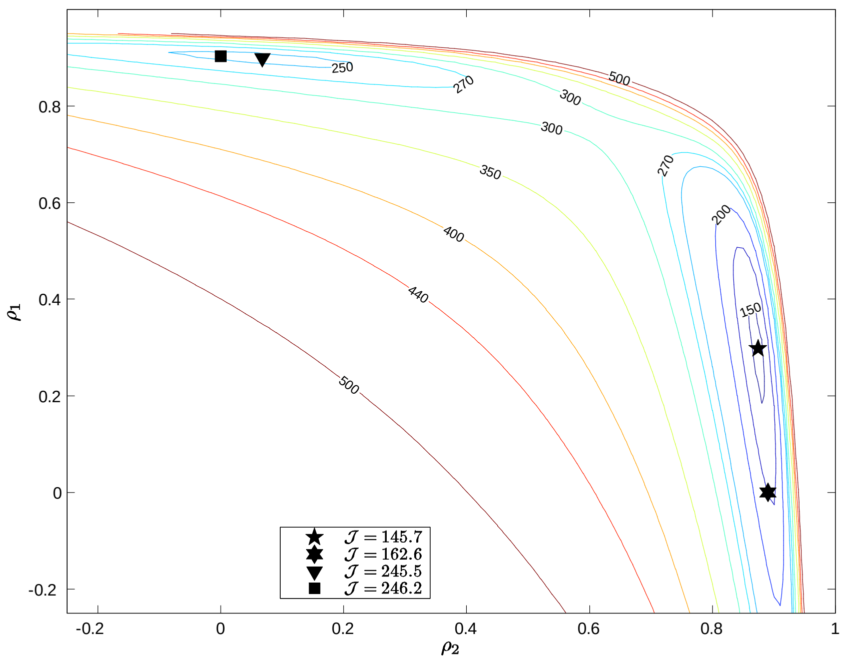

Figure 5: Minimum Distance Estimation Result, *Contingent Dominant Case*

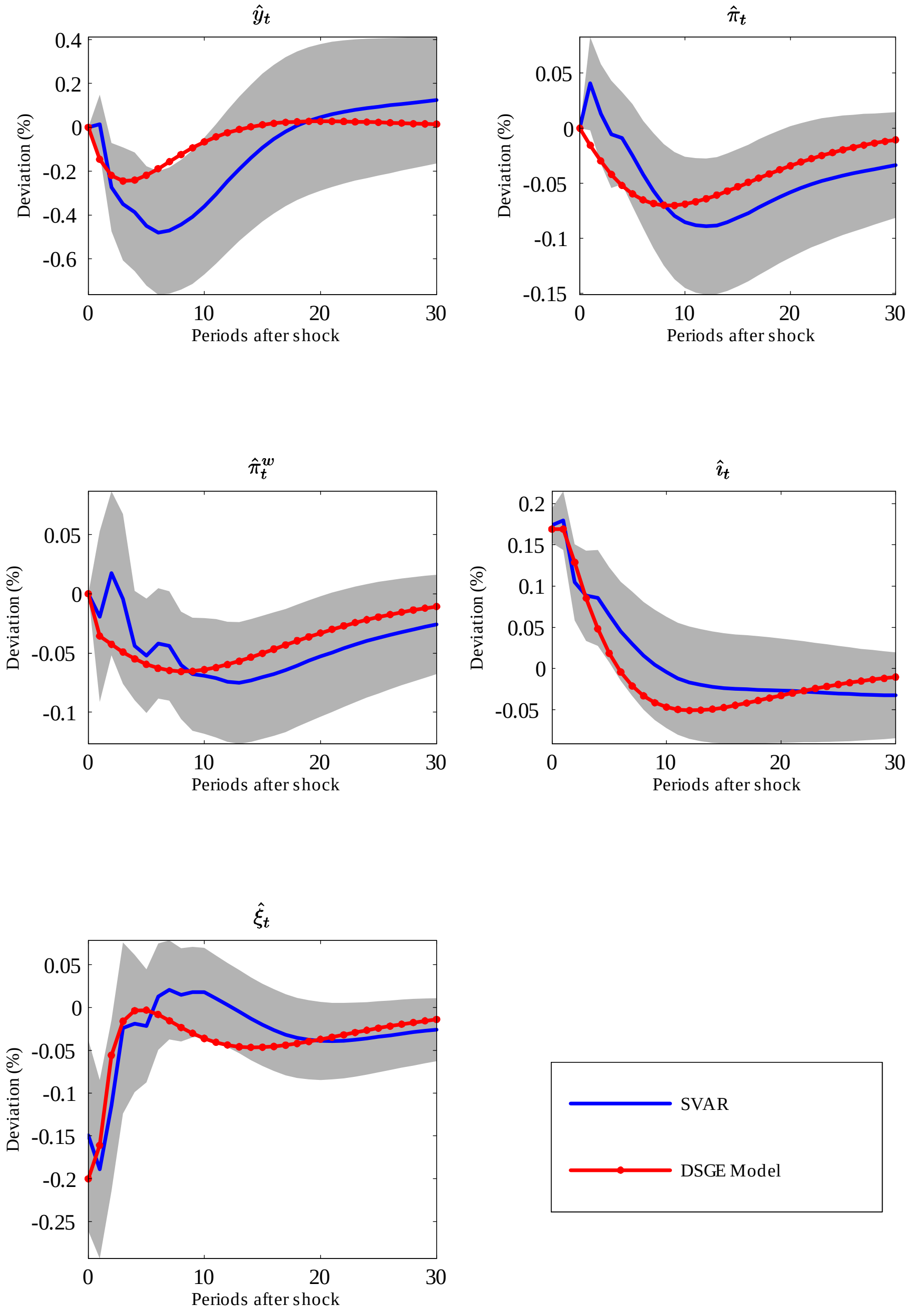

Figure 6: Minimum Distance Estimation Result, *Smoothing Dominant Case*

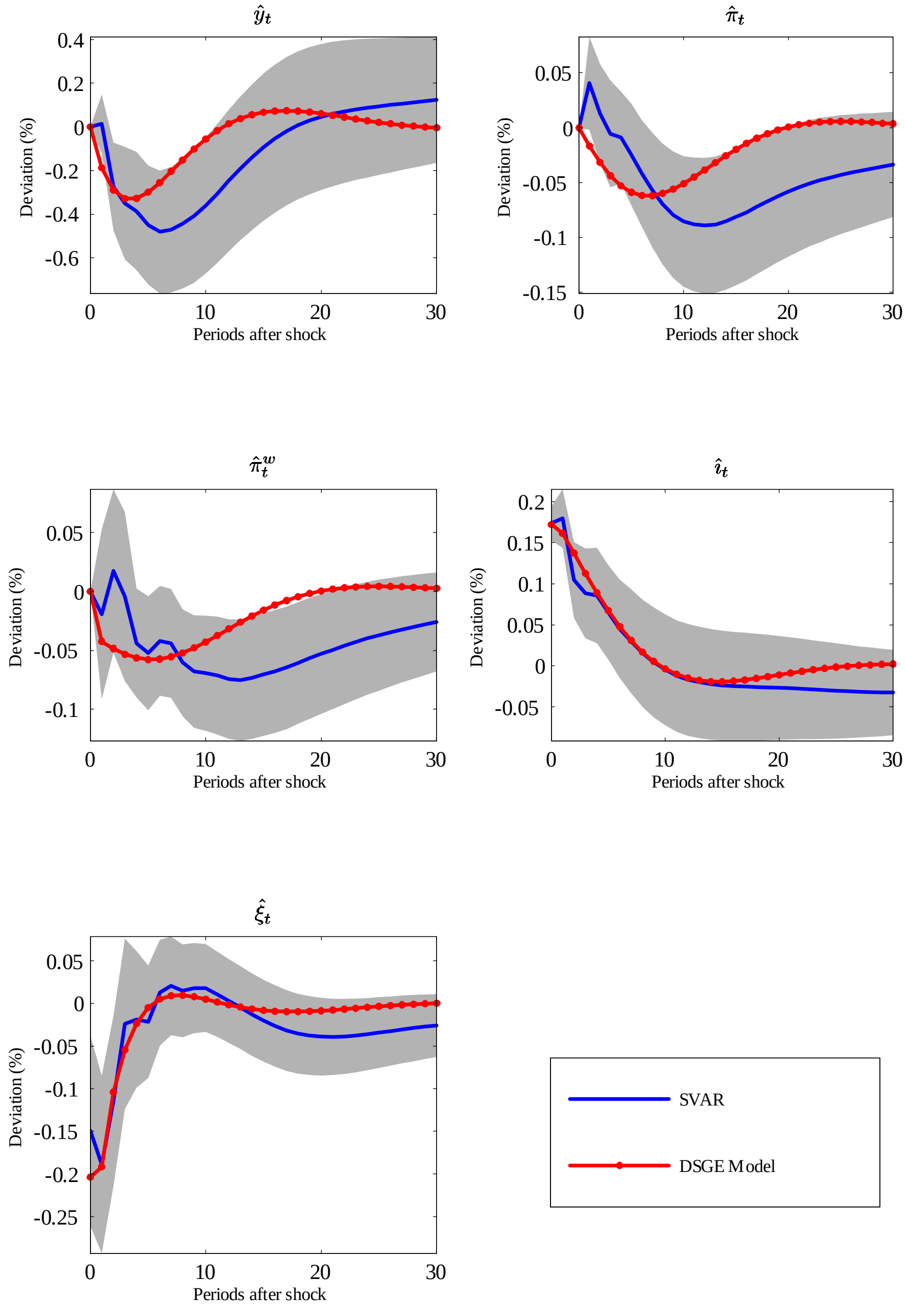

Figure 7: Value of $\mathcal{J}_T$ as a function of ρ_2 and ρ_1.

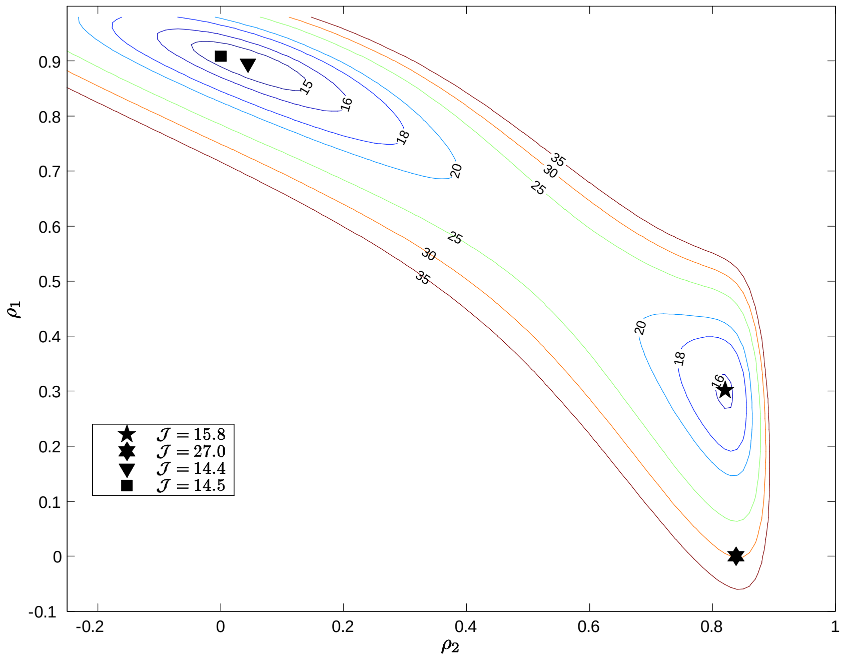

Figure 8: Decomposition of the $\mathcal{J}_T$ statistic as a function of the time horizon

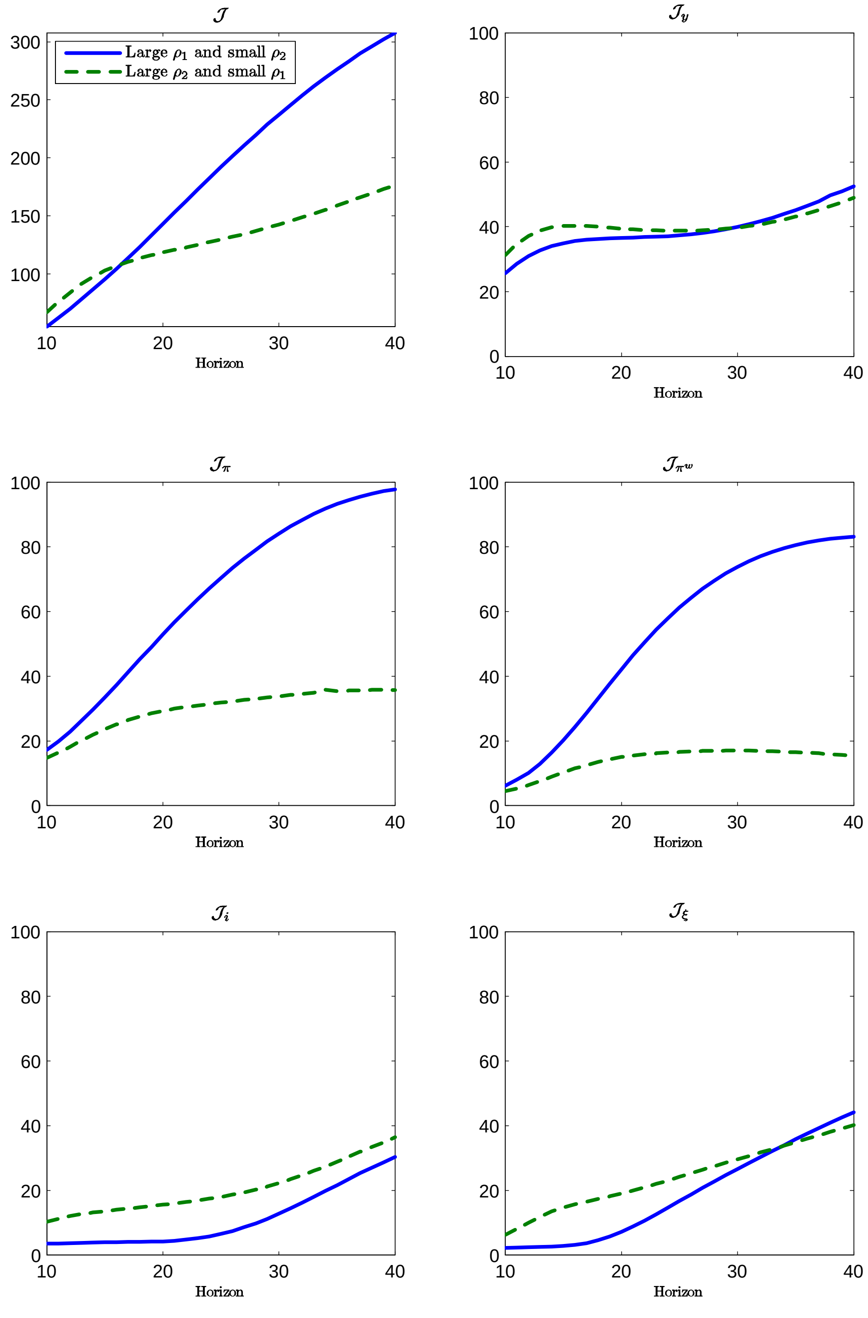

Figure 9: Role of Timing Restrictions – *Contingent Dominant Case*

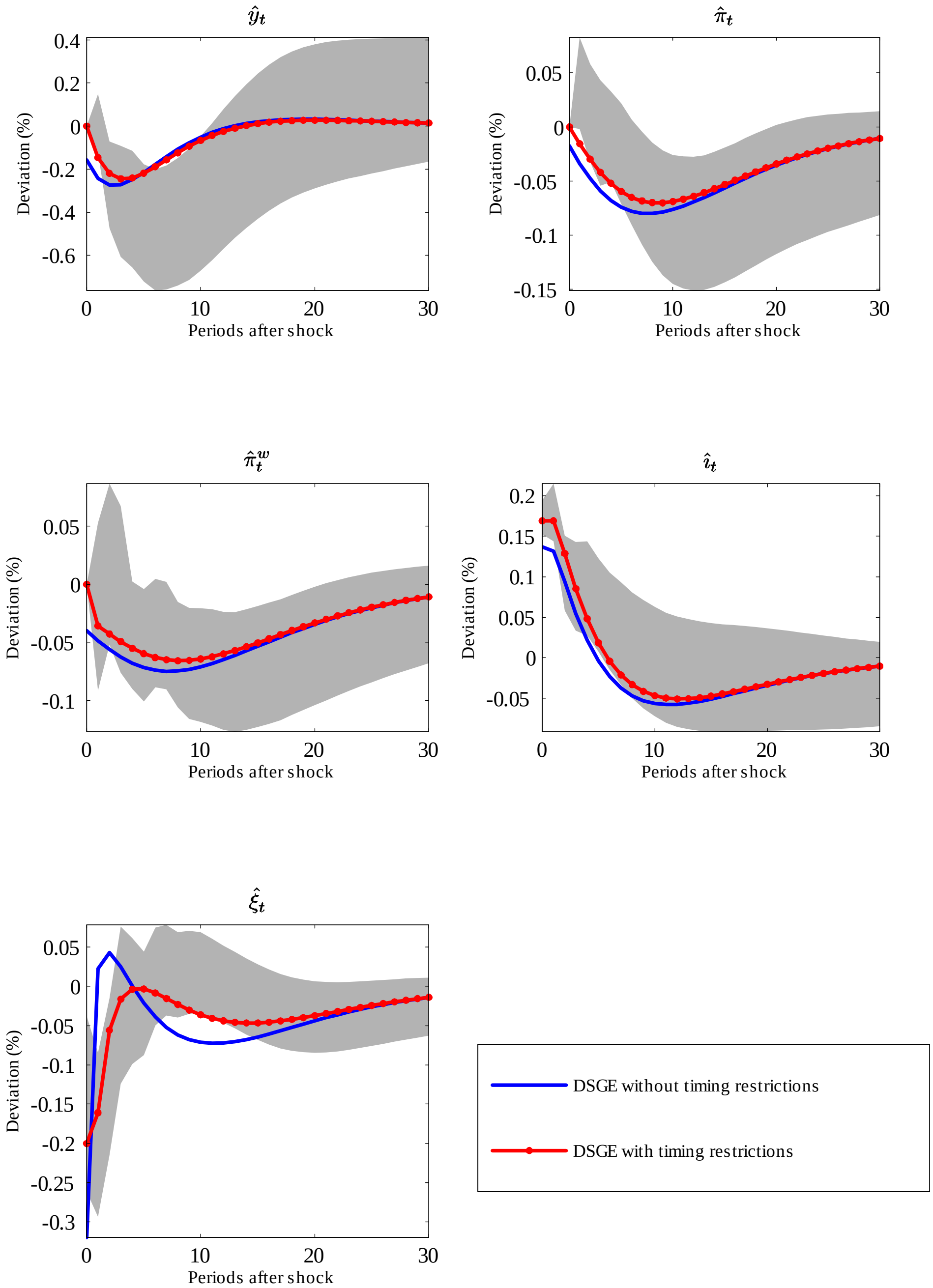

Figure 10: Role of Timing Restrictions – *Smoothing Dominant Case*

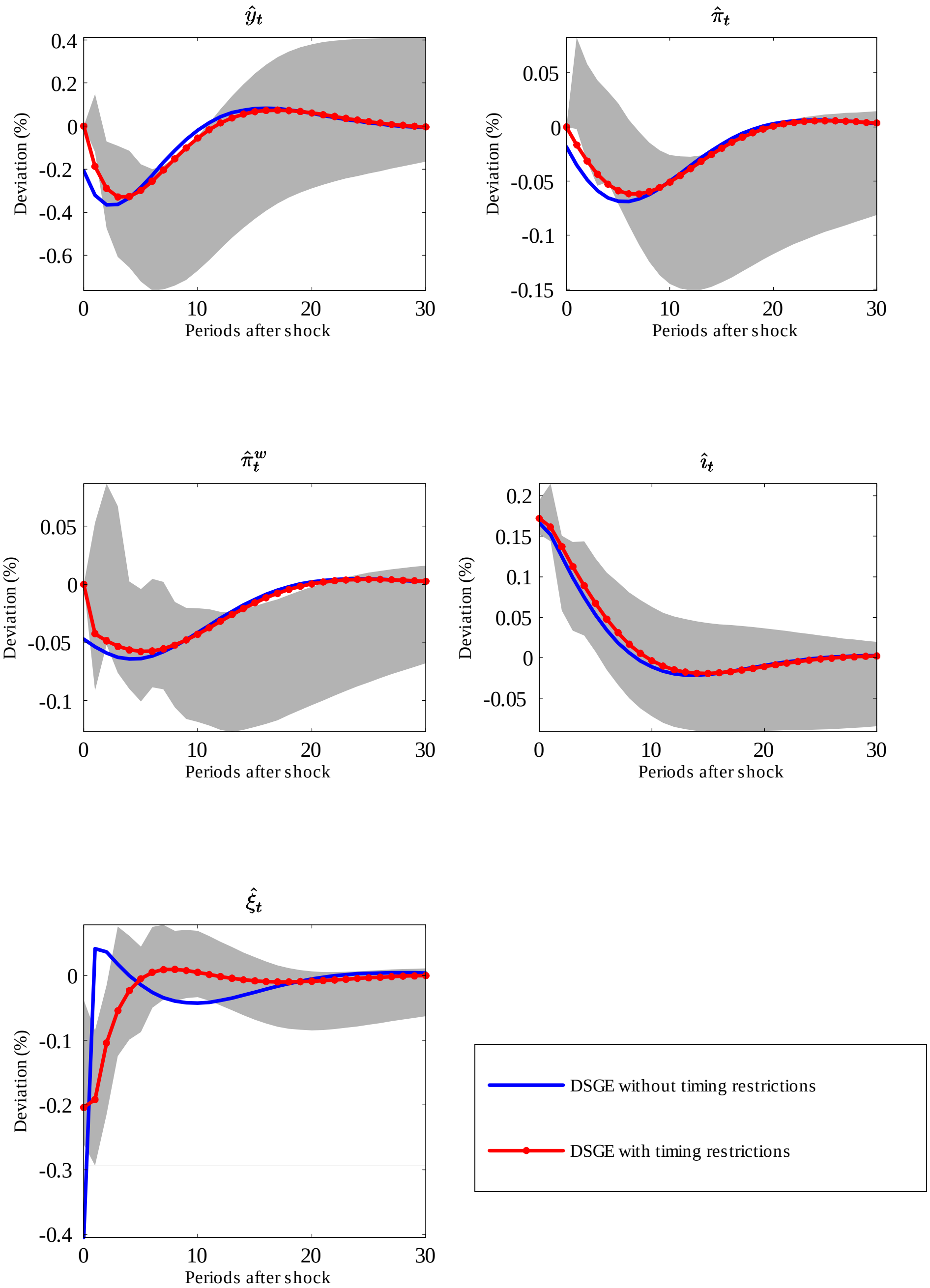

Figure 11: Minimum Distance Estimation Result, Forward-looking Taylor Rule

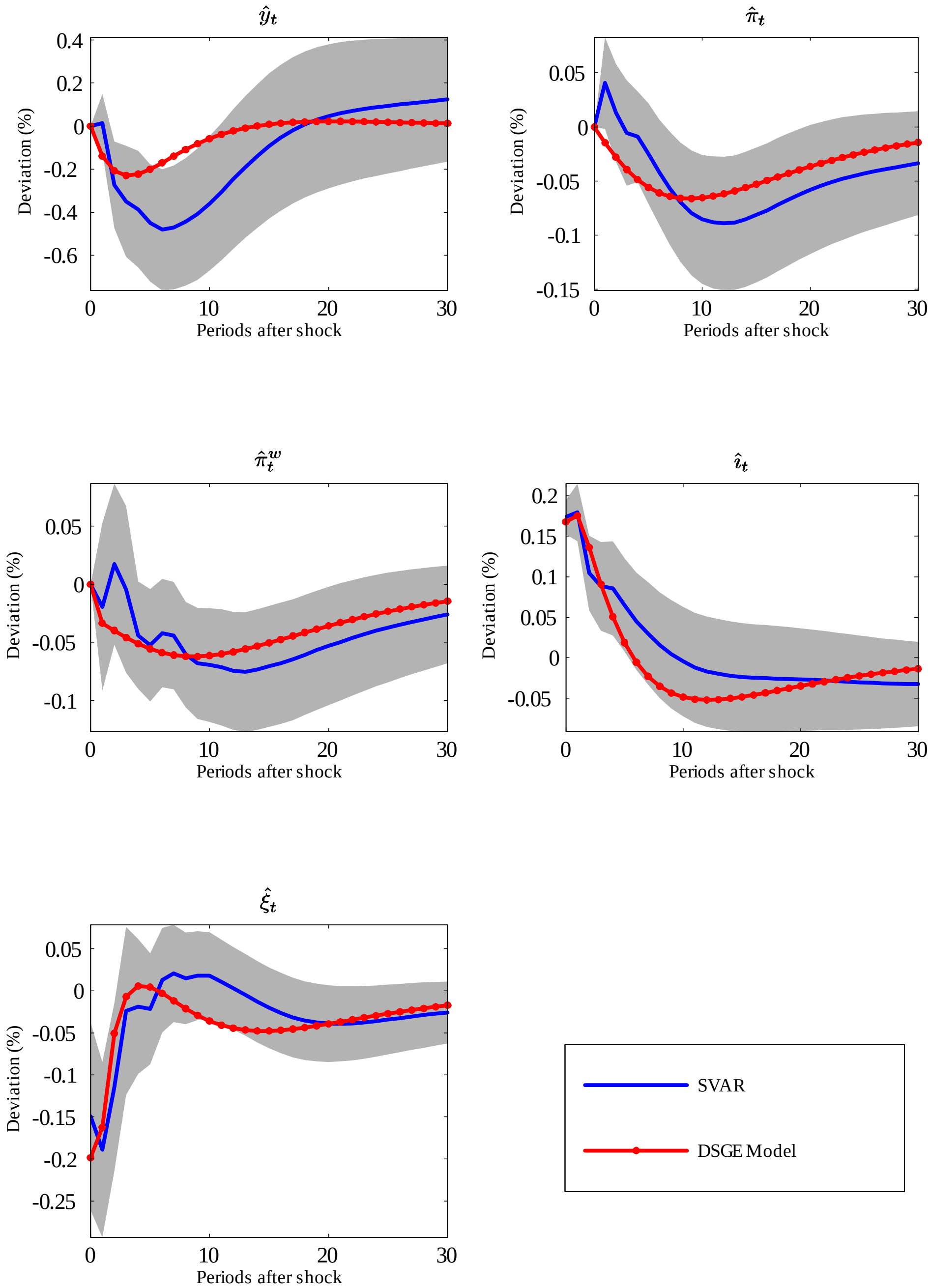

Table 1. Variance Decomposition

Forecast Horizon	0	4	8	20	30
$\hat{y}_t$	0.00 [*]	0.07 [0.02,0.19]	0.16 [0.04,0.32]	0.13 [0.04,0.29]	0.13 [0.05,0.29]
$\hat{\pi}_t$	0.00 [*]	0.01 [0.01,0.05]	0.04 [0.02,0.13]	0.20 [0.05,0.34]	0.22 [0.05,0.36]
$\hat{\pi}_t^w$	0.00 [*]	0.01 [0.00,0.05]	0.04 [0.01,0.10]	0.15 [0.04,0.27]	0.17 [0.04,0.28]
$\hat{\imath}_t$	0.86 [0.73,0.92]	0.40 [0.24,0.51]	0.27 [0.15,0.40]	0.21 [0.12,0.37]	0.22 [0.12,0.39]
$\hat{\xi}_t$	0.08 [0.03,0.15]	0.14 [0.06,0.25]	0.13 [0.06,0.25]	0.13 [0.06,0.25]	0.14 [0.07,0.26]

Notes: Estimated forecast error variance decomposition from the SVAR. The values in brackets are the confidence intervals based on 1000 bootstrap replications of the estimated VAR.

Table 2. 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_x	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	Prob. of no price adj.	0.66
α_w	Prob. of no wage adj.	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

Table 3. 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_1	0.2976 (0.061)	0.8986 (0.009)	0.0000 (∗)	0.9026 (0.007)	0.7535 (0.001)	0.3018 (0.079)	0.8951 (0.049)	0.0000 (∗)	0.9091 (0.035)	0.5606 (0.002)
ρ_2	0.8740 (0.006)	0.0676 (0.092)	0.8900 (0.007)	0.0000 (∗)		0.8207 (0.012)	0.0440 (0.135)	0.8383 (0.016)	0.0000 (∗)	
σ_ν	0.1691 (0.010)	0.1720 (0.010)	0.1882 (0.009)	0.1760 (0.009)	0.1232 (0.005)	0.1731 (0.011)	0.1742 (0.011)	0.1898 (0.010)	0.1754 (0.010)	0.1679 (0.006)
$\mathcal{J}$	145.66 [62.94]	245.54 [0.00]	162.64 [28.17]	246.15 [0.00]	246.52 [0.00]	15.78 [96.86]	14.36 [98.44]	26.97 [57.32]	14.48 [98.86]	21.63 [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_1 small and ρ_2 large; (2): initialization with ρ_2 small and ρ_1 large; (3) constrained case $\rho_1 = 0$; (4) constrained case $\rho_2 = 0$; (5) constrained case $\rho_1 = \rho_2$. In columns (3) and (4), a star denotes a standard error not available.

Table 4. Sensitivity to Calibration Based on $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}_t^w, \hat{\imath}_t, \hat{\xi}_t)'$

		Large ρ_2, Small ρ_1						Large ρ_1, Small ρ_2					
Parameters	Value	$\mathcal{J}$	$\mathcal{J}_y$	$\mathcal{J}_\pi$	$\mathcal{J}_{\pi^w}$	$\mathcal{J}_i$	$\mathcal{J}_\xi$	$\mathcal{J}$	$\mathcal{J}_y$	$\mathcal{J}_\pi$	$\mathcal{J}_{\pi^w}$	$\mathcal{J}_i$	$\mathcal{J}_\xi$
Preferences													
b	0.00	396 [0.00]	73	36	19	238	31	528 [0.00]	97	117	123	121	71
ω_w	10.00	120 [97.00]	35	29	12	19	25	218 [0.04]	35	78	65	14	26
η_i	3.00	181 [5.56]	40	34	17	35	55	275 [0.00]	39	87	76	22	51
Technology													
ϕ	1.00	103 [99.99]	32	22	10	17	21	199 [0.65]	31	68	62	13	24
θ_p	11.00	110 [99.50]	33	25	13	18	22	211 [0.11]	32	72	68	13	25
θ_w	11.00	183 [4.33]	48	41	25	30	39	281 [0.00]	49	94	89	16	32
Price/Wage Setting													
γ_p	0.00	235 [0.00]	26	93	73	16	26	300 [0.00]	28	108	108	18	38
γ_w	0.00	266 [0.00]	30	79	62	36	59	308 [0.00]	29	107	101	23	49
α_p	0.00	356 [0.00]	67	109	71	57	52	#	#	#	#	#	#
α_w	0.00	481 [0.00]	76	49	70	212	73	499 [0.00]	85	118	154	69	74
Target Level													
a_π	3.00	133 [86.45]	46	27	11	23	27	264 [0.00]	42	93	83	17	30
a_y	0.50	147 [60.38]	44	29	13	28	33	216 [0.05]	40	75	61	15	25

Notes: P-value in brackets. A # in the "Large ρ_1, Small ρ_2" panel refers to the corresponding figure in the "Large ρ_2, Small ρ_1" panel.

Table 5. Sensitivity to Calibration Based on $X_t = (\hat{\imath}_t)$

		Large ρ_2, Small ρ_1	Large ρ_1, Small ρ_2
Parameters	Value	$\mathcal{J}_T$	$\mathcal{J}_T$
Preferences			
b	0.00	25 [62.69]	#
ω_w	10.00	13 [99.25]	14 [99.02]
η_i	3.00	16 [97.20]	14 [98.48]
Technology			
ϕ	1.00	11 [99.89]	13 [99.25]
θ_p	11.00	11 [99.78]	13 [99.21]
θ_w	11.00	19 [88.62]	16 [96.47]
Price/Wage Setting			
γ_p	0.00	10 [99.93]	16 [96.47]
γ_w	0.00	22 [76.44]	19 [90.21]
α_p	0.00	20 [87.69]	#
α_w	0.00	24 [70.36]	#
Target Level			
a_π	3.00	15 [98.29]	16 [96.77]
a_y	0.50	18 [92.15]	14 [98.88]

Notes: P-value in brackets. A # in the "Large ρ_1, Small ρ_2" panel refers to the corresponding figure in the "Large ρ_2, Small ρ_1" panel.

Table 6. Estimation Results with Forward-Looking Taylor Rules

Parameter	Based on $X_t = (\hat{y}_t, \hat{\pi}_t, \hat{\pi}^w_t, \hat{\imath}_t, \hat{\xi}_t)'$					Based on $X_t = (\hat{\imath}_t)$				
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**
ρ_1	0.3713 (0.088)	#	0.0000 $(*)$	0.8660 (0.006)	0.6812 (0.046)	0.3683 (0.014)	0.8118 (0.066)	0.0000 $(*)$	0.8743 (0.055)	0.5727 (0.055)
ρ_2	0.7845 (0.007)	#	0.8194 (0.006)	0.0000 $(*)$		0.7220 (0.011)	0.2366 (0.130)	0.7724 (0.009)	0.0000 $(*)$	
σ_ν	0.2310 (0.016)	#	0.2737 (0.012)	0.2060 (0.010)	0.1837 (0.010)	0.2112 (0.019)	0.1904 (0.011)	0.2438 (0.015)	0.2003 (0.010)	0.1871 (0.011)
$\mathcal{J}$	141.46 [71.92]	#	164.53 [24.78]	204.73 [0.33]	166.12 [22.14]	16.59 [95.63]	15.48 [97.29]	30.59 [38.51]	20.49 [87.71]	18.90 [92.37]
$\mathcal{J}_y$	43.13	#	47.38	36.89	40.76	—	—	—	—	—
$\mathcal{J}_\pi$	30.16	#	25.04	64.10	45.35	—	—	—	—	—
$\mathcal{J}_{\pi^w}$	13.63	#	10.34	49.64	27.19	—	—	—	—	—
$\mathcal{J}_i$	24.36	#	40.09	20.85	30.67	116.59	15.48	30.59	20.49	18.90
$\mathcal{J}_\xi$	30.19	#	41.69	33.25	22.15	—	—	—	—	—

Notes: Standard errors in parentheses, P-value in brackets. (1): initialization with ρ_1 small and ρ_2 large; (2): initialization with ρ_2 small and ρ_1 large; (3) constrained case $\rho_1 = 0$; (4) constrained case $\rho_2 = 0$; (5) constrained case $\rho_1 = \rho_2$. A # in column (2) refers to the corresponding figure in column (1). In columns (3) and (4), a star denotes a standard error not available.

Notes d'Études et de Recherche

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2. C. Bordes, « Variabilité de la vitesse et volatilité de la croissance monétaire : le cas français », février 1989.

3. C. Bordes, M. Driscoll and A. Sauviat, "Interpreting the Money-Output Correlation: Money-Real or Real-Real?," May 1989.

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5. B. Bensaid, S. Federbusch et R. Gary-Bobo, « Sur quelques propriétés stratégiques de l'intéressement des salariés dans l'industrie », juin 1989.

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8. M. Boutillier and B. Cabrillac, "Foreign Exchange Markets: Efficiency and Hierarchy," October 1990.

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12. H. Pagès and H. He, "Consumption and Portfolio Decisions with Labor Income and Borrowing Constraints," August 1990.

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