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Testing the structural interpretation of the price puzzle with a cost channel model

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Testing the structural interpretation of the price puzzle with a cost channel model

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Abstract

We estimate a new-Keynesian DSGE model with the cost channel to assess its ability to replicate the price puzzle ie the inflationary impact of a monetary policy shock typically arising in VAR analysis. In order to correctly identify the monetary policy shock, we distinguish between a standard policy rate shifter and a shock to trend inflation ie the time-varying inflation target set by the Fed. While offering some statistical support to the cost channel, our estimated model clearly implies a *negative* inflation reaction to a tightening of monetary policy. We offer a discussion of the possible sources of mismatch between the VAR evidence and our own.

Keywords: cost channel, inflation dynamics, price puzzle, trend inflation.

JEL classification numbers: E30, E52

Selittääkö yritysten rahoituskustannusten vaihtelu rahapolitiikan epätavalliset hintavaikutukset?

Suomen Pankin keskustelualoitteita 20/2009

Efrem Castelnuovo
Rahapolitiikka- ja tutkimusosasto

Tiivistelmä

Tässä tutkimuksessa estimoidaan uuskeynesiläinen rahapolitiikan makromalli, jossa ns. kustannuskanava eli yritysten rahoituskustannusten vaihtelu on keskeinen rahapolitiikan vaikutusten kannalta. Estimoidun mallin kykyä selittää ns. hintapähkinä eli rahapolitiikan kiristymisen inflatorisia vaikutuksia arvioidaan estimointitulosten perusteella. Hintapähkinä on tavanomainen tulos, kun rahapolitiikan yllättävien muutosten, sokkien, vaikutuksia estimoidaan paljon käytetyillä vektoriautoregressiivisillä malleilla. Keskeinen piirre näiden VAR-mallien soveltamisen kannalta rahapolitiikan vaikutusten arviointiin liittyy rahapolitiikan todellisten ei-päätösperäisten, yllättävien muutosten identifiointiin. Korrektin identifioinnin varmistamiseksi rahapolitiikan ohjauskorkoon kohdistuvat sokit erotetaan tässä tutkimuksessa Yhdysvaltain keskuspankin ajan mittaan vaihtelevaan inflaatiotavoitteeseen kohdistuvista sokeista. Vaikka tilastohavainnot tukevat kustannuskanavan olemassaoloa, rahapolitiikan kiristymistä seuraa estimoidussa mallissa inflaatiovauhdin hidastuminen. Estimoitu malli ei siis selitä hintapähkinää ja se on tältä osin ristiriidassa VAR-malleihin perustuvan empiirisen näytön kanssa. Työssä pohditaan mahdollisia syistä tähän ristiriitaan.

Avainsanat: kustannuskanava, inflaatiodynamiikka, hintapähkinä, inflaatiotavoite

JEL-luokittelu: E30, E52

Contents

1 Introduction

What is the short-run reaction of inflation to an unexpected and temporary monetary policy tightening? Macroeconomic textbooks suggest that inflation should react negatively to such a monetary policy move (Woodford, 2003a; Galì, 2008). However, empirical investigations based on the VAR-methodology cast doubts on this prediction.

Figure 1 depicts the impulse response functions produced with a VAR estimated with post-WWII US data.[1] An unexpected one-shot increase in the policy rate leads to i) a significantly positive reaction of the policy rate, ii) a significantly negative reaction of the business cycle, and iii) a significantly *positive* reaction of inflation. This evidence stands in stark contrast with conventional wisdom. Eichenbaum (1992) labeled this short-run positive conditional correlation between policy rate and inflation the 'price puzzle' (the 'VAR evidence' henceforth). Importantly, Castelnuovo and Surico (2009) show that this result is *robust* to the implementation of an alternative identification strategy, based on sign restriction, which does not assume recursiveness, and it is then consistent with the timing of models such as the popular standard new-Keynesian framework.

In fact, a possible interpretation of this VAR empirical regularity is offered by models embedding the 'supply channel', otherwise known as the 'cost channel'. The idea is simple. Cash-constrained firms must borrow money from financial intermediaries to pay the wage-bills to workers before the goods market opens. Consequently, the interest rate paid on borrowings enters firms' marginal costs and influences firms' price setting, so giving a structural role to the presence of the policy rate in the new-Keynesian Phillips curve. This creates an extra-link between monetary policy moves and aggregate inflation fluctuations (Ravenna and Walsh, 2006; Chowdhury, Hoffmann, and Schabert, 2006; Kilponen and Milne, 2007; Surico, 2008; Tillmann, 2009; and Llosa and Tuesta, 2009). Clearly, if the inflationary impact induced by monetary policy moves via the supply channel is stronger than the one operating via the standard 'demand channel', a *positive* reaction of inflation to a monetary policy tightening may very well realize.

The plausibility of such a structural interpretation, however, is ultimately an *empirical* issue. This paper employes Bayesian techniques to estimate a new-Keynesian small-scale DSGE model embedding the cost channel. The model is an extension of the baseline, aggregate-demand based set up widely employed to scrutinize US inflation (Benati and Surico, 2008a; Benati and Surico, 2008b; Benati, 2008a; Canova, 2009). The model has the potential to replicate the VAR evidence, because the supply channel may in *principle* prevail. Our exercise aims exactly to understand if, *empirically,* the strength of the cost channel is actually such to induce a positive correlation between inflation and the policy rate conditional on a monetary policy shock.

[1] Impulse responses related to a trivariate VAR with GDP deflator inflation, CBO output gap, and federal funds rate, sample: 1954:III–2008:II. Similar evidence, originally reported by Sims (1992), is also put forward by Stock and Watson (2001) and Rabanal (2007). Hanson (2004) shows that this result is robust to the introduction of commodity prices as well as a variety of other inflation predictors in the VAR.

Our main result is twofold. First, we do find some statistical support for the cost channel. Due to the presence of the short-term interest rate in the inflation equation, the cost channel model fits the data better with respect to the standard 'demand channel only' new-Keynesian framework. Second, we do *reject* the structural interpretation of the price puzzle. Clearly, the data prefer a parameterization of the model for which the demand channel is relatively stronger than the cost channel in transmitting the monetary policy impulses to inflation. The estimated degree of interest rate smoothing is the ingredient boosting the demand side's relative strength. Possibly, this is so because inflation expectations are strongly influenced by a *gradual* monetary policy conduct (Woodford, 1999; Woodford, 2003b). Then, an increase in marginal costs driven by drifts in the policy rate is less inflationary than it would be under a less persistent policy conduct. Our results are robust to a variety of perturbations of the baseline analysis – sample selection, different measure of inflation, different models for the policy shock. We then conclude that the VAR price puzzle is not a fact, but instead an *artifact* possibly due to model misspecification.

Before moving to the next Section, we make contacts with some strictly related contributions. Barth and Ramey (2001) analyze different US sectors and find that sectorial differences in the working capital may rationalize the heterogeneous impact across sectors of a monetary policy shock. Similar results are obtained by Gaiotti and Secchi (2006) for Italy, and Dedola and Lippi (2005) for France, Germany, Italy, and the United Kingdom. In a single-equation framework, Ravenna and Walsh (2006) support the presence of the cost-channel for the US economy, Tillmann (2008) for the US, UK, and Euro Area, and Chowdhury, Hoffmann, and Schabert (2006) for Canada, France, Italy, the UK and the United States. There is then support for the empirical relevance of the cost-channel for a variety of countries, the US being among them.

As regards the structural interpretation of the VAR evidence, Chowdhury, Hoffmann, and Schabert (2006) couple an estimated Phillips curve embedding the cost channel with a *calibrated* demand side, and show that such model replicate the VAR evidence as for Italy, the UK, and the United States. Christiano, Eichenbaum, and Evans (2005) estimate a model featuring several nominal and real rigidities by *indirect inference* (impulse response matching), and also replicate such a fact. With the same econometric strategy, Henzel, Hulsewig, Mayer, and Wollmershaeuser (2007) obtain similar results for the Euro Area. While offering stimulating results, this evidence is not conclusive. Calibrated models may lead to dynamics that are at odds with the data. Moreover, estimation techniques such as impulse-response matching are bias-prone, and may lead to fragile conclusions when confronted to alternatives such as Bayesian estimation (see Canova and Sala, 2009).

The paper closest to ours is probably Rabanal (2007). He investigates the sign and magnitude of the inflation reaction to a monetary policy shock by estimating a medium-scale model a la Christiano, Eichenbaum, and Evans (2005) with Bayesian techniques, and finds evidence supporting the 'textbook' monetary policy transmission mechanism. In his paper, the key-drivers for this result are i) a less than full wage indexation, ii) a moderate wage stickiness, and iii) a high price stickiness. Our paper differs from Rabanal's (2007) along several dimension. First, we jointly model the standard monetary policy shock *and* the shock to 'trend inflation', ie the time-varying inflation target set by the Fed. This is a key-modeling choice. In fact, if the Fed had actually pursued a time-varying inflation target, in assuming a constant target we would force the dynamics of the inflation target to enter the 'residual' of the policy rule, and we would label as 'policy shock' what, *de facto*, is a convolution of the true policy innovation and the inflation target dynamics. Bache and Leitemo (2008) show that this misspecification can dramatically bias impulse responses to a monetary policy shock in autoregressive models. Given the empirical evidence pointing towards trend inflation in the US (Ireland, 2007; Cogley and Sbordone, 2008; Cogley, Primiceri, and Sargent, 2009, we believe that the separate identification of these two monetary policy shocks is important for the issue at stake). Second, we relax the unitary upper bound to the cost channel parameter imposed by Rabanal (2007) when conducting his estimates. While such an upper bound represents a natural imposition when interpreting the cost channel parameter as the share of financially constrained firms in the economy, frictions on the financial markets may indeed suggest a pass-through from the policy rate to the lending rate larger than one (Chowdhury, Hoffmann, and Schabert, 2006; Tillmann, 2008). Moreover, model misspecification – eg a too simplistic banking sector – may easily turn the structural cost channel parameter to a reduced form capturing the direct effect of the policy rate to inflation. At that point, however, the imposition of a unitary upper bound would not necessarily be warranted. For these reasons, we allow the cost channel parameter to take values above one, so possibly putting the model in a more favorable condition to replicate the VAR fact. All in all, our contribution should be seen as complementary with respect to Rabanal's (2007).

The paper develops as follows. Section 2 presents the new-Keynesian model with the cost channel we work with, and scrutinizes the role that the model's parameters plays in shaping the inflation reaction to a monetary policy shock. Section 3 presents and discusses our empirical findings. In Section 4 we show that our baseline results are robust to a variety of perturbations – sample selection, different inflation measures, different structure of the policy shock. A discussion on the reasons underlying the possible misspecification of the policy shock in VAR analysis is then offered in Section 5. Section 6 concludes.

2 A model with the cost channel

As discussed above, we concentrate on a new-Keynesian DSGE model flexible enough to replicate the VAR evidence.

2.1 Structure of the model

The model reads as follows:[2]

$$
\begin{aligned}
\pi_t &= \frac{\beta}{1+\alpha\beta} E_t \pi_{t+1} + \frac{\alpha}{1+\alpha\beta}\pi_{t-1} + \kappa[(\sigma+\eta)x_t + \psi R_t] + \varepsilon_t^{\pi}, && (2.1)\\
x_t &= \frac{1}{(1+h)} E_t x_{t+1} + \frac{h}{(1+h)} x_{t-1} - \frac{(1-h)}{\sigma(1+h)}(R_t - E_t\pi_{t+1}) + v_t^x && (2.2)\\
R_t &= \phi_R R_{t-1} + (1-\phi_R)[\phi_\pi(\pi_t - \pi_t^*) + \phi_x x_t] + \varepsilon_t^R, && (2.3)\\
\pi_t^* &= \rho_* \pi_{t-1}^* + \varepsilon_t^*, && (2.4)\\
v_t^x &= \rho_x v_{t-1}^x + \varepsilon_t^x, && (2.5)\\
\varepsilon_t^j &\sim i.i.d.N(0, \sigma_j^2), j \in \{\pi, x, R, *\}. && (2.6)
\end{aligned}
$$

Eq. (2.1) is an expectational new-Keynesian Phillips curve (NKPC) in which π_t stands for the inflation rate, β identifies the discount factor, α indicates price indexation to past-inflation, x_t identifies the business cycle – the 'output gap' – whose impact on current inflation is influenced by the slope-parameter κ (a convolution of the discount factor and the probability of non-reoptimizing prices by firms), σ is the representative consumers' degree of relative risk aversion, η is the inverse of the labor supply elasticity, and ε_t^{π} is interpreted as 'inflation' shock, or 'supply' shifter. Differently with respect to the 'demand channel only' models, eq. (2.1) embeds a *direct* impact of the nominal interest rate R_t on the inflation rate, which is active as long as the 'cost channel' parameter $\psi > 0$.

Eq. (2.2) is obtained by log-linearizing the consumption Euler equation stemming from the household's intertemporal problem. Output fluctuations are driven both by expectations on future realizations of the business cycle and by the ex-ante real interest rate, whose loading (the intertemporal elasticity of substitution, ie IES) is a convolution of habits and relative risk aversion. The demand shock v_t^x, which is autoregressive as suggested by eq. (2.5), is interpreted as households' preference shock or a fiscal shock.[3] Eq. (2.3) is a Taylor rule postulating the systematic reaction of the policy rate to movements

[2] The variables in the model are expressed in log-deviations with respect to their non-stochastic steady state values or, as for ouptut, in deviations with respect to its long-run trend.

[3] In a former version of the paper we modeled also the inflationary shock as an autoregressive process of order one. Possibly, this is due to the presence in the NKPC of the policy rate, which inherits (part of the) persistence of the trend inflation process. We eventually decided against it because we verified that i) the estimated persistence of such shock in this application is very low, and ii) its contribution to the fit of the model is negligible.

in the inflation *gap* and the output gap. Past policy decisions matter, and their impact is captured by the interest rate smoothing parameter ϕ_R, as in Clarida, Galì, and Gertler (2000). The zero-mean *i.i.d.* random shock ε_t^R stands for the monetary policy innovation. The evolution of the inflation target – formalized by eq. (2.4) – is dictated by the autoregressive parameter ρ_* as well as the volatility σ_* of its innovation ε_t^*. This process is typically assumed to be a random walk or a very-persistent variance-stationary process capturing the low-frequency component of the inflation rate, which are likely to be sensible approximations of the time-varying target set by monetary-policy authorities. The innovation processes (2.6) close the model.

A set up similar to the one hereby presented (time-varying inflation target aside) has recently been object of *theoretical* investigations by Ravenna and Walsh (2006), Chowdhury, Hoffmann, and Schabert (2006), Kilponen and Milne (2007), Surico (2008), Tillmann (2009), and Llosa and Tuesta (2009).

2.2 Investigating the mechanism

Before moving to estimation, it is of interest to perform some simulations in order to scrutinize the ability of the model to match the VAR evidence, ie to replicate the price puzzle. Given the timing of the model, very standard in the new-Keynesian arena, inflation may immediately (contemporaneously) jump up in reaction to a policy tightening. This *upward* jump, which is at odds with conventional wisdom, would i) prove the model to be able to replicate the VAR fact, ii) corroborate the structural interpretation of such fact, and iii) points towards the cost channel as the ingredient possibly needed to replicate it.

We perform the following exercise. First, we calibrate the model with sensible parameter values – we employ $\beta = 0.99, \psi = 1.75, \kappa = 0.05, \alpha = 0.5, \sigma = 3, \eta = 1, h = 0.7, \phi_\pi = 1.7, \phi_x = 0.3, \phi_R = 0.3$.[4] We then explore the robustness of the conclusions coming from the baseline calibration by perturbing the parameters one-by-one. To have a sense of the different impact that those perturbations may have on the inflation dynamics, we repeat the same exercise also for the 'demand channel only' model – ie under $\psi = 0$.[5] For comparability reasons, we normalize the monetary policy shock so to induce an on-impact 25 basis points jump of the policy rate.

Figure 2 displays the model consistent *on-impact* inflation reactions. First of all, the model is clearly able to produce a *positive* inflation reaction to an unexpected interest rate hike, which is indeed the outcome of the baseline calibration (identified by vertical black dotted lines). Second, the

[4]To simulate the effects of a monetary policy shock assumed to be uncorrelated to all other shocks, we set all other shocks' standard deviations to zero. Consequently, we need *not* specify the values for the roots of the AR(1) exogenous processes of our model.

[5]We solve the model numerically by searching for the unique equilibrium under rational expectations. For a discussion of the uniqueness conditions in the model (1)–(3) with no endogenous persistence of any sort and with a central bank just reacting to inflation fluctuations, see Brueckner and Schabert (2003). Regarding this issue, Surico (2008) investigates the role played by the systematic reaction to output gap fluctuations. Llosa and Tuesta (2009) study the effects of the cost-channel on determinacy and learnability of the rational expectations equilibrium.

qualitative impact of perturbations to the model parameters is mostly in line with expectations. Obviously, a higher cost channel renders more likely to get the inflationary effect out of a policy tightening. The same holds true as for habit formation, which decreases IES. The (inverse of the) labor elasticity η works instead in favor of the demand channel: The higher η, the lower the labor supply elasticity to real wage fluctuations (ie the higher the slope of the labor supply), the higher the real wage and marginal cost drop after a policy shock, the more intense the negative push on inflation exerted by the demand channel. The slope of the NKPC and the relative risk aversion parameter's impacts are a priori unclear. On the one hand, they enhance the deflationary demand pressure due to an economic bust triggered by a conservative monetary policy shock. One the other, the relative risk aversion has a negative impact on the IES, and both κ and σ enhance also the cost channel effect (see the NKPC (2.1)). It turns out that, under the baseline calibration, both parameters increase the likelihood of the inflationary effect. Interestingly, the weakening of the demand channel by σ is present also when the 'demand channel only' framework is considered (green dotted line). Obviously, the opposite holds as for the slope of the NKPC, which, absent the supply channel, clearly magnifies the deflationary impact of a policy tightening. Price indexation monotonically increases the probability of a positive policy rate-inflation conditional correlation, possibly because of the persistence induced in the inflation process that is translated into higher inflation expectations. The Taylor rule parameters on the inflation gap also strengthen the policy inflationary effect: Given that the baseline calibration suggests 'inflation up', the higher the policy rate reaction to inflation, the stronger the supply channel. The reaction to the output gap seems to be hardly influent in this simulation. By contrast, and not surprisingly, the interest rate smoothing parameter clearly boosts the demand side's relative strength. This is so because inflation expectations are strongly influenced by the monetary policy conduct given the latter's persistence (Woodford, 1999; Woodford, 2003b), then an increase in marginal costs driven by borrowing costs is less inflationary than it would be under a less gradual policy conduct. Finally, when switching off the cost-channel, we are obviously left with negative inflation realizations no matter what the calibration employed is.

Wrapping up, different *calibrations* may lead to different on-impact reactions. We then move to the *estimation* of the structural model to let the data free to speak as regards magnitude and sign of the inflation reaction to a monetary policy shock.

3 Bayesian estimation

We estimate the model (2.1)–(2.6) with Bayesian methods (for an extensive survey, see An and Schorfheide, 2007) for the sample 1954:III–2008:II. We limit our study to the second quarter of 2008 so to avoid dealing with the acceleration of the financial crises began with the bankruptcy of Lehman Brothers in September 2008, which triggered non-standard policy moves by the Fed

(Brunnermeier, 2009). Importantly, the use of a full system approach is likely to limit the weak instruments problem affecting GMM when applied to hybrid schedules displaying rational expectations among the drivers of the modeled variables (Mavroeidis, 2004 and Canova and Sala, 2009). Moreover, the full system estimation enables us to account for cross equation restrictions clearly affecting the estimation of NKPC's parameters (for a maximum likelihood application to the Euro-area NKPC, see Fanelli, 2008). In the model we focus on, a notable example regards the time-varying inflation target, which enters (also) the solution of the inflation rate and shapes its persistence, so clearly affecting the estimate of the cost channel parameter as well as others (eg price indexation).

To estimate the model, we employ three observables. The output gap is computed as percentualized log-deviation of the real GDP with respect to the potential output as computed by the Congressional Budget Office.[6] The inflation rate is the quarterly growth rate of the GDP deflator. Finally, for the short-term nominal interest rate we consider the effective federal funds rate (averages of monthly values) expressed in quarterly terms. The source of the data is the Federal Reserve Bank of St. Louis (FREDII). All the transformed data are demeaned before estimation.

3.1 Priors

Our Bayesian estimation calls for the imposition of prior densities on the model parameters. First and foremost, we have to set a prior for the cost channel parameter ψ. Exploring US data with Bayesian techniques, Rabanal (2007) estimates it to be 0.15 in a full sample analysis, and 0.56 for the 1980s and 1990s. Ravenna and Walsh (2006) appeal to single-equation GMM estimation and find it to be 1.276 (benchmark estimate of the battery they provide), a value very close to that put forward by Chowdhury et al (2006), who propose 1.3 on the basis of GMM estimation. Christiano, Eichenbaum, and Evans (2005) set it to 1. In order not to play against the cost channel interpretation of the inflation reaction found with VARs, we assume ψ to be Normally distributed with mean 1.75 (that is the value we employed in investigating the mechanism, see Section 2.2). However, to remain relatively agnostic on this parameter, we allow for a fairly large standard deviation, ie 0.7. We also impose dogmatic priors as for the inverse of the labor supply elasticity η and the slope of the NKPC. As for the former parameter, since we do not employ labor data in the estimation, and given the identification issues regarding it, we

[6] Ravenna and Walsh (2006) point out that under an active cost-channel, the welfare relevant output gap is the one computed by considering flexible output conditional on a constant and positive nominal interest rate. Interestingly, Justiniano and Primiceri (2008a) show that the correlation between the theoretical output gap – computed with a medium-scale DSGE model – and empirical measures of the output gap like the one employed in this paper is high. Moreover, the model we focus on does not display capital, then the model-consistent natural level of output could very well be misspecified. Finally, we ran some exercised in which we switched off the reaction of the Fed to business cycle fluctuations. The results presented in this paper are robust to this perturbation of the model.

calibrated it to 1, a standard value in the literature. Preliminary attempts to estimate the slope of the NKPC led to implausibly low realizations, a problem encountered by eg Ireland (2004). We then set κ to 0.05, a value in line with recent empirical evidence (Benati and Surico, 2008a and 2008b). As for the trend inflation process, we follow Cogley, Primiceri, and Sargent (2009) and set the autoregressive parameter ρ_* to 0.995 so to force the trend inflation process to capture low-frequency movements in inflation. Following the convention, we also fix the discount factor $\beta = 0.99$ (corresponding to an annual discount rate of approximately 4%). Given the relevance of the interest rate smoothing degree ϕ_R – shown in Section 2.2 – we assume a beta distribution with a conservative prior mean – 0.5 – and a large standard deviation – 0.285. The remaining priors are standard, and in line with Benati and Surico (2008a) and Benati and Surico (2008b) and Cogley, Primiceri, and Sargent (2009) as for the parameters in common. Table 1 collects our prior densities.

3.2 Posterior densities and Bayesian impulse responses

Given the vector $\xi = (\psi, \beta, \alpha, \kappa, \eta, \sigma, h, \phi_\pi, \phi_x, \phi_R, \rho_x, \rho_*, \sigma_\pi, \sigma_x, \sigma_R, \sigma_*)'$ of structural parameters, the vector of endogenous variables $z_t = [x_t, \pi_t, R_t]'$, the autoregressive demand shock $\varepsilon_t = [v_t^x]'$, the vector of innovations $\eta_t = [\varepsilon_t^x, \varepsilon_t^\pi, \varepsilon_t^R, \varepsilon_t^*]'$, and the vector of observable variables we aim at tracking $Y_t = [x_t, \pi_t, R_t]'$, we write the model in state space form, we relate the latent processes to the observable variables via the measurement equation (without assuming any measurement errors),[7] and we employ the Kalman filter to evaluate the likelihood $L(\{Y_t\}_{t=1}^T \mid \xi)$. The posterior distribution $p(\xi \mid \{Y_t\}_{t=1}^T)$ is then proportional to the product of the likelihood function $L(\{Y_t\}_{t=1}^T \mid \xi)$ and the priors $\Pi(\xi)$.[8]

[7] Estimations obtained by adding white noise measurement errors – not shown for the sake of brevity, but available upon request – confirmed the robustness of our findings.

[8] To perform our Bayesian estimation we employed Dynare 4.0, available at http://www.cepremap.cnrs.fr/dynare/. The model is estimated by implementing a two-step strategy. First, we estimate the mode of the posterior distribution by maximizing the log-posterior density, which combines our priors on the parameters of interest with the likelihood function. Second, we employ the random-walk Metropolis-Hastings algorithm to estimate the posterior distribution. The mode of each parameter's posterior distribution was computed by using the 'csminwel' algorithm elaborated by Chris Sims. A check of the posterior mode, performed by plotting the posterior density for values around the computed mode for each estimated parameter in turn, confirmed the goodness of our optimizations. We then exploited such modes for initializing the random walk Metropolis-Hastings algorithm to simulate the posterior distributions. In particular, the inverse of the Hessian of the posterior distribution evaluated at the posterior mode was used to define the variance-covariance matrix of the chain. The initial VCV matrix of the forecast errors in the Kalman filter is set to be equal to the unconditional variance of the state variables. We initialized the state vector in the Kalman filter with steady-state values. We simulated two chains of 500,000 draws each, and discarded the first 50% as burn-in. To scale the variance-covariance matrix of the random walk chain we used factors implying an acceptance rate belonging to the [23%, 40%] interval. We verified the convergence towards the target posterior distribution via the Brooks and Gelman (1998) convergence checks. As typically done in the literature, we discarded all the draws not implying a unique equilibrium of the system.

Our posterior estimates are reported in Table 1. First, we focus on the cost channel parameter ψ. Its posterior mean reads 1.18, clearly smaller than the prior mean – the latter being 1.75. Still, it is 'significant', ie the 5th percentile of its posterior density reads 0.60, a value clearly larger than zero. Notably, this result does not appear to be driven by our prior choice. When re-estimating the model either with the prior $\psi \sim N(1, 0.7)$ – so with a marked downward shift of the mean – or with $\psi \sim N(1, 1.3)$ – with a much larger variance – we obtained densities comparable to those displayed in Table 1. Interestingly, when shutting down the cost channel and re-estimating the model, the marginal likelihood deteriorates of about 2 log-points,[9] ie the Bayes factor amounts to $exp(2) \approx 7.4$. This deterioration offers 'positive' evidence in favor of the cost channel model.[10] The posterior mean is close to the point estimates put forward by Ravenna and Walsh (2006) and Chowdhury, Hoffmann, and Schabert (2006), and it is somewhat larger than the one by Rabanal (2007).

The remaining estimated parameters assume values in line with previous contributions (eg Rabanal, 2007; Smets and Wouters, 2007; Justiniano and Primiceri, 2008a). In particular, the NKPC turns out to be purely-forward looking, a finding recently discussed, among others, by Benati (2009), Cogley and Sbordone (2008) in a NKPC in which trend inflation appears as a further driver.[11] Kleibergen and Mavroeidis (2009) perform GMM estimation with identification-robust methods of the semi-structural version of the NKPC we focus on (cost channel aside), and cannot reject the null of purely forward looking inflation process. The demand shock is fairly persistent, but the estimated autoregressive parameter value is far from unity. This suggests that in the model there is a propagation mechanism of the shocks capable to replicate the unit-root like dynamics of output without the need of imposing a unit-root (or almost unit-root) IS disturbance.

Figure 3 – top panel contrasts actual series with the model's one-step-ahead predictions. Indeed, the model is successful in predicting our observables. Figure 3 – bottom panel contrasts actual inflation to the estimated time-varying inflation target set by the Fed in the post-WWII. The pattern followed by the estimated target is quite sensible, and clearly falls in the credible set estimated by Cogley and Sbordone (2008). We take the evidence proposed in Figure 3 as supporting the empirical ability of our framework to fit the post-WWII US data.

[9] We computed the log-marginal likelihoods both by means of the Laplace approximation around the posterior mode (based on a normal distribution) and via the modified harmonic mean estimator (Geweke, 1998), which exploits the draws from the posterior distribution. The two methods deliver virtually identical results. This is due to the close-to-normal distribution of all the estimated posteriors. Given the large computational gains implied by the Laplace approximation, we employ this approximation for our model comparison.

[10] According to Kass and Raftery (1995), a Bayes factor between 1 and 3 is 'not worth more than a bare mention', between 3 and 20 suggests a 'positive' evidence in favor of one of the two models, between 20 and 150 suggests a 'strong' evidence against it, and larger than 150 'very strong' evidence.

[11] Technically, Benati (2009) and Cogley and Sbordone (2008) consider NKPC curves log-linearized around a positive value for the inflation rate in steady-state, ie 'trend inflation' as popularized by Ascari (2004). Differently, we consider a model consistent with a zero inflation rate in steady-state.

Ultimately, however, this analysis aims at pinning down the reaction of inflation to a monetary policy shock. Figure 4 – top panel shows the estimated dynamics responses to such a shock. The data speak clearly. The presence of an active cost channel is far from overturning conventional wisdom: A monetary policy tightening opens a recession, and such downward demand pressure leads to a statistically significant *deflationary* phase. The output gap then follows an hump-shaped convergence pattern towards the steady state. The policy rate and the inflation rate also gradually go back to their steady states.

While being statistically significant, the economic role of the cost channel appears to be present but limited. Figure 5 displays the impulse response of the benchmark model to the four estimated shocks, and contrasts them with those estimated in the 'no cost channel' scenario. Indeed, differences appear to be marginal, with the exception of the reaction of the output gap to a trend inflation shock, which is clearly milder when the cost channel is considered, possibly due to more moderate interest rate deviations from the steady state. These responses suggest that the relevance of the cost channel might be conditional to the type of shock a researcher is interested into.

3.3 Interpreting the result: The role of interest rate smoothing

We have established that, at an empirical level, the demand channel overwhelms the cost channel in translating the impact of the policy rate hike on inflation. Why is it so? A scrutiny of the estimated parameters of the model and their impact on inflation reaction guided us to the *role of interest rate smoothing*. Getting back to Figure 1, one may indeed notice that the sensitivity of the sign of the inflation reaction under a plausible calibration is mainly driven, in terms of *magnitude*, by such parameter. Possibly, this is so because inflation expectations are strongly influenced by the monetary policy conduct given the latter's persistence (Woodford, 1999; Woodford, 2003b), then an increase in marginal costs driven by borrowing costs is less inflationary than it would be under a less firm policy conduct. Indeed, when re-estimating the model by setting $\phi_R = 0$, one may observe a substantially weakened estimated effect of monetary policy on inflation, as shown by Figure 6. Still, an interest rate hike is deflationary (confidence bands, not shown, confirm that it is statistically so). This is due to the fact that we re-estimated the entire model under $\phi_R = 0$, so allowing all the other parameters to adjust to fit the data at best. In particular, the cost channel, habit formation, and Taylor rule parameters all adjust downwards, so partly 'counter-balancing' the inflationary effect of the missing interest rate smoothing (results available upon request). However, the marginal likelihood of the model with no interest rate smoothing clearly works against it, reading -680.15, ie a deterioration of about 257 log-points. Conditionally to the model at hand, the data clearly suggest a rejection of the structural interpretation of the price puzzle.

4 Robustness checks

We check the robustness of our findings along three dimensions:

- *Subsample stability.* The analysis developed so far has relied on the assumption of stability of the structural parameters in the sample at hand, as in Smets and Wouters (2007) and Justiniano and Primiceri (2008a). However, the appointment of Paul Volcker as Chairman of the Fed has been associated to a break in the US monetary policy conduct (Clarida, Galì, and Gertler, 2000; Lubik and Schorfheide, 2004; Boivin and Giannoni, 2006; Benati and Surico, 2008b; and Mavroeidis, 2009). To control for these break, we re-estimate the model by focusing on the subsamples 1954:III–1979:II and 1982:IV–2008:II. We do not include the span 1979:III–1982:III not to deal with the 'Volcker experiment', ie the period during which Chairman Paul Volcker targeted non-borrowed reserves, a monetary policy hard to describe with a Taylor rule.

  Our results are displayed in Table 1 (2nd and 3rd column) and Figure 3 (2nd and 3rd row). The two main messages are robust to this subsample analysis: i) There is an active cost channel, whose importance is supported by the marginal likelihood comparison in the second subsample. In fact, the first one is less supportive, but it is still hard to clearly reject its importance; ii) the effect of a monetary policy tightening is clearly deflationary. As regards the remaining parameters, one may notice that the systematic reaction to inflation gap fluctuations is larger in the second subsample, a finding in line with several recent studies (Lubik and Schorfheide, 2004; Boivin and Giannoni, 2006; Cogley, Primiceri, and Sargent, 2009; Benati and Surico, 2008b; and Mavroeidis, 2009). In the first subsample, such estimated reaction is larger than what typically found in the literature. This is due to the fact that we truncate the parameter space and concentrate on the 'determinacy territory'.[12] Moreover, in this model the object targeted by the Fed is the inflation *gap*, as opposed to the *raw* inflation rate typically considered in Taylor rule estimations – notable exceptions being Castelnuovo, Greco and Raggi (2008) and Cogley, Primiceri, and Sargent (2009). The estimated shocks' standard deviations are clearly smaller in the 'Great Moderation' subsample, a finding already put forward by Justiniano and Primiceri (2008b). The remaining parameters, cost channel parameter included, display stability over subsamples, with the exception of habit formation,

[12] Clarida, Galì, and Gertler (2000), Lubik and Schorfheide (2004), Boivin and Giannoni (2006), and Mavroeidis (2009) offer support to the 'indeterminacy' hypothesis to explain the US macroeconomic dynamics in the 1970s. Castelnuovo and Surico (2009) show that indeterminacy may offer a rationale for the price puzzle typically found when estimating the effects of a monetary policy shocks with VAR models. Surico (2006) discusses the perils coming from merging two subsamples featuring different equilibria. However, Sims and Zha (2006), Justiniano and Primiceri (2008b) and Cogley, Primiceri and Sargent (2009) cast doubs on multiple equilibria as a relevant feature to describe the dynamics of the 1960s and 1970s. Moreover, Castelnuovo (2009) shows that the equilibrium selection strategy one implements under indeterminacy may importantly drive the model consistent theoretical volatilities. We then decided to stick to the uniqueness scenario.

which increases.

- *Alternative inflation measures.* In our benchmark exercise we consider the GDP deflator inflation as relevant measure of inflation. However, alternative measures of inflation may be considered. We then repeat our exercise either with the growth rate of Personal Consumption Expenditure (PCE) or Consumer Price Index (CPI) (less food and energy). Figure 7 contrasts the benchmark analysis with alternatives. Some remarks are in order. The alternative measures of inflation suggest a weaker impact of the real ex-ante interest rate on aggregate demand and, consequently, a less deflationary effect. This is in part due to the rightward shift of the cost channel parameter's posterior densities, whose mean is 1.36 and [5th, 95th] percentiles are $[0.82, 1.96]$ in the case of PCE inflation, and $1.83 - [1.14, 2.48]$ in that of CPI inflation. Interestingly, the model with cost channel is favored by the data also when these alternative inflation indicators are employed. In particular, the marginal likelihood suggests a difference of $(-413.59)-(-417.37) \approx 3.8$ log-points when PCE inflation is considered, and $(-420.63) - (-427.64) \approx 7$ log-points as for CPI core inflation. However, and most importantly for our study, inflation's reaction is still clearly negative.

- *Autocorrelated monetary policy shocks.* Rudebusch (2002) states that the smooth behavior of the policy rate observed in the US (and a variety of other countries) is not intentionally implemented by the Fed, but it is instead caused by serially correlated monetary policy shocks. The Taylor rule then reads as follows:

$$
\begin{aligned}
R_t &= \phi_R R_{t-1} + (1-\phi_R)[\phi_\pi(\pi_t - \pi_t^*) + \phi_x x_t] + \varepsilon_t^R \\
\varepsilon_t^R &= \rho_\varepsilon \varepsilon_{t-1}^R + \eta_t^R
\end{aligned}
$$

According to this policy rule monetary authorities behave in a past-dependent fashion also under $\phi_R = 0$. This is so because the shock ε_t^R is a state variable of the system. We assume the prior density $\rho_\varepsilon \sim \beta(0.5, 0.285)$. Figure 7 shows also the estimated reactions under this alternative specification of the policy rule. The presence of an autoregressive monetary policy shock reinforces the impact of the monetary policy move by inducing a more severe recession and a larger deflation. In terms of model fit, we notice that the model with autoregressive policy shocks's marginal likelihood reads -419.80, and it is then to be preferred to the 'interest rate smoothing only' framework – the marginal likelihood difference, in terms of log-points, is about 4. However, this result does not offer support to Rudebusch's (2002) 'interest rate smoothing illusion' argument. In fact, the marginal likelihood of a model estimated under $\phi_R = 0, \rho_\varepsilon > 0$ reads -506.60, a dramatic deterioration of the model fit – around 83 log-points. This finding corroborates previous research by English, Nelson, and Sack (2003) and Castelnuovo (2003), who support the hypothesis of gradualism intentionally pursued by the Fed in the post-WWII sample.

5 VAR misspecification and the price puzzle

Our exercise leads to a rebuttal of the structural interpretation of the VAR evidence. Rabanal (2007) reaches the same conclusion by focusing on the importance of different sources of persistence in the Christiano, Eichenbaum, and Evans (2005) model. Then, if estimated models do not offer a rationale for the price puzzle, why do we observe it in VARs?

Sims (1992) was the first to point out that the price puzzle is likely to be due to a *misspecification* of the monetary policy shock. In fact, if the central bank reacts to expected inflation, then a predicted upcoming surge in inflation will be followed by an increase in the policy rate, a decrease in the output gap, and – as long as the monetary policy tightening is not such to fully offset the inflationary shock – a rise in current inflation. If the VAR omits expected inflation, and if expected inflation and current inflation are not strictly linked (ie current inflation is not a 'sufficient statistic' for expected inflation), then the supposed-to-be monetary policy shock in a trivariate VAR in inflation, output gap, and policy rate will somewhat naturally capture the positive correlation between inflation and the policy rate, ie it will produce a price puzzle. Sims (1992) proposed to add an indicator of nascent inflation (commodity prices) to the vector of variables of interest. While not solving the price puzzle problem, this trick clearly renders the picture less puzzling.

The omitted variable issue is also tackled by Bernanke, Boivin, and Eliasz (2005) and Boivin, Giannoni, and Mihov (2009), who show that by allowing some factors extracted by a large panel of variables to enter the vector autoregression (as 'endogenous variables') the price puzzle tends to disappear. Forni and Gambetti (2008) focus on an open economy VAR and show that both the price puzzle and the forward discount puzzle – which refers to the small scale VAR evidence of a 'delayed overshooting' – disappear when a data-rich approach in the context of a structural factor model is considered. Castelnuovo and Surico (2009) show that the price puzzle evidence is actually limited to the pre-Volcker subsample – similar evidence is provided by Barth and Ramey (2001), Hanson (2004) and Boivin and Giannoni (2006). Working with a model in which they simulate a policy shift resembling the one estimated for the US case, Castelnuovo and Surico (2009) show that a standard trivariate VAR estimated on pseudo-data may indeed produce a price puzzle when, in fact, the model generating such pseudo-data suggests a *negative* inflation reaction to a policy tightening. They show that this is possibly due to the omission in the VAR of inflation expectations under the weak monetary policy scenario.

Interestingly, some of the best predictors of future inflation turn out to be useless for correcting the bias in the dynamics of inflation, as shown by Hanson (2004). Other recent contributions have pointed towards other types of VAR misspecifications. Leeper and Roush (2003) show that money is important for well specifying the monetary policy shock when studying economies in which a double-causal link between money and interest rate might have occurred. In particular, if the central bank reacts contemporaneously to monetary aggregates, and if money demand is contemporaneously driven by the nominal interest rate, then the omission of money would lead to a misspecification of the monetary policy shock. A different issue is raised by

Giordani (2004), who shows that the omission of potential output in standard trivariate VARs may severely bias impulse responses and be the responsible of the price puzzle. In fact, potential output appears in all the equations of a standard new-Keynesian AD/AS model. Hence, its omission will lead supposed-to-be shocks to be residuals correlated across the VAR equations, and consequently to produce biased impulse response functions. Romer and Romer (2004) stick on a standard trivariate VAR but produce a careful measure of the monetary policy shocks based on changes in the intended federal funds rate and the Fed's expectations on future inflation and output. Such new measure of monetary policy shock does not imply any price puzzle in their estimated VARs.

While presenting somewhat different views on how to model a monetary policy shock in a VAR framework, these papers clearly express a common vieon the 'price puzzle', ie they qualify it as an 'artifact' due to model misspecification, more than a genuine 'fact'.

6 Conclusions

This paper showed that a new-Keynesian model embedding the cost channel may hardly offer a rationale for the price puzzle typically found when conducting VAR analysis. Under some particular parameterizations of the model, a positive inflation reaction to an unexpected, restrictive monetary policy may actually arise. However, when taking the model to the data, the structural interpretation of the VAR evidence is clearly rebutted. The impact exerted by the estimated systematic monetary policy gradualism is shown to possibly drive this result. Our findings are robust to several perturbations to the baseline analysis, including different sample selection, alternative inflation indicators, and a different statistical model for the monetary policy shock. We think of this result as being important for understanding the sign (and the magnitude) that monetary policy actions should take in response to shocks moving inflation off target.

We stress that this paper does *offer* some evidence in favor of the cost channel. In particular, the presence of such channel seems to be economically important when assessing the reaction of output to a trend inflation shock. In general, the structural role of the interest rate in the Phillips curve calls for a serious re-thinking of optimal monetary policy in presence of supply effects. Contributions along this path have recently been proposed by Ravenna and Walsh (2006) and Kilponen and Milne (2007). Moreover, given the uncertainty surrounding the magnitude of the cost channel parameter, more research is needed both for the quantification of the importance of the cost channel and for the design of an optimal monetary policy in presence of cost channel uncertainty, an issue recently tackled by Tillmann (2009). Llosa and Tuesta (2009) analyze the relationship between uniqueness and learnability of equilibrium in presence of supply effects, a topic of great relevance for policymakers. We plan to participate to this exciting agenda with further investigations in the close future.

References

An, S – Schorfheide, F (2007) **Bayesian Analysis of DSGE Models.** Econometric Reviews, 26, 113–172.

Ascari, G (2004) **Staggered Prices and Trend Inflation: Some Nuisances.** Review of Economic Dynamics, 7, 642–667.

Bache, I – Leitemo, K (2008) **The Price Puzzle: Mixing the Temporary and Permanent Monetary Policy Shocks** Norges Bank Working Paper 2008/18.

Barth, M – Ramey, V (2001) **The Cost of Monetary Transmission.** NBER Macroeconomic Annual, 199–239

Benati, L (2008a) **Investigating Inflation Persitence Across Monetary Regimes.** The Quarterly Journal of Economics, 123(3), 1005–1060.

Benati, L (2009) **Are 'Intrinsic Inflation Persistence' Models Structural in the Sense of Lucas (1976)?** European Central Bank Working Paper No. 1038.

Benati, L – Surico, P (2008a) **Evolving U.S. Monetary Policy and the Decline of Inflation Predictability.** Journal of the European Economic Association, 6(2–3), 634–646.

Benati, L – Surico, P (2008b) **VAR Analysis and the Great Moderation.** The American Economic Review, forthcoming.

Bernanke, B – Boivin, J – Eliasz, P (2005) **Measuring Monetary Policy: A Factor Augmented Vector Autoregressive (FAVAR) Approach.** The Quarterly Journal of Economics, 120(1), 387–422.

Boivin, J – Giannoni, M (2006) **Has Monetary Policy Become More Effective?** The Review of Economics and Statistics, 88(3), 445–462.

Boivin, J – Giannoni, M – Mihov, I (2009) **Sticky Prices and Monetary Policy: Evidence from Disaggregated U.S. Data.** The American Economic Review, 99(1), 350–384.

Brooks, S – Gelman, A (1998) **General Methods for Monitoring Convergence of Iterative Simulations.** Journal of Computational and Graphical Statistics, 7(4), 434–455.

Brueckner, M – Schabert, A (2003) **Supply Side Effects of Monetary Policy and Equilibrium Multiplicity.** Economics Letters, 79, 205–211.

Brunnermeier, M (2009) **Deciphering the Liquidity and Credit Crunch 2007–2008.** Journal of Economic Perspectives, 23(1), 77–100.

Canova, F (2009) **What Explains the Great Moderation in the US? A Structural Analysis.** Journal of the European Economic Association, 7(4), 697–721.

Canova, F – Sala, L (2009) **Back to Square One: Identification Issues in DSGE Models.** Journal of Monetary Economics, 56(4), 431–449.

Castelnuovo, E (2003) **Taylor Rules, Omitted Variables, and Interest Rate Smoothing in the US.** Economics Letters, 81(1), 55–59.

Castelnuovo, E (2009) **Policy Switch and the Great Moderation: The Role of Equilibrium Selection.** University of Padua, mimeo.

Castelnuovo, E – Greco, L – Raggi, D (2008) **Estimating Regime Switching Taylor Rules with Trend Inflation.** Bank of Finland Discussion Paper 20/2008.

Castelnuovo, E – Surico, P (2009) **Monetary Policy Shifts, Inflation Expectations and the Price Puzzle.** Economic Journal, forthcoming.

Chowdhury, I – Hoffmann, M – Schabert, A (2006) **Inflation Dynamics and the Cost Channel of Monetary Transmission.** European Economic Review, 50, 995–1016.

Christiano, L – Eichenbaum, M – Evans, C (2005) **Nominal Rigidities and the Dynamic Effects of a Shock to Monetary Policy.** Journal of Political Economy, 113(1), 1–45.

Clarida, R – Galì, J – Gertler, M (2000) **Monetary Policy Rules and Macro-economic Stability: Evidence and Some Theory.** Quarterly Journal of Economics, 115, 147–180.

Cogley, T – Primiceri, G E – Sargent, T (2009) **Inflation-Gap Persistence in the U.S.** New York University and Northwestern University, mimeo.

Cogley, T – Sbordone, A (2008) **Trend Inflation, Indexation, and Inflation Persistence in the New Keynesian Phillips Curve.** The American Economic Review, 98(5), 2101–2126.

Dedola, L – Lippi, F (2005) **The Monetary Transmission Mechanism: Evidence from the Industries of 5 OECD Countries.** European Economic Review, 49(6), 1543–1569.

Eichenbaum, M (1992) **Comment on 'Interpreting the Macroeconomic Time Series Facts: The Effects of Monetary Polisy' by C A Sims.** European Economic Review, 36(5), 1001–1011.

English, W – Nelson, W – Sack, B (2003) **Interpreting the Significance of the Lagged Interest Rate in Estimated Monetary Policy Rules.** Contributions to Macroeconomics, 3(1), Article 5.

Fanelli, L (2008) **Testing the New Keynesian Phillips Curve Through Vector Autoregressive Models: Results from the Euro Area.** Oxford Bulletin of Economics and Statistics, 70(1), 53–66.

Forni, M – Gambetti, L (2008) **The Dynamic Effects of Monetary Policy: A Structural Factor Model Approach.** University of Modena-Reggio Emilia and UAB.

Gaiotti, E – Secchi, A (2006) **Is There a Cost Channel of Monetary Policy Transmission? An Investigation Into the Pricing Behavior of 2,000 Firms.** Journal of Money, Credit and Banking, 38(8), 2013–2037.

Galì, J (2008) **Monetary Policy, Inflation and the Business Cycle.** Princeton University Press.

Geweke, J (1998) **Using Simulation Methods for Bayesian Econometric Models: Inference, Development and Communication.** Federal Reserve Bank of Minnesota Staff Report 249.

Giordani, P (2004) **An Alternative Explanation to the Price Puzzle.** Journal of Monetary Economics, 51, 1271–1296.

Hanson, M (2004) **The 'price Puzzle' Reconsidered.** Journal of Monetary Economics, 51, 1385–1413.

Henzel, S – Hulsewig, O – Mayer, E – Wollmershaeuser, T (2007) **The Price Puzzle Revisited: Can the Cost Channel Explain a Rise in Inflation After a Monetary Policy Shock?** CESIfo Working Paper Series, 2039.

Ireland, P (2004) **Technology Shocks in a New Keynesian Model.** The Review of Economics and Statistics, 86(4), 923–936.

Ireland, P (2007) **Changes in Federal Reserve's Inflation Target: Causes and Consequences.** Journal of Money, Credit and Banking, 39(8), 1851–1882.

Justiniano, A – Primiceri, G (2008a) **Potential and Natural Output.** Federal Reserve Bank of Chicago and Northwestern University, mimeo.

Justiniano, A – Primiceri, G (2008b) **The Time-Varying Volatility of Macroeconomic Fluctuations.** The American Economic Review, 98(3), 604–641.

Kilponen, J – Milne, A (2007) **The Lending Channel under Optimal Choice of Monetary Policy.** Bank of Finland Discussion Paper No. 33/2007.

Kleibergen, F – Mavroeidis, S (2009) **Weak Instrument Robust Test in GMM and the New Keynesian Phillips Curve.** Journal of Business and Economic Statistics, forthcoming.

Leeper, E – Roush, J (2003) **Putting 'M' Back in Monetary Policy.** Journal of Money, Credit and Banking, 35(2), 1217–1256.

Llosa, L – Tuesta, V (2009) **Learning About Monetary Policy Rules When the Cost-Channel Matters.** Journal of Economic Dynamics and Control, forthcoming.

Lubik, T – Schorfheide, F (2004) **Testing for Indeterminacy: An Application to U.S. Monetary Policy.** The American Economic Review, 94(1), 190–217.

Mavroeidis, S (2004) **Weak Identification of Forward-Looking Models in Monetary Economics.** Oxford Bulletin of Economics and Statistics, 66 (Supplement), 609–635.

Mavroeidis, S (2009) **Monetary Policy Rules and Macroeconomic Stability: Some New Evidence.** The American Economic Review, forthcoming.

Rabanal, P (2007) **Does Inflation Increase After a Monetary Policy Tightening? Answers Based on an Estimated DSGE Model.** Journal of Economic Dynamics and Control, 31, 906–937.

Ravenna, F – Walsh, C (2006) **Optimal Monetary Policy with the Cost Channel.** Journal of Monetary Economics, 53, 199–216.

Romer, C – Romer, D (2004) **A New Measure of Monetary Policy Shocks: Derivation and Implications.** The American Economic Review, 94(4), 1055–1084.

Rudebusch, G (2002) **Term Structure Evidence of Interest Rate Smoothing and Monetary Policy Inertia.** Journal of Monetary Economics, 49, 1161–1187.

Sims, C (1992) **Interpreting the Macroeconomic Time Series Facts: The Effects of Monetary Policy.** European Economic Review, 36(5), 975–1000.

Sims, C – Zha, T (2006) **Were There Regime Switches in U.S. Monetary Policy?** The American Economic Review, 96(1), 54–81.

Smets, F – Wouters, R (2007) **Shocks and Frictions in US Business Cycle: A Bayesian DSGE Approach.** The American Economic Review, 97(3), 586–606.

Stock, J – Watson, M (2001) **Vector Autoregressions.** Journal of Economic Perspectives, 15, 101–115.

Surico, P (2006) **Monetary Policy Shifts and Inflation Dynamics.** In Cobham, D (ed.): The Travails of the Eurozone, Palgrave, London.

Surico, P (2008) **The Cost Channel of Monetary Policy and Indeterminacy.** Macro-economic Dynamics, 12, 724–735.

Tillman, P (2008) **Does the Cost Channel on Monetary Transmission Explain Inflation Dynamics?** Journal of Economic Dynamics and control, 32, 2723–2744.

Tillman, P (2009) **Optimal Monetary Policy with an Uncertain Cost Channel.** Journal of Money, Credit, and Banking, forthcoming.

Woodford, M (1999) **Optimal Monetary Policy Inertia.** Princeton University, mimeo.

Woodford, M (2003a) **Interest and Prices: Foundations of a Theory of Monetary Policy.** Princeton University Press.

Woodford, M (2003b) **Optimal Interest-Rate Smoothing.** Review of Economic Studies, 70, 861–886.

Param.	*Prior Dens.*	*Posterior Means* [$5h,95th$]		
		1954:III–2008:II	1960:I–1979:II	1982:IV–2008:II
ψ	$N(1.75, 0.7)$	1.18 [0.60,1.75]	1.01 [0.22,1.74]	1.12 [0.36,1.84]
α	$\beta(0.5, 0.285)$	0.01 [0.00,0.02]	0.02 [0.00,0.05]	0.01 [0.00,0.03]
σ	$N(1, 0.05)$	0.89 [0.80,0.96]	0.94 [0.86,1.02]	0.94 [0.85,1.02]
h	$\beta(0.7, 0.15)$	0.78 [0.71,0.86]	0.68 [0.56,0.80]	0.82 [0.74,0.90]
ϕ_π	$N(1.7, 0.3)$	1.87 [1.49,2.24]	1.70 [1.31,2.06]	1.91 [1.54,2.31]
ϕ_x	$\Gamma(0.3, 0.2)$	0.72 [0.43,1.01]	0.47 [0.23,0.69]	0.61 [0.30,0.91]
ϕ_R	$\beta(0.5, 0.285)$	0.93 [0.91,0.95]	0.90 [0.86,0.94]	0.94 [0.92,0.96]
ρ_x	$\beta(0.5, 0.285)$	0.40 [0.27,0.52]	0.45 [0.29,0.62]	0.52 [0.38,0.68]
σ_x	$I\Gamma(0.1, 0.25)$	0.43 [0.36,0.49]	0.51 [0.38,0.63]	0.22 [0.16,0.28]
σ_π	$I\Gamma(0.1, 0.25)$	0.25 [0.21,0.30]	0.34 [0.26,0.42]	0.21 [0.16,0.26]
σ_R	$I\Gamma(0.1, 0.25)$	0.22 [0.21,0.24]	0.19 [0.16,0.21]	0.14 [0.12,0.15]
σ_*	$I\Gamma(0.1, 0.25)$	0.05 [0.03,0.08]	0.06 [0.03,0.09]	0.04 [0.03,0.06]
$Log(ML)$	–	−423.64	−229.62	−79.84
$Log(ML\,\vert_{\psi=0})$	–	−425.59	−227.78	−96.11

Table 1: **Bayesian estimates of the benchmark model.** Full sample and subsample posterior densities. Prior densities: Figures indicate the (mean,st.dev.) of each prior distribution. Posterior densities: Figures reported indicate the posterior mean and the [5th, 95th] percentile of the estimated densities. Details on the estimation procedure provided in the text. Marginal likelihoods computed via Laplace approximation.

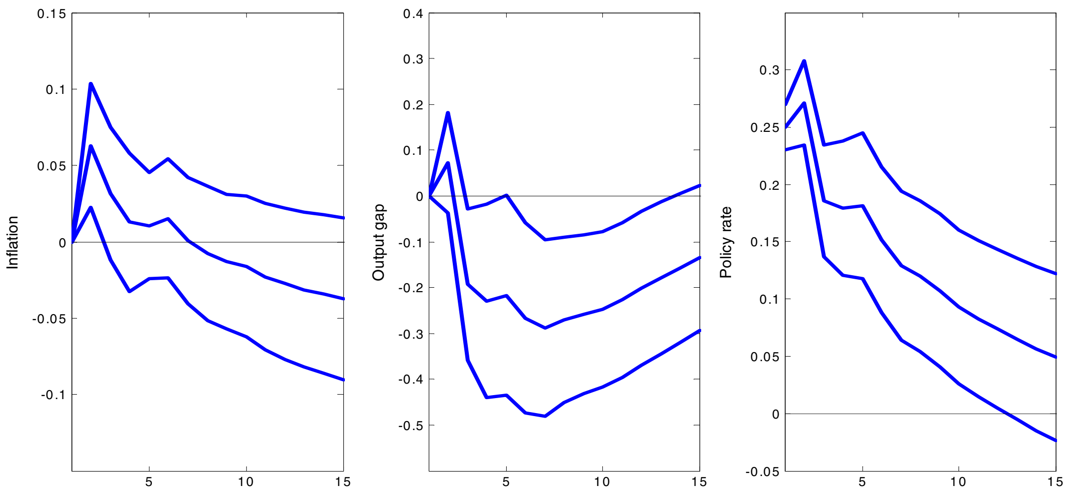

Figure 1: **SVAR impulse response functions to a monetary policy shock.** Sample: 1954:III–2008:II. Variables: Quarterly GDP inflation, CBO output gap, quarterly federal funds rate – source: FREDII. Identification of the monetary policy shock via Cholesky decomposition (lower triangular matrix, ordering: quarterly inflation, output gap, quarterly federal funds rate). Solid line: Mean response. Dotted lines: 90% confidence bands (analytically computed). VAR estimated with four lags.

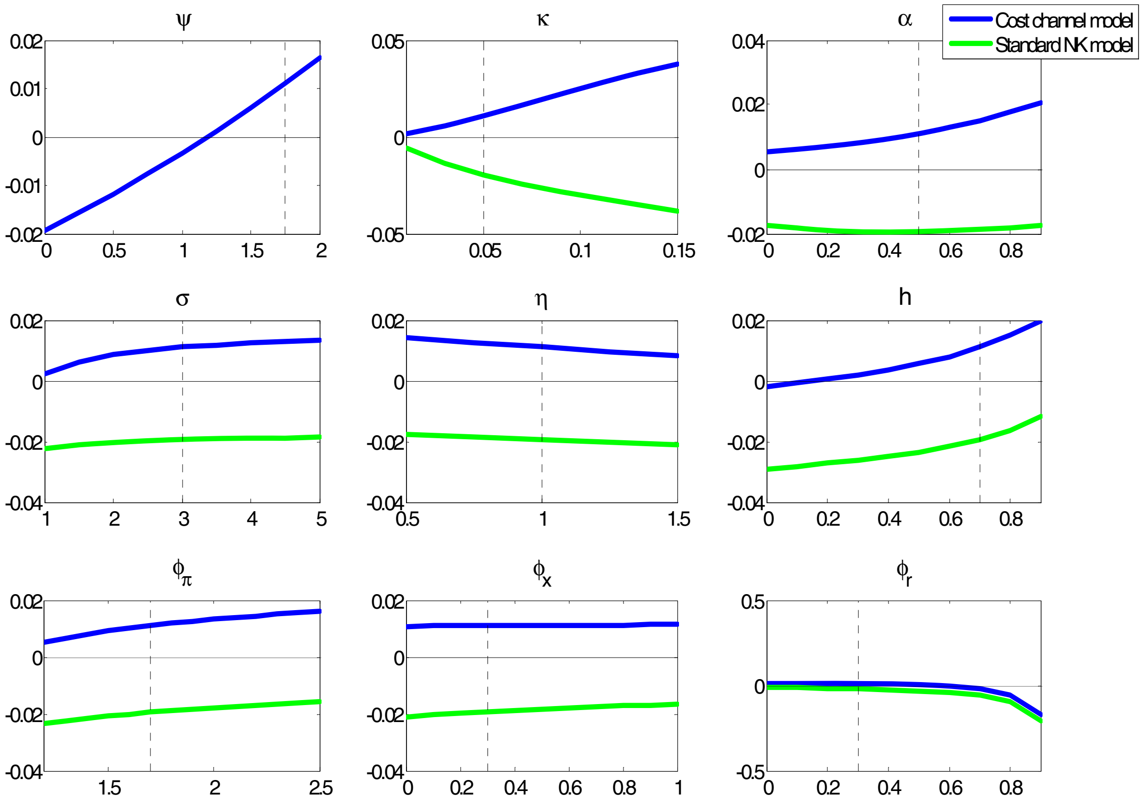

Figure 2: **On-impact inflation reaction to a monetary policy shock under different calibrations.** Baseline calibration (indicated by vertical black dotted lines): $\beta = 0.99, \psi = 1.75, \kappa = 0.05, \alpha = 0.5, \sigma = 3, \eta = 1, h = 0.7, \phi_\pi = 1.7, \phi_x = 0.3, \phi_R = 0.3$, Blue solid line: Cost channel model. Green dotted line: Model with $\psi = 0$. Monetary policy shock calibrated to induce a 25 basis points jump of the policy rate right after the shock (not shown).

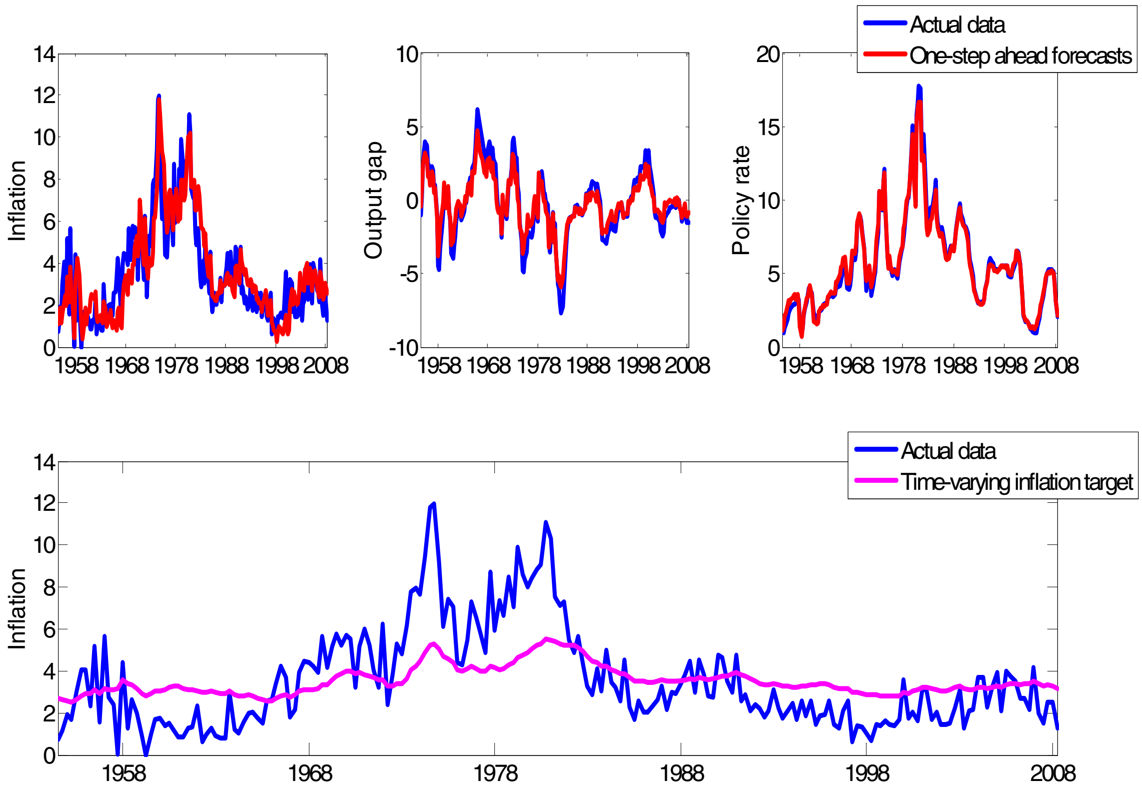

Figure 3: **Estimated U.S. dynamics.** Top panels: Solid blue lines: Actual series. Dotted red lines: model-based one-step-ahead forecasts (filtered latent factors). Bottom panel: Solid blue line: Actual inflation. Dotted magenta line: Estimated time-varying inflation target (smoothed latent factor). Sample means added back to latent and actual series in a model-consistent fashion.

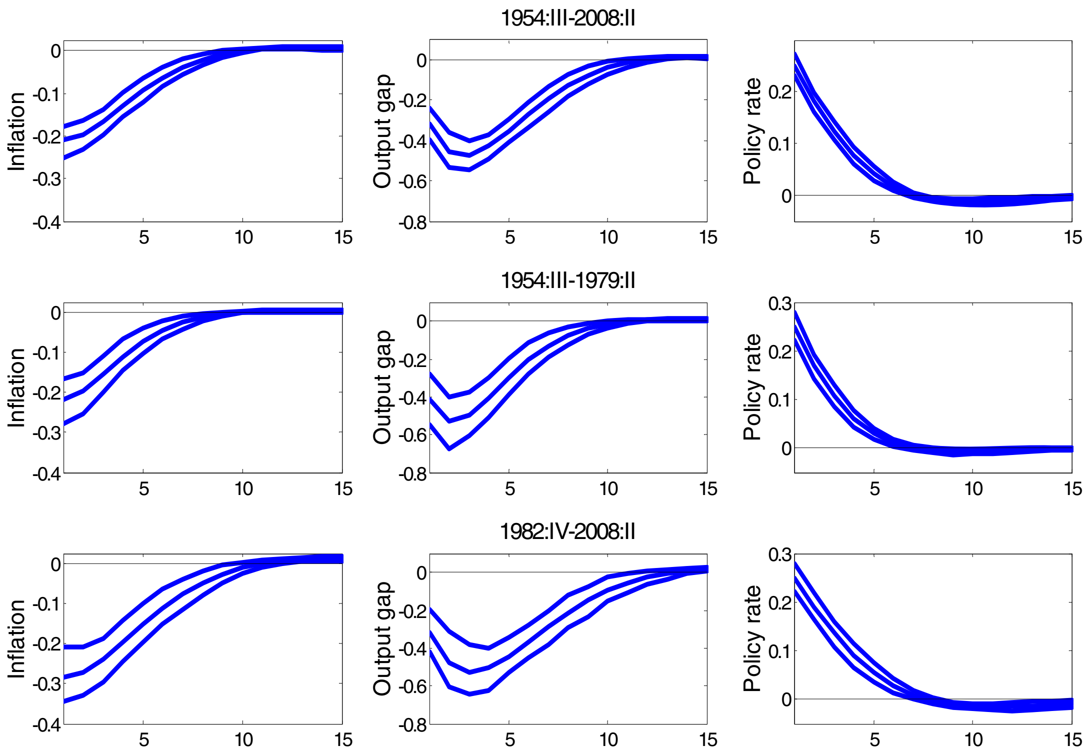

Figure 4: **Bayesian impulse response functions to a monetary policy shock.** Solid lines: mean impulse response. Dotted lines: 5th and 95th percentiles of the posterior distributions. Shock size normalized so to induce a 25 basis point-jump of the quarterly policy rate.

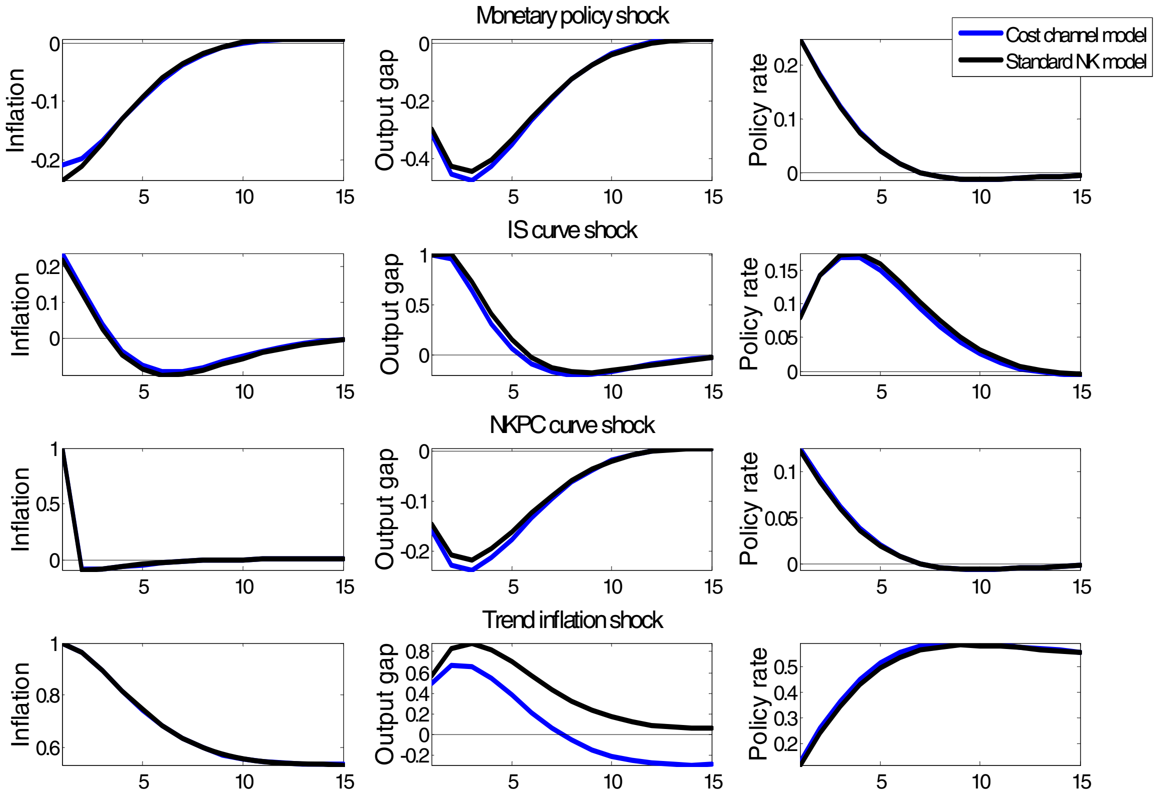

Figure 5: **Bayesian impulse responses: Role of the cost channel.** Standard NK model: Model estimated under $\psi = 0$. Shocks normalized so to render the dynamic responses of the two models comparable.

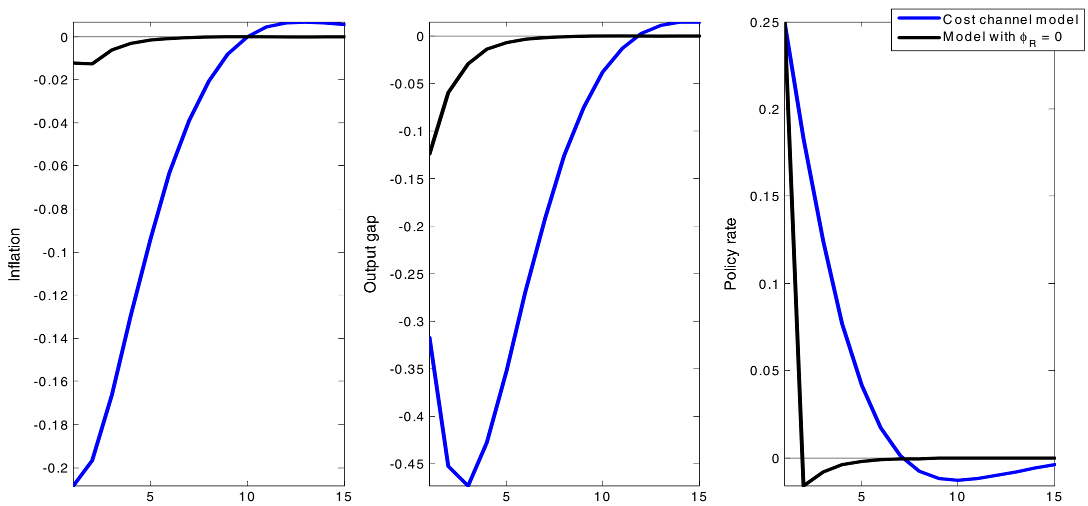

Figure 6: **Bayesian impulse responses: The role of interest rate smoothing.** Shock size normalized so to induce a 25 basis point-jump of the quarterly policy rate.

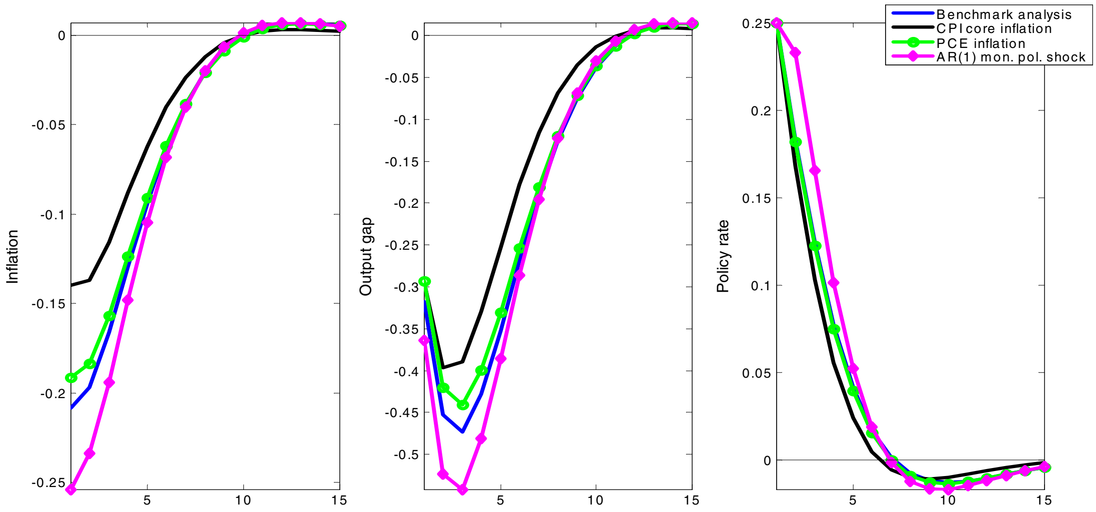

Figure 7: **Bayesian impulse responses: Alternative scenarios.** Shock size normalized so to induce a 25 basis point-jump of the quarterly policy rate.

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