In Cholesky-VARs We Trust? An Empirical Investigation with U.S. Data*

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Abstract

Two estimated new-Keynesian DSGE models are employed in turn to generate artificial data in a MonteCarlo exercise in which the effects of a monetary policy shock on inflation and output are computed via Cholesky-VARs. Compelling empirical evidence documenting the substantial distortions that affect such responses is provided. The wrongly assumed policy delays consistent with the recursive identification scheme are responsible for this result. Our MonteCarlo Cholesky-VAR impulse responses are shown to replicate to a large extent those obtained with actual U.S. data as for the great moderation phase. This result offers a novel interpretation of the mild-to-muted reactions of inflation and output in the post-Volcker era.

JEL classification: C22, E47, E52.

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1 Introduction

Vector AutoRegressions (VARs) have been employed to estimate the effects of monetary policy shocks on output and inflation for at least three decades (Sims (1980)). A researcher typically i) estimates a reduced-form VAR, ii) applies an identification scheme to isolate the effects of a policy shock, and iii) produces moments of interest such as, e.g., impulse responses to the identified shock.

The most popular identification strategy as for monetary policy shocks is the "Cholesky" scheme. Such scheme orders the monetary policy indicator (typically, a short-term interest rate) after "slow moving" variables such as inflation and output in the vector of modeled variables. The underlying assumptions are that i) monetary authorities contemporaneously react to macroeconomic indicators; ii) inflation and output are affected by policy shocks with a lag. The nice feature of this strategy is that it does not require the researcher to take a position on the identification of other shocks (see Christiano, Eichenbaum, and Evans (1999) for an extensive discussion on this issue). Once identification is achieved, an analysis of the magnitude, persistence, and significance of the estimated impulse responses to the "Cholesky policy shock" can be conducted. The idea is, first of all, that of understanding if the modeled variables do react to a policy shock or not. It this context, a "representative" statement is " ... given that variable X displays an insignificant reaction to the policy shock, we infer that monetary policy is not able to affect such variable ... ".

Evidence of mild-to-muted reactions of inflation and output to a policy shock has actually been found by a number of contributions as for the post-Volcker era (Hanson (2004), Boivin and Giannoni (2006), Castelnuovo and Surico (2010)). Boivin and Giannoni (2006) and Boivin, Kiley, and Mishkin (2010) confirm this evidence with a Factor Augmented VAR approach embedding information coming from large datasets.[1] Figure 1 just replicates this evidence.[2] One possible interpretation involves the role of financial

[1]Different results are typically obtained when dealing with longer samples including the 1970s (e.g. Christiano, Eichenbaum, and Evans (2005)). Our Appendix reports results confirming that VARs estimated with samples involving the 1970s do suggest significant reactions of inflation and output to a monetary policy shock. However, Bagliano and Favero (1998) find signs of misspecification affecting VAR models estimated over long samples. In light of the instabilities affecting VAR responses over different samples, we focus our study on the great moderation period.

[2]Evidence obtained with a trivariate VAR including quarterly GDP deflator inflation, a measure of the output gap produced by the Congressional Budget Office, and the federal funds rate (average of monthly observations). Giordani (2004) shows that the estimated responses to a monetary policy shock are likely to be biased if a measure of potential output is omitted from the VAR. The uncertainty surrounding our point estimates, which is comparable to the one in Boivin and Giannoni (2006) and Castelnuovo and Surico (2010), may be due to the moderate volatility of our series in the sample under

innovations occurred in the U.S. in the early 1980s, which may have sharpened households' ability to smooth their consumption out and, therefore, harmed policymakers' ability to affect the demand channel of the monetary policy transmission mechanism. Another interpretation refers to the change in the U.S. systematic monetary policy, which may have stabilized inflation and output more successfully since the advent of Paul Volcker as Chairman of the Federal Reserve at the end of the 1970s.[3]

This paper asks the following question:

Are mild-to-muted Cholesky-VARs responses to a monetary policy shock necessarily a sign of inability by the Federal Reserve to influence the macroeconomic environment?

Our answer is *negative*. Indeed, this paper shows that the evidence proposed in Figure 1 is *fully consistent* with a monetary policy whose shocks exert a substantial influence on the macroeconomic environment.

We support our answer by conducting a MonteCarlo exercise in which two estimated new-Keynesian models of the business cycle are employed in turn as Data-Generating Processes (DGPs) to generate artificial data with which we feed Cholesky-VARs (CVARs). We find evidence of substantial differences between the DSGE-consistent impulse responses and those recovered with our CVARs. In line with conventional wisdom, the estimated DSGE models predict a drop in output and inflation in response to a monetary policy shock. Differently, our CVARs return, on average, *mild-to-muted* reactions of these two variables. This is due the zero restrictions associated to the Cholesky-identification scheme. Such restrictions lead to a misspecified CVAR policy "shock" which is, in fact, a convolution of truly structural shocks exerting offsetting effects on inflation and output. Therefore, our MonteCarlo experiment reveals that *mild-to-muted macroeconomic reactions estimated with a standard CVAR are fully compatible with a monetary policy shock exerting "textbook" effects on inflation and output.*

We conduct our empirical exercise by considering two different DSGE models of the business cycle in turn. The first one is a standard small-scale new-Keynesian model à la King (2000) and Woodford (2003), which features three equations responsible for the evolution of inflation, output, and the policy rate. Despite of its simplicity, this model

scrutiny and/or to the sample size, which is smaller than those typically employed when conducting post-WWII investigations.

[3] "Econometric" intepretations involve small-sample bias issues, which might be severe in a sample like ours, and the misspecification of the monetary policy shock due to the underestimation of the set of variables the Federal Reserve has reacted to. On this latter point, see Barakchian and Crowe (2010), who employ monthly data in their analysis. The relevance of their results at quarterly frequencies as for the great moderation sample is material for future research.

has recently been shown to possess good forecasting properties as for inflation and output when compared to larger scale frameworks (Herbst and Schorfheide (2011)). One interesting feature of this model is that its reduced form is an exact VAR(2) representation, which implies that issues like truncation biases or non-fundamentalness of identified shocks are not theoretically relevant in this context. This enables us to focus on the role and consequences of imposing Cholesky-restrictions under the null hypothesis of contemporaneous monetary policy effects. Then, we move to the medium-scale model à la Smets and Wouters (2007), which has been adopted by a number of central banks for some years now. This model features a variety of nominal and real frictions as well as a number of structural shocks that enhance its ability to track the autocovariances of the macroeconomic data of interest. When conducting our exercise with an estimated version of the Smets and Wouters (2007) model for the post-Volcker era, we find results very similar to those achieved with the small-scale representation of the economy, i.e., mild-to-muted reactions of inflation and output to a monetary policy shock identified with the Cholesky assumption. From a policy standpoint, our conclusions are reassuring, in that this paper proves that econometric evidence as the one shown in Figure 1 may very well arise in a world in which monetary policy shocks do exert an effect on inflation and output.

The reason of this result is the following. Dynamic Stochastic General Equilibrium (DSGE) models typically admit an *immediate* reaction of inflation and output to monetary policy impulses. Differently, CVARs model a *lag* in such reactions. Consequently, under the null hypothesis of the DSGE model being the DGP, CVARs offer a misspecified representation of policy shocks and their effects, because they confound a pure monetary policy innovation with a linear combination of a number of structural shocks exerting offsetting effects on inflation and output. Importantly, from a *theoretical* standpoint the severity of this misspecification may very well range from substantial down to negligible. Canova and Pina (2005) and Carlstrom, Fuerst, and Paustian (2009) show that alternative calibrations of the DGP may give rise to very different assessment on the ability of Cholesky-VARs to recover the effects of monetary policy shocks. Their result is a possibility result, i.e., these papers suggest that the CVAR "can go wrong". But does it go wrong? To stress the relevance of model estimation (as opposed to calibration) in answering this question, we anticipate two possible outcomes of our MonteCarlo exercises in Figure 2. Two alternative calibrations of the same DSGE model (which we present in the next Section) are employed; a set of pseudo-data concerning output, inflation, and the nominal interest rate is generated with our different calibrations in turn; finally, recursive VARs are estimated with our pseudo-data, and the impact of a monetary policy shock is assessed. This Figure clearly points to the relevance of model calibration when assessing Cholesky-VARs. "Calibration A", which is behind the results plotted in the top-row panels, suggests an excellent performance by the CVAR in recovering the "true" effects of a monetary policy shock. Quite differently, "Calibration B" induces a "price puzzle", and "output gap" puzzle, and a largely overestimated policy rate reaction. Interestingly, the difference between these two calibrations concerns a single parameter (the root of the "technology" shock). In other words, exercises based on calibrated models may give rise to orthogonal judgements as for CVARs reliability. This is the reason why, in conducting our Monte Carlo investigations, we will employ *estimated* DSGE models of the business cycle, whose calibration is data-driven. Therefore, our paper represents a further step in the research agenda that aims at understanding the pros and cons of using VARs to identify monetary policy shocks, in that we *empirically* assess their reliability. As anticipated, our results point towards a very poor performance by Cholesky VARs along this dimension.

It is important to note that not all DSGE models in the literature assume an immediate transmission of the monetary policy impulse. Christiano, Eichenbaum, and Evans (2005) and Altig, Christiano, Eichenbaum, and Lindé (2011) are models satisfying the relevant Cholesky-identifying assumptions. If these models were the true DGP, a VAR would probably be able to uncover the true impulse response functions from a monetary policy shock. The models used in this paper feature a contemporaneous timing as in Smets and Wouters (2007), Rabanal (2007), Justiniano and Primiceri (2008), Justiniano, Primiceri, and Tambalotti (2010), Justiniano, Primiceri, and Tambalotti (2011), among others. Formal evidence against zero-restrictions as for the reactions of output and inflation to monetary policy surprises is offered by Del Negro, Schorfheide, Smets, and Wouters (2007) (as for output) and Faust, Swanson, and Wright (2004) (inflation). To the best of our knowledge, however, no empirical test on the relevance of transmission delays (as opposed to contemporaneous effects) in the DSGE modeling context has been conducted so far. We leave this important question to future research.

The paper develops as follows. Section 2 presents and estimates a small-scale standard new-Keynesian DSGE model with U.S. data. Such model is employed as DGP in Section 3, which sets up our MonteCarlo experiment. In this Section we contrast the impulse responses generated with our estimated small-scale DSGE with those coming from the CVARs in a controlled environment, and show that substantial differences arise. An interpretation of our results, which highlight the role of non-policy structural shocks for the Cholesky-VAR responses, is also provided. Section 4 presents our analysis based on the medium-scale model by Smets and Wouters (2007), which confirms our results. Section 5 relates this work to some contributions in the literature. Section 6 concludes.

2 A small-scale DSGE model as DGP

2.1 Model presentation

The first framework we consider is a standard small-scale DSGE model (King (2000), Woodford (2003), Carlstrom, Fuerst, and Paustian (2009)). The log-linearized version of the model is the following:

$$\begin{aligned} \pi_t &= (1+\alpha\beta)^{-1}[\beta E_t \pi_{t+1} + \alpha\pi_{t-1} + \kappa y_t + \varepsilon_t^{\pi}], & (1)\\ y_t &= \gamma E_t y_{t+1} + (1-\gamma) y_{t-1} - \sigma^{-1}(R_t - E_t\pi_{t+1}) + Q(\rho_a - 1)a_t, & (2)\\ R_t &= (1-\tau_R)(\tau_\pi \pi_t + \tau_y y_t) + \tau_R R_{t-1} + \varepsilon_t^R, & (3) \end{aligned}$$

Eq. (1) is an expectational new-Keynesian Phillips curve (NKPC) in which π_t stands for the inflation rate, β represents the discount factor, y_t identifies the output gap, whose impact on current inflation is influenced by the slope-parameter κ, α identifies indexation to past inflation, and ε_t^{π} may be interpreted as a "cost-push shock"; γ is the weight of the forward-looking component in the intertemporal IS curve (2); σ^{-1} is the households' intertemporal elasticity of substitution; the convolution $Q \equiv (1+\nu)(\sigma+\nu)^{-1}$ involves the inverse of the Frisch labor elasticity ν, and a_t is a stochastic component that works as a "technology shock"; τ_π, τ_y, and τ_R are policy parameters in the Taylor rule (3); the monetary policy shock ε_t^R allows for a stochastic evolution of the policy rate.

The model is closed with the following stochastic processes:

$$\begin{bmatrix} \varepsilon_t^{\pi} \\ a_t \\ \varepsilon_t^R \end{bmatrix} = \boldsymbol{F} \begin{bmatrix} \varepsilon_{t-1}^{\pi} \\ a_{t-1} \\ \varepsilon_{t-1}^R \end{bmatrix} + \begin{bmatrix} u_t^{\pi} \\ u_t^a \\ u_t^R \end{bmatrix}, \boldsymbol{F} \equiv \begin{bmatrix} \rho_\pi & 0 & 0 \\ 0 & \rho_a & 0 \\ 0 & 0 & \rho_R \end{bmatrix}, \qquad (4)$$

where the martingale differences, mutually independent processes $\boldsymbol{u}_t$ are distributed as

$$\begin{bmatrix} u_t^\pi \\ u_t^a \\ u_t^R \end{bmatrix} \sim N\left(\begin{bmatrix} 0 \\ 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sigma_\pi^2 & 0 & 0 \\ 0 & \sigma_\pi^2 & 0 \\ 0 & 0 & \sigma_R^2 \end{bmatrix}\right). \tag{5}$$

This or similar small-scale models have successfully been employed to conduct empirical analysis concerning the U.S. economy. Clarida, Galí, and Gertler (2000) and Lubik and Schorfheide (2004) have investigated the influence of systematic monetary policy over the U.S. macroeconomic dynamics; Boivin and Giannoni (2006), Benati and Surico (2009), Canova (2009), and Lubik and Surico (2010) have replicated the U.S. great moderation; Benati (2008) and Benati and Surico (2008) have investigated the drivers of the U.S. inflation persistence; Ireland (2007) and Cogley, Primiceri, and Sargent (2010) have scrutinized the role of shocks to the low frequency component of the U.S. inflation. The employment of this model, on top of offering a first assessment on CVARs' ability to recover the effects of structural shocks, enable us to control for truncation biases possibly arising when employing DSGE models having VARMA representations like Smets and Wouters' (2007), therefore enabling us to focus on the timing discrepancy issue.

2.2 Model estimation

We estimate the model (1)-(5) with Bayesian methods. We work with quarterly U.S. data, sample: 1984:I-2008:II, which roughly corresponds to the great moderation (McConnell and Perez-Quiros (2000)). Our sample ends in 2008:II to exclude the acceleration of the financial crises began with the bankruptcy of Lehman Brothers in September 2008, which triggered non-standard policy moves by the Federal Reserve (Brunnermeier (2009)). We employ three observables, which we demean prior to estimation. The output gap is computed as log-deviation of the real GDP with respect to the potential output estimated by the Congressional Budget Office. The inflation rate is the quarterly growth rate of the GDP deflator. For the short-term nominal interest rate we consider the effective federal funds rate expressed in quarterly terms (averages of monthly values). The source of the data is the Federal Reserve Bank of St. Louis' website.

The vector $\boldsymbol{\xi} = [\beta, \upsilon, \kappa, \alpha, \gamma, \sigma, \tau_\pi, \tau_y, \tau_R, \rho_a, \rho_\pi, \rho_R, \sigma_a, \sigma_\pi, \sigma_R]^T$ collects the parameters characterizing the model. We set $\beta = 0.99$ and $\nu = 1$, a very standard calibration in the literature.[4] The remaining priors, which are standard in this literature, are collected

[4] Perturbations of this baseline calibration confirmed the robustness of our results.

in Table 1. Details on the Bayesian algorithm are relegated in an Appendix available upon request.

Our posterior estimates are reported in Table 1. All the estimated parameters take conventional values. The parameters of the policy rule suggest an aggressive conduct to dampen inflation fluctuations, and a high degree of policy gradualism; the estimated degree of price indexation (posterior mean) is 0.09 (90% credible set: $[0.01, 0.17]$); the estimated weight of the forward looking component in the IS curve is 0.78 (90% credible set: $[0.70, 0.86]$).[5]

3 DSGE vs. CVARs: A MonteCarlo exercise

We now turn to the assessment of the ability of a CVAR to recover the effects of the structural monetary policy shock u_t^R. Basically, we aim at comparing the DSGE-consistent impulse responses to those produced with a VAR whose monetary policy shock is identified, as typically done in the literature, with a Cholesky decomposition of the variance-covariance matrix of a vector autoregression in which "slow-moving" variables such as inflation and output are ordered before the policy rate. Our algorithm works as follows.

For $k = 1$ to K, we

1. sample a realization of the vector $\boldsymbol{\xi}^k$ from the estimated posterior density $p(\boldsymbol{\xi} \,|\, \boldsymbol{Y})$, where $\boldsymbol{Y}$ is the set of observables employed to estimate our model;

2. compute the DSGE model-consistent impulse responses conditional on $\boldsymbol{\xi}^k$ to an unexpected nominal interest rate hike, and store them in the $[3xHxK]$ $\boldsymbol{DSGE\_IRFs}$ matrix, which accounts for the $[3x1]$ vector of variables we focus on, the $h \in \{1, ..., H\}$ step-ahead of the impulse responses of interest, and the $k \in \{1, ..., K\}$ draw of the vector of structural parameters $\boldsymbol{\xi}$;

3. feed the CVAR impulse responses to a normalized monetary policy shock hike with the artificial data $\boldsymbol{x}^k_{ps,[3:T]}$ (ordering: inflation, output gap, nominal rate) generated with the DSGE model conditional on $\boldsymbol{\xi}^k$, and store them in the $[3xHxK]$ $\boldsymbol{CVAR\_IRFs}$ matrix.[6]

[5] A comparison involving the actual series employed to estimate the model and the estimated DSGE's one step ahead predictions confirms that this model is not a "straw man" from an empirical standpoint, in that its short-term forecasting ability is satisfactory. More details are reported in our Appendix.

[6] Given that the DSGE model we deal with features a finite VAR(2) representation, our CVARs are

We run this algorithm by setting the number of repetitions $K = 5,000$, the horizon of the impulse response functions $H = 15$, and the length of the pseudo-data sample $T = 98$. This sample length coincides with that of the actual data sample (1984:I-2008:II) we employed to estimate both our DSGE model and the CVAR whose impulse responses are plotted in Figure 1. Monetary policy shocks are normalized to induce an on-impact equilibrium reaction of the nominal rate equivalent to 25 quarterly basis points.

Figure 3 contrasts the impulse responses obtained with the DSGE model with those generated with our CVARs. This figure is extremely informative. The estimated DSGE predicts a "significant" reaction of both inflation and the output gap, i.e., the zero value does not belong to the 90% credible set for all the realizations of the variables of interest. In particular, the unexpected interest rate hike induces an immediate recession, with the output level getting back to potential after some quarters. Such recession leads to a persistent deflationary phase, which lasts for more than two years. Evidently, our estimated model supports the U.S. monetary policy's ability to affect inflation and the business cycle.

A dramatically different picture arises when turning to our CVARs. On average, our CVARs return muted responses of inflation and output to a monetary policy shock. This result is extremely relevant in light of the popularity of Cholesky-VARs as for the quantification of the effects of a monetary policy shock.

The similarity between the CVAR responses shown in Figure 3 and those reported in Figure 1 is impressive. In both cases, a monetary policy "shock" identified with the Cholesky-scheme induces flat reactions of inflation and output. Therefore, our MonteCarlo evidence suggests that the flat reactions reported in Figure 1, more than a genuine fact, may actually be an *artifact* due to the imposition of the (wrong) Cholesky identification scheme. Therefore, *mild-to-muted CVAR responses* to a (misspecified) monetary policy shock turn out to be *fully consistent* with a monetary policy actually *able to affect* the macroeconomic environment.

Why do we get distorted impulse responses with our CVARs? The fundamental reason is the discrepancy in the *timing assumptions* entertained by the DSGE vs. CVAR models. While the first one allows for an *immediate* impact of the policy shock on

estimated with two lags. Robustness checks dealing with the optimal choice of the VAR lag-length based on the Schwarz criterion supported the solidity of our results. We also verified the robustness of our results to the imposition of DSGE model-consistent matrices to the VAR structure, and to the employment of an upper triangular (as opposed to the lower triangular used here) impulse matrix. These robustness checks are available upon request.

inflation and output, the CVAR imposes a *delayed* reaction. As shown by Carlstrom, Fuerst, and Paustian (2009), in a case like ours one may express the Cholesky-"shocks" $\boldsymbol{\varphi_t}$ in terms of the DSGE shocks $\boldsymbol{u}_t$ as follows:[7]

$$\boldsymbol{\varphi}_t = \Phi \boldsymbol{u}_t = \begin{bmatrix} \phi_{11} & \phi_{12} & \phi_{13} \\ \phi_{21} & \phi_{22} & \phi_{23} \\ \phi_{31} & \phi_{32} & \phi_{33} \end{bmatrix} \begin{bmatrix} u_t^{\pi} \\ u_t^{a} \\ u_t^{R} \end{bmatrix}, \tag{6}$$

Therefore, the mapping going from the true DSGE shocks to the CVAR monetary policy "shock" reads

$$\varphi_t^R = \phi_{31} u_t^{\pi} + \phi_{32} u_t^{a} + \phi_{33} u_t^{R}. \tag{7}$$

The stochastic component φ_t^R is, in fact, a misspecified representation of the true monetary policy shock u_t^R. The standard Cholesky identification scheme would recover the true policy shock only under the restrictions $\phi_{31} = \phi_{32} = 0$. This would occur if the structural DSGE model featured delays in the impact of the true monetary policy shock u_t^R on inflation and output, which is not the case in our DGPs. Then, the effects of the structural monetary policy shocks on inflation and output must be offset by another structural shock. In this model, the only possible candidate is the "technology shock" u_t^a, because it is the only shock able to move inflation and output in the same direction as the monetary policy shock does, therefore offsetting the macroeconomic effects induced by the latter. Therefore, our Cholesky-VARs wrongly return mild-to-muted reactions of inflation and output because they confound the impact of two structural shocks, i.e., a monetary policy shock and a supply shock interpretable as a technology shock.[8]

Given the close similarity between Figures 1 and 3, we conclude that the evidence obtained with actual U.S. data (Figure 1) could very well be an artifact due to the misspecification of the monetary policy shock, more than a true fact.

4 MonteCarlo exercise with Smets and Wouters (2007)

As a matter of fact, virtually all central banks and a large number of researchers have drifted their attention to the richer medium-scale framework à la Smets and Wouters

[7] A full derivation is presented in Carlstrom, Fuerst, and Paustian (2009) and in our Appendix.

[8] Some simulations, confined in our Appendix, confirm that the weight of the cost-push shock in the linear combination (7) is basically zero. Differently, the loading of the technology shock is clearly negative, which is what one should expect from that shock to offset the effects of a monetary policy tightening.

(2007) for some years now. This model features a variety of nominal and real frictions as well as a set of shocks that can be given a structural interpretation. We refer to Smets and Wouters (2007) and to our Appendix for a full description of the model. Clearly, it is of interest to understand if our result carries over when employing such a richer structure as DGP in our MonteCarlo simulations.[9]

We then estimate Smets and Wouters' (2007) framework with Bayesian techniques over the great moderation sample, 1984:I-2008:II. Following Smets and Wouters (2007), we use seven observables (quarterly growth rates of GDP, consumption, investments, and wages, all expressed in per-capita, real terms; log of hours; GDP deflator quarterly inflation; and federal funds rate). The model features a deterministic growth rate driven by labor-augmenting technological progress, so that the data do not need to be detrended before estimation. Tables 2 and 3 document the priors we employed, which are the same as in Smets and Wouters (2007), along with some posterior moments. Basically, our results are in line with most of the literature focusing on the estimation of DSGE models for the U.S. economy with great moderation data. In particular, we find a strong systematic policy reaction to inflation, a mild reaction to the model-consistent output gap, and a slightly stronger one to output growth. Monetary policy is conducted with a fair amount of gradualism. Our evidence points to a fairly large degree of habit formation in consumption, and lends support to the modeling of frictions in capital formation. The posterior means of the Calvo price and wage parameters are comparable with a large number of estimates obtained with macroeconomic U.S. data. Shocks to TFP, Government spending, price and wage mark-ups feature a high degree of correlation, also considering the MA(1) component of these last two shocks.

Given that this model fits the growth rate of real GDP, it is of interest to understand if the results shown in Figure 1, obtained with an empirical measure of the output gap, still hold when modeling a CVAR with actual U.S. data, 1984:I-2008:II, involving inflation, output *growth*, and the federal funds rate. Figure 4 depicts the outcome of our VAR regressions. The mean reactions of inflation and output are clearly in line with conventional wisdom. However, the reaction of inflation is surrounded by a substantial amount of uncertainty. The reaction of output growth is more precisely estimated, with the 68% confidence set signalling negative values for a few quarters beginning one year

[9] Medium-scale models like Smets and Wouters' (2007) are also affected by misspecification, as found by Del Negro, Schorfheide, Smets, and Wouters (2007). However, as stressed by Del Negro, Schorfheide, Smets, and Wouters (2007) themselves, such model's impulse responses to a monetary policy shock are statistically comparable to those produced by the best fitting version of its DSGE-VAR "counterpart". This result supports the use of the Smets and Wouters (2007) model for our exercise.

after the shock. According to the 90% confidence set, however, the reaction is not significant.

The outcome of our MonteCarlo exercise based on the Smets and Wouters (2007) model is depicted in Figure 5. Three considerations are in order. First, the responses of inflation and output estimated with our Cholesky-VARs are substantially distorted. While the structural DSGE model suggests a negative and persistent reaction of inflation and output to a policy tightening, our VARs predict mild reactions surrounded by a large amount of uncertainty. Therefore, the MonteCarlo exercise based on the Smets and Wouters (2007) model confirms that Cholesky-VARs are likely to produce substantially distorted macroeconomic reactions to a policy surprise under the null hypothesis of contemporaneous timing.[10] Second, our simulations return Cholesky-VAR responses very similar to those obtained with actual post-Volcker U.S. data. A noticeable discrepancy between Figures 4 and 5 regards the reaction of output growth, which appears slightly more precisely estimated with actual data than in our MonteCarlo simulations. However, the match between these two Figures is evident. This leads us to our third consideration, i.e., also an exercise conducted with a medium-scale model à la Smets and Wouters (2007) suggest that mild reactions of inflation and output to a policy shock are likely to be an artifact induced by the employment of the Cholesky-identification scheme, more than a fact.

The MonteCarlo experiment based on the Smets and Wouters (2007) model involves a number of possible reasons behind the failure of Cholesky-VARs to recover the effects of a structural policy shocks. On top of the assumption on delayed effects, our VARs are also likely to be affected by truncation biases and non-fundamentalness issues. Truncation biases may arise due to the VARMA representation of the Smets and Wouters (2007) framework, which is driven first of all by the presence of ARMA(1,1) price and wage mark-up shocks in the system. This implies a theoretical VAR(∞) representation of the DSGE model, which can clearly suffer from truncation bias-issues. Non-fundamentalness is likely to arise due to the omission of relevant factors from the VARs, a notable one being potential output. As a matter of fact, it is complicated to correctly identify and quantify the relative role played by non-monetary policy struc-

[10] The policy rule in the Smets and Wouters (2007) model features a systematic reactions of the policy rate to current inflation, the output gap, and output growth. In our MonteCarlo exercise, which assumes the Smets and Wouters (2007) model to be the DGP, CVARs do not feature any measure of the output gap. However, an estimated version of the Smets and Wouters (2007) model featuring a systematic policy reaction to inflation and output growth only leaves our MonteCarlo results unchanged (see our Appendix).

tural shocks vs. truncation bias vs. omitted relevant factors in this analysis. However, a clear story is told by this exercise, i.e., a Cholesky-VARs may very well confound a monetary policy able to affect the economic system with monetary policy ineffectiveness.

5 Relation to the literature

The papers closest to ours are probably Canova and Pina (2005) and Carlstrom, Fuerst, and Paustian (2009). Canova and Pina (2005) set up a Monte Carlo exercise in which they consider two calibrated small-scale DSGE models (a limited participation model and a sticky price-sticky wage economy) as DGPs to estimate a variety of short-run "zero restrictions" VAR identification schemes. They find substantial differences between the predictions coming from the structural models and those implied by the estimated CVARs. Carlstrom, Fuerst, and Paustian (2009) propose a theoretical investigation on the consequences of the timing discrepancy between DSGE and CVARs as for the macroeconomic reactions to a monetary policy shock. They show that, depending on the chosen calibration of their DSGE models, CVARs may return a variety of predictions, including price and output puzzles, responses in line with the true DSGE reactions, muted responses, and so on. These papers make a theoretical point. Our contribution is empirical, in that we employ two *estimated* models of the business cycle to *measure* the ability of Cholesky-VARs to identify a U.S. monetary policy shock. On top of it, we draw a comparison between CVARs' responses estimated with artificial vs. actual U.S. data, from which we are able to offer a novel interpretation of the mild-to-muted macroeconomic reactions to a monetary policy shock typically found when focusing on the post-Volcker sample.

Fernández-Villaverde and Rubio-Ramírez (2006) and Fernández-Villaverde, Rubio-Ramírez, Sargent, and Watson (2007) derive a necessary condition to ensure the existence of the VAR representation of a DSGE model (i.e. to check if the DSGE model is "invertible").[11] Ravenna (2007) discusses under which conditions a finite VAR representation exists, and shows that truncated VARs may provide misleading indications when the true DGP is an infinite order VAR. Further investigations on the distortions coming from the truncation bias, in the context of the identification of the effects of technology shocks on hours worked, are offered by Christiano, Eichenbaum, and Vig-

[11] A VAR is invertible if its innovations map into the shocks of the economic model in population and under the correct identification scheme. Non-invertibilities typically arise when some relevant state variables of the model are not included in the VAR (for instance, because they are not observable). The relevance of non-invertibility is, of course, an empirical issue - see e.g. Sims (2009).

fusson (2006) and Chari, Kehoe, and McGrattan (2008). Our focus is different, in that we are interested in understanding to what extent the VAR evidence on the effects of monetary policy shocks during the post-Volcker era is driven by the choice of the identification scheme *per se*. Moreover, the small-scale model we use has an exact VAR(2) representation, which (in principle) allows the researcher to correctly recover the effects of a structural monetary policy shocks. Differently, the Smets and Wouters (2007) model we use has a VARMA representation, which leads to a VAR representation with infinite lags. This does not invalidate our exercise, however, in that our ultimate goal is to mimic the behavior of an econometrician endowed with some time series and willing to study the effects of a policy shock with a Cholesky-VAR. Hence, a model like Smets and Wouters's (2007) is clearly suited for our MonteCarlo exercise.

Del Negro and Schorfheide (2004) exploit the structure of a DSGE model in a data-driven fashion, in that the fit of a Bayesian VAR is maximized by opportunely relaxing the restrictions imposed by the DSGE framework on such VAR representation. The authors note that their methodology enables the econometrician to identify structural shocks with the VAR in a non-recursive fashion. An application of their methodology to a medium-scale model à la Smets and Wouters (2007) is provided by Del Negro, Schorfheide, Smets, and Wouters (2007), who find the Cholesky restrictions to be implausible due to the very likely immediate reaction of output to a policy shock. These papers approximate the DSGE model by a VAR, then they systematically relax the implied cross-equation restrictions and document how the model fit changes. Differently, as it is customary in MonteCarlo exercises, we assume our (DSGE) models to be "true". Then, conditional on this assumption, we conduct a MonteCarlo experiment to assess the ability of Cholesky VARs to replicate the impulse response stemming from our DGPs. Faust, Swanson, and Wright (2004) show that the zero response of prices to a monetary policy shock imposed by a standard Cholesky-identification scheme is not supported by the data when disturbances are inferred using futures data in a two-step procedure. That paper deals, however, with the issue of identification schemes within structural VARs (for which the authors provide econometric testing), but it is silent on structural models.

6 Conclusions

This paper shows that *mild-to-muted* impulse responses produced with a Cholesky-VAR estimated with U.S. 1984:I-2008:II data are *fully consistent* with *monetary policy shocks* *exerting substantial effects on inflation and output.* We make this point by proceeding in two steps. Firstly, we estimate two new-Keynesian DSGE frameworks with Bayesian techniques, and verify that they predict negative, persistent reactions of inflation and output to an unexpected monetary policy tightening. Then, we set up a MonteCarlo experiment in which we feed Cholesky-VARs with pseudo-data generated by our estimated new-Keynesian frameworks in turn. We show that Cholesky-VARs generate, on average, falsely mild-to-muted responses of inflation and output. A misspecification of the policy shock due to the timing discrepancy existing between the structural DSGE model (that allows for an immediate impact of the policy shock on inflation and output) and the Cholesky-VAR (that models a transmission lag from the policy shock to inflation and output) is shown to be the driver of this result. Because of this timing discrepancy, Cholesky-VARs' monetary policy "shocks" are, in fact, a convolution of true structural shocks exerting offsetting effects on our macroeconomic indicators.

Which are the implications of our study? To be clear, our results do *not* call for a rejection of the VAR approach. Vector autoregressions are clearly useful to establish stylized facts when different, competing models are *a-priori* equally sensible. As Fernandéz-Villaverde, Rubio-Ramírez, Sargent, and Watson (2007, page 1025) puts it, "Despite pitfalls, it is easy to sympathize with the enterprise of identifying economic shocks from VAR innovations if one is not dogmatic in favor of a particular fully specified model." However, our results suggest that i) the evidence on the macroeconomic reactions to a monetary policy shock identified with a standard recursive scheme should be interpreted with great care; ii) alternative identification schemes should be adopted - for a recent survey discussing pros and cons of alternatives to the recursive identification scheme, see Kilian (2011). DSGE models are likely to be misspecified. But, as shown by Del Negro and Schorfheide (2004) and Del Negro, Schorfheide, Smets, and Wouters (2007), also misspecified DSGE models, when combined with VARs, may provide useful information to estimate the contemporaneous and dynamic effects of a monetary policy shock. We see their proposal as a promising alternative to Cholesky-VARs.

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Param.	*Interpretation*	*Priors*	*Posterior Means* $_{[5h,95th]}$
β	Discount factor	*Calibrated*	0.99 [−]
v^{-1}	Frisch elasticity	*Calibrated*	1 [−]
κ	NKPC, slope	$Normal(0.1, 0.015)$	0.12 [0.10,0.14]
α	Price indexation	$Beta(0.5, 0.2)$	0.09 [0.01,0.17]
γ	IS, forw. look. degree	$Beta(0.5, 0.2)$	0.78 [0.70,0.86]
σ	Inverse of the IES	$Normal(3, 1)$	5.19 [3.95,6.45]
τ_π	T. Rule, inflation	$Normal(1.5, 0.3)$	2.21 [1.85,2.56]
τ_y	T. Rule, output gap	$Gamma(0.3, 0.2)$	0.16 [0.05,0.25]
τ_R	T. Rule, inertia	$Beta(0.5, 0.285)$	0.81 [0.77,0.86]
ρ_a	AR tech. shock	$Beta(0.5, 0.285)$	0.89 [0.84,0.94]
ρ_π	AR cost-push shock	$Beta(0.5, 0.285)$	0.98 [0.97,0.99]
ρ_R	AR mon. pol. shock	$Beta(0.5, 0.285)$	0.43 [0.30,0.56]
σ_a	Std. tech. shock	$InvGamma(1.5, 0.2)$	1.50 [1.10,1.91]
σ_π	Std. cost-push. shock	$InvGamma(0.35, 0.2)$	0.09 [0.07,0.11]
σ_R	Std. mon. pol. shock	$InvGamma(0.35, 0.2)$	0.14 [0.12,0.15]

Table 1: **Bayesian estimates of the small-scale DSGE model.** 1984:I-2008:II U.S. data. 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.

Param.	*Interpretation*	*Priors*	*Posterior Means* [5h,95th]
φ	Capital adj. cost elasticity	$Normal(4, 1.5)$	6.06 [4.22,7.96]
σ_c	Risk aversion	$Normal(1.5, 0.375)$	1.39 [1.16,1.62]
h	Habit formation	$Beta(0.7, 0.1)$	0.63 [0.50,0.75]
ξ_w	Wage stickiness	$Beta(0.5, 0.1)$	0.64 [0.49,0.79]
σ_l	Elast. lab. supply	$Normal(2, 0.75)$	1.76 [0.78,2.74]
ξ_p	Price stickiness	$Beta(0.5, 0.1)$	0.71 [0.62,0.80]
ι_w	Wage indexation	$Beta(0.5, 0.15)$	0.52 [0.28,0.76]
ι_p	Price indexation	$Beta(0.5, 0.15)$	0.40 [0.20,0.59]
ψ	Capacity utiliz. elast.	$Beta(0.5, 0.15)$	0.69 [0.54,0.85]
$\Phi - 1$	Fixed c. in prod. (share)	$Normal(0.25, 0.125)$	0.44 [0.30,0.57]
r_π	T. Rule, inflation	$Normal(1.5, 0.25)$	2.10 [1.78,2.43]
ρ	T. Rule, inertia	$Beta(0.75, 0.10)$	0.83 [0.80,0.87]
r_y	T. Rule, output gap	$Normal(0.125, 0.05)$	0.05 [0.02,0.09]
$r_{\Delta y}$	T. Rule, output growth	$Normal(0.125, 0.05)$	0.16 [0.11,0.20]
$\overline{\pi}$	St. state inflation rate	$Gamma(0.625, 0.10)$	0.64 [0.55,0.73]
$100(\beta^{-1} - 1)$	St. state interest rate	$Gamma(0.25, 0.10)$	0.25 [0.10,0.40]
$\bar{l}$	St. state hours worked	$Normal(0, 2)$	0.87 [−0.84,2.57]
$\overline{\gamma}$	Trend growth rate	$Normal(0.4, 0.1)$	0.42 [0.37,0.47]
α	Share of capital in prod.	$Normal(0.3, 0.05)$	0.32 [0.25,0.39]

Table 2: **Bayesian estimates of the Smets and Wouters' (2007) DSGE model - Structural Parameters.** 1984:I-2008:II U.S. data. 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.

Param.	*Interpretation*	*Priors*	*Posterior Means* [$5h$,$95th$]
σ_a	TFP shock, st.dev.	$InvGamma(0.1, 2)$	0.41 [0.36,0.46]
σ_b	Risk-premium shock, st.dev.	$InvGamma(0.1, 2)$	0.16 [0.10,0.21]
σ_g	Gov. spending shock, st.dev.	$InvGamma(0.1, 2)$	0.41 [0.36,0.46]
σ_I	Invest.-specific tech. shock, st.dev.	$InvGamma(0.1, 2)$	0.35 [0.27,0.42]
σ_r	Mon. policy shock, st.dev.	$InvGamma(0.1, 2)$	0.12 [0.10,0.14]
σ_p	Price mark-up shock, st.dev.	$InvGamma(0.1, 2)$	0.10 [0.08,0.12]
σ_w	Wage mark-up shock, st.dev.	$InvGamma(0.1, 2)$	0.29 [0.23,0.35]
ρ_a	TFP shock, AR(1) coeff.	$Beta(0.5, 0.2)$	0.95 [0.92,0.97]
ρ_b	Risk-premium shock, AR(1) coeff.	$Beta(0.5, 0.2)$	0.32 [0.04,0.62]
ρ_g	Gov. sp. shock, AR(1) coeff.	$Beta(0.5, 0.2)$	0.95 [0.93,0.97]
ρ_I	Invest.-spec. tech. shock, AR(1) coeff.	$Beta(0.5, 0.2)$	0.74 [0.65,0.85]
ρ_r	Mon. pol. shock, AR(1) coeff.	$Beta(0.5, 0.2)$	0.30 [0.15,0.44]
ρ_p	Price mark-up shock., AR(1) coeff.	$Beta(0.5, 0.2)$	0.89 [0.81,0.98]
ρ_w	Wage mark-up shock, AR(1) coeff.	$Beta(0.5, 0.2)$	0.93 [0.88,0.98]
μ_p	Price mark-up shock, MA(1) coeff.	$Beta(0.5, 0.2)$	0.66 [0.47,0.85]
μ_w	Wage mark-up shock, MA(1) coeff.	$Beta(0.5, 0.2)$	0.71 [0.53,0.88]
ρ_{ga}	Gov.spending-TFP shocks, correlation	$Beta(0.5, 0.2)$	0.44 [0.28,0.61]

Table 3: **Bayesian estimates of the Smets and Wouters' (2007) DSGE model - Shock processes.** 1984:I-2008:II U.S. data. 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.

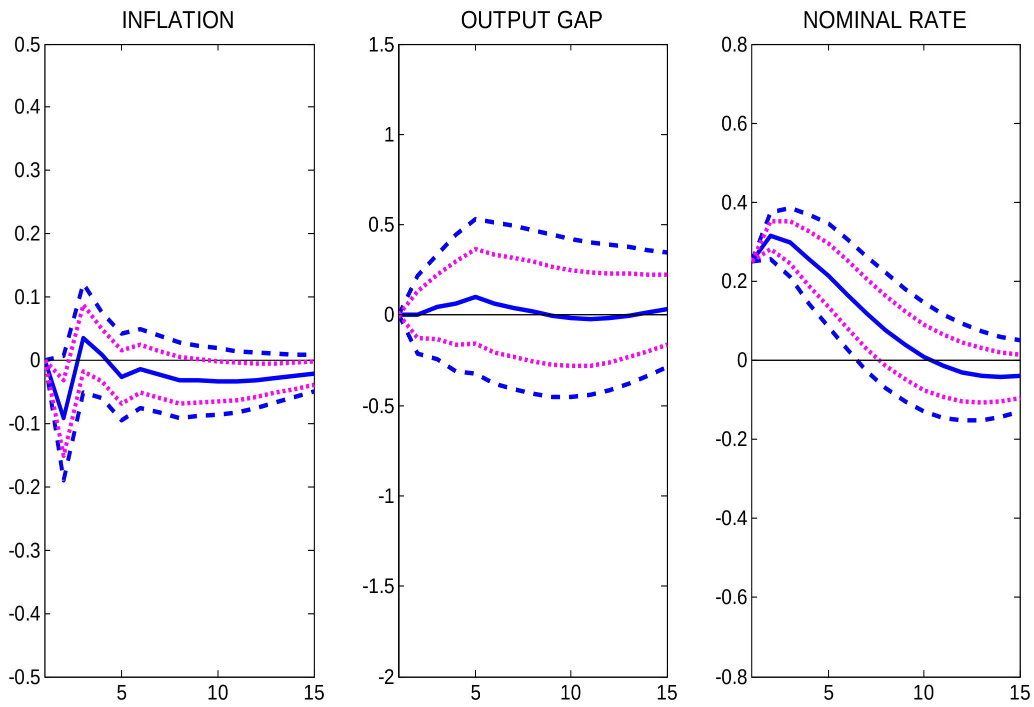

Figure 1: **CVAR impulse response functions to a monetary policy shock.** Sample: 1984:I-2008:II. Variables: Quarterly GDP inflation, CBO output gap, quarterly federal funds rate - source: FREDII. Identification of the monetary policy shock via Cholesky decomposition (lower triangular matrix, ordering: inflation, output gap, federal funds rate). Solid blue line: Mean response; Dashed blue lines: [5th,95th] percentiles; Magenta dotted lines: [16th,84th] percentiles (bootstrapped, 500 repetitions). VAR estimated with a constant, a linear trend, and three lags.

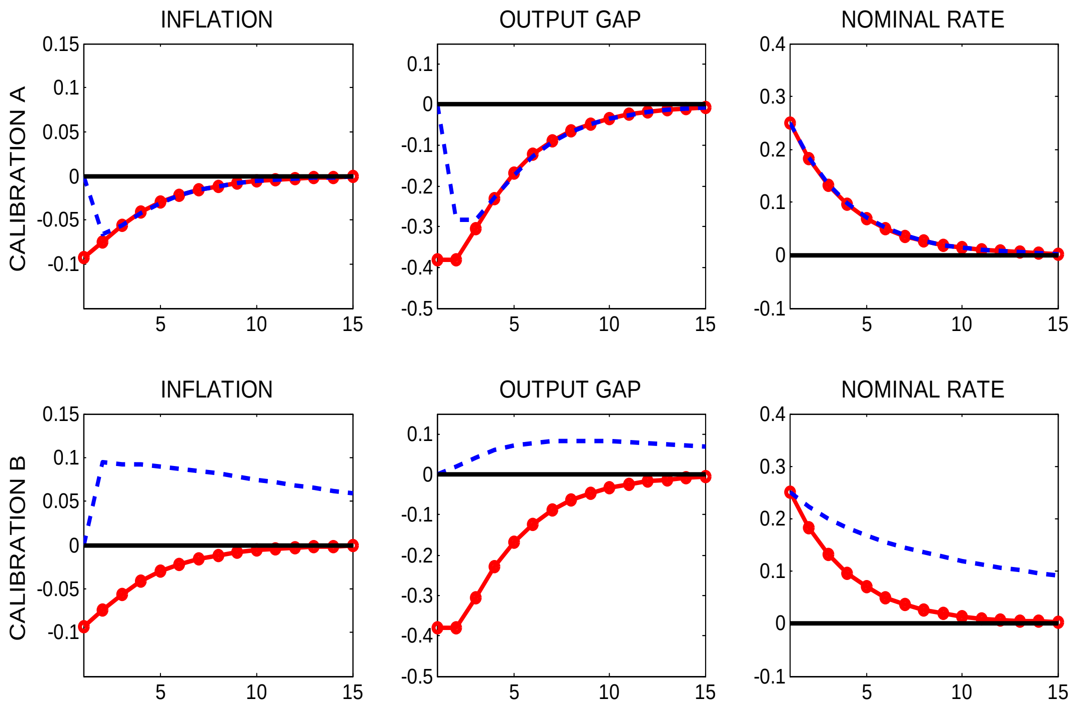

Figure 2: **DSGE vs. CVARs impulse response functions to a monetary policy shock - the role of calibration.** Solid red lines: DSGE impuse responses. Dashed blue lines: CVAR impulse responses (population moments). Identification of the monetary policy shock via Cholesky decomposition (lower triangular matrix, ordering: inflation, output gap, short-term interest rate). VAR estimated with two-lags. Model calibration, common values: $\beta = 0.99$, $\upsilon = 1$, $\kappa = 0.05$, $\alpha = 0$, $\gamma = 1$, $\sigma = 4$, $\tau_\pi = 1.2$, $\tau_y = 0$, $\tau_R = 0.8$, $\rho_\pi = \rho_R = 0$, $\sigma_a = 4$, $\sigma_\pi = 0.8$, $\sigma_R = 0.3$. Calibration A: $\rho_a = 0$; Calibration B: $\rho_a = 0.95$.

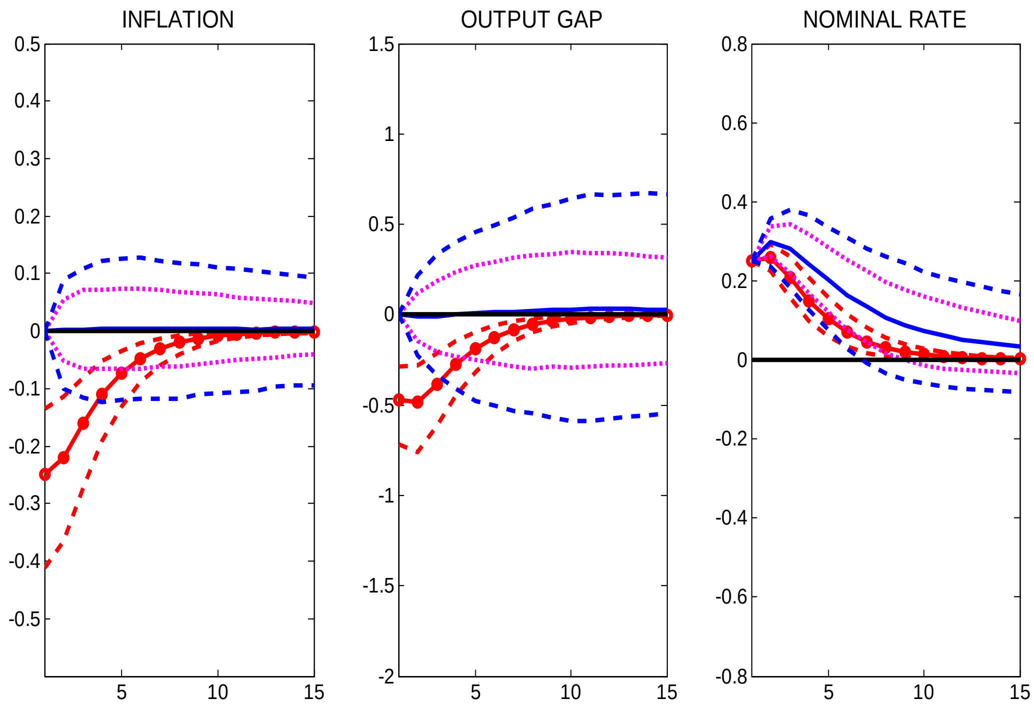

Figure 3: **Small-scale DSGE vs. CVAR impulse response functions to a monetary policy shock.** Circled red lines: DSGE Bayesian mean impulse responses; Dashed red lines: 90% credible sets. Solid blue line: CVAR mean impulse responses; Dashed blue lines: [5th,95th] percentiles; Magenta dotted lines: [16th,84th] percentiles. Moments computed the impulse response function distributions simulated by drawing 5,000 realizations of the vector of parameters of the DSGE model, which is also used to generate the pseudo-data to feed the CVARs. Identification of the monetary policy shock via Cholesky decomposition (lower triangular matrix, ordering: inflation, output gap, nominal rate). VAR estimated with two lags.

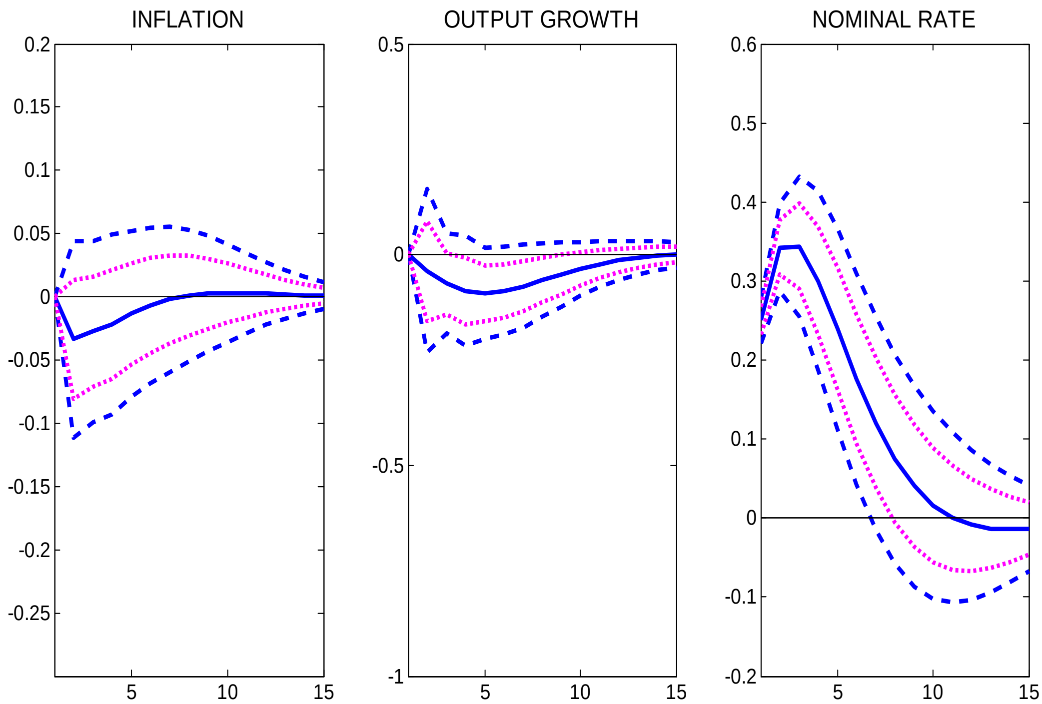

Figure 4: **CVAR impulse response functions to a monetary policy shock, 1984:I-2008:II**. Variables: Quarterly GDP inflation, quarterly output growth, quarterly federal funds rate - source: FREDII. Identification of the monetary policy shock via Cholesky decomposition (lower triangular matrix, ordering: inflation, output growth, federal funds rate). Solid blue line: Mean response; Dashed blue lines: [5th,95th] percentiles; Magenta dotted lines: [16th,84th] percentiles (bootstrapped, 500 repetitions). VAR estimated with a constant, a linear trend, and two lags.

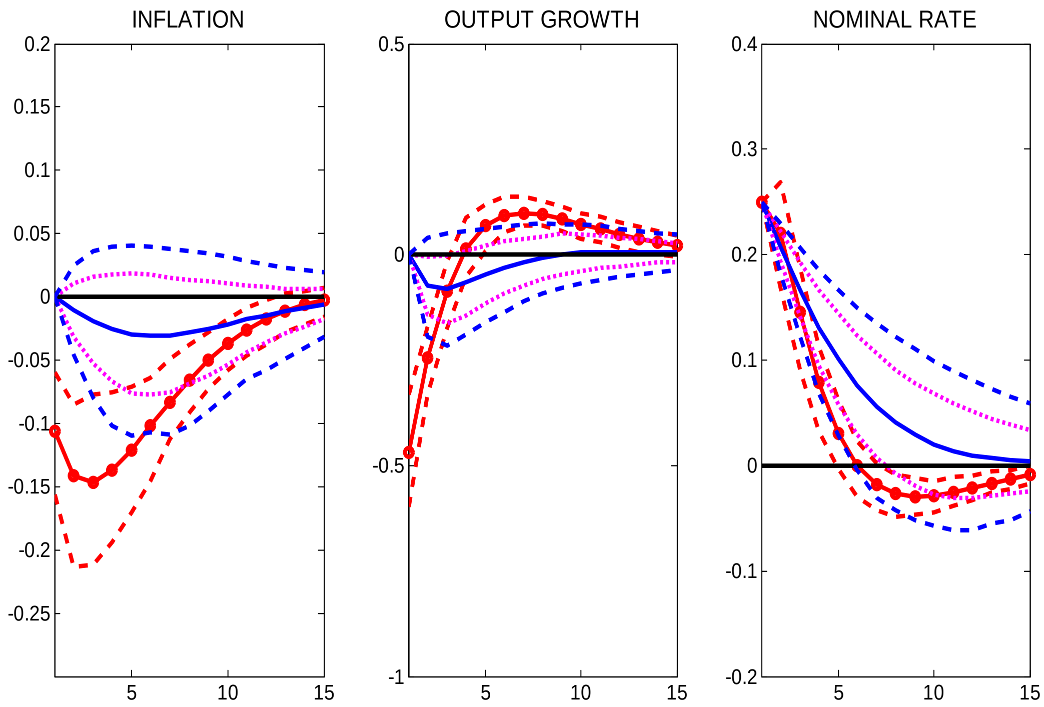

Figure 5: **DSGE à la Smets and Wouters (2007) vs. CVAR impulse response functions to a monetary policy shock.** Circled red lines: DSGE Bayesian mean impulse responses; Dashed red lines: 90% credible sets. Solid blue line: CVAR mean impulse responses; Dashed blue lines: [5th,95th] percentiles; Magenta dotted lines: [16th,84th] percentiles. Moments computed the impulse response function distributions simulated by drawing 5,000 realizations of the vector of parameters of the DSGE model, which is also used to generate the pseudo-data to feed the CVARs. Identification of the monetary policy shock via Cholesky decomposition (lower triangular matrix, ordering: inflation, output growth, nominal rate). VAR estimated with a number of lags determined (per each given VAR) by the Schwarz criterion.