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

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CEA@Cass Working Paper Series

WP–CEA–03-2011

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

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October 2011

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 pointing to those responses being substantially distorted is provided. The wrongly assumed policy delays consistent with the Cholesky-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 for the mild-to-muted reactions of inflation and output in the post-Volcker era.

JEL classification: C22, E47, E52.

Keywords: Monetary policy shocks, Cholesky identification, VARs, Dynamic Stochastic General Equilibrium models, MonteCarlo simulations.

*We thank Michal Andrle, Guido Ascari, Fabio Canova, Chuck Carlstrom, Martin Eichenbaum, Martin Ellison, Luca Fanelli, Massimo Franchi, Eric Leeper, Salvatore Nisticò, Matthias Paustian, Harald Uhlig, and participants at the 10th Econometric Society World Congress (Shanghai), the 5th BOMOPA Economics Meetings (Padua), the 6th Dynare Conference (Helsinki), and seminars held at the National Bank of Slovakia, the Universities of Oxford, Milan, Milano 'Bicocca', Milano 'Cattolica', Rome 'La Sapienza', Namur, Bologna, and Verona for comments and suggestions. Part of this paper was written while visiting the Department of Economics of the University of Oxford, whose hospitality is gratefully acknowledged. All remaining errors are ours. Paper previously circulated with the title "Monetary Policy Shocks, Cholesky Identification and DSGE Models: An Empirical Investigation for the U.S.". Author's details: Efrem Castelnuovo, Department of Economics, University of Padova, Via del Santo 33, I-35123 Padova (PD). E-mail: efrem.castelnuovo@unipd.it. Phone: +39 049 8274257. Fax: +39 049 8274211.

1 Introduction

Vector AutoRegressions (VARs) have been employed to estimate the effects of policy shocks on output and inflation for at least three decades. 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 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), and Barakchian and Crowe (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

[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. Bagliano and Favero (1998) find signs of misspecification affecting VAR models estimated over long periods involving the 1970s jointly with the great moderation sample. Castelnuovo (2011) finds real balances to be important to replicate the volatility of output in the 1970s, but not during the great moderation. 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.

interpretation involves the role of financial 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 standard new-Keynesian model 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 predicts 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 *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 (alternatively) considering two different DSGE models of the business cycle. The first one is a standard small-scale new-Keynesian model à la King (2000) and Woodford (2003), which features three equations responsible

[3] "Econometric" intepretations involve small-sample bias issues, which might be severe in a sample like the 1980s-2000s as for the U.S. economy, 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.

for the evolution of inflation, output, and the policy rate. Despite of its simplicity, this model 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 will 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. Therefore, we conclude that evidence of mild-to-muted macroeconomic reactions to policy shocks obtained with Cholesky-VARs may very well be consistent with monetary policy shocks being truly effective in affecting the macroeconomic environment. 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. This paper represents a further step in this research agenda, in that it *empirically* assesses Cholesky-VARs's ability to identify the effects of policy surprises. As anticipated, our results point towards a very poor performance by Cholesky VARs along this dimension.[4]

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 DSGE with those coming from the SVARs in a controlled environment, and show that substantial differences arise. An interpretation of our results, highligthing the role of non-policy structural shocks for the Cholesky-VAR responses, is also provided. Section 5 presents our analysis based on the medium-scale model by Smets and Wouters (2007), which confirms our results. Section 6 relates this work to some contributions in the literature. Section 7 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:

$$\pi_t = (1+\alpha\beta)^{-1}[\beta E_t \pi_{t+1} + \alpha \pi_{t-1} + \kappa y_t + \varepsilon_t^{\pi}], \quad (1)$$

$$y_t = \gamma E_t y_{t+1} + (1-\gamma) y_{t-1} - \sigma^{-1}(R_t - E_t \pi_{t+1}) + Q(\rho_a - 1) a_t, \quad (2)$$

$$R_t = (1-\tau_R)(\tau_\pi \pi_t + \tau_y y_t) + \tau_R R_{t-1} + \varepsilon_t^R, \quad (3)$$

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 intepreted as a "cost-push shock"; γ is the weight of the forward-looking component in the intertemporal IS curve (2); σ^{-1} is the

[4]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 certainly uncover the true impulse response functions from a monetary policy shock. Our paper assumes 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). To the best of our knowledge, no investigation on the relevance of modeling the transmission delays in a DSGE framework from an empirical standpoint has been conducted so far. We leave this important research question to future research.

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 "technological 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:

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

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.[5] The remaining priors are collected in Table 1. Notice that such priors are fairly uninformative, above all as regards the autoregressive parameters, which are important drivers of the possible biases arising when imposing the (wrong) Cholesky-factorization to identify the monetary policy shock (Carlstrom, Fuerst, and Paustian (2009)). 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: $[0.70, 0.86]$).[6]

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 with 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

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

[6] 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.

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 $\boldsymbol{x}_{ps,[3:T]}^k$ (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.[7]

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 2 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.

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

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 2 and those reported in Figure 1 is impressive. In both cases, a "monetary policy shock" identified with the Cholesky recursive 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 SVAR 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 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:[8]

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

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 mistakingly return mild-

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

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 intepretable as a technology shock.[9]

Given the close similarity between Figures 1 and 2, 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 (2007). 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.

We then estimate Smets and Wouters' (2007) framework with 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. We employ Bayesian techniques and assume their prior distributions as for the parameters of the estimated model. For the same reasons already expressed in Section 2.2, we condition our analysis on the great moderation sample 1984:I-2008:II. Then, by taking the estimated version of the Smets and Wouters (2007) model as our DGP, we conduct our MonteCarlo exercise as explained in Section 3. We document the results of our Bayesian estimation in our Appendix [to be added]. It is worth noticing, however, that they are in line with those provided by most of the literature.

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 vector of actual U.S. data, 1985:I-2008:II, involving inflation, output *growth*, and the federal funds rate. Figure 3 depics the outcome of

[9] Some simulations, available upon request, 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.

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 a substantial amound 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 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 4. 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, also an exercise based on the Smets and Wouters (2007) model as DGP suggests 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 MonteCarlo exercises return Cholesky-VAR responses very simular to those obtained with actual post-Volcker U.S. data. A noticeable discrepancy between Figures 3 and 4 regard 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-to-muted 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 of 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(∞) rep-

[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, VARs 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).

resentation 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 structural 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) 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 *quantify* 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 rep-

[11] A VAR is invertible if its innovations do map into the shocks of the economic model even 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).

resentation 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 Vigfusson (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) propose a methodology to exploit the restrictions implied by a DSGE model of the business cycle to educate the estimation of VAR coefficients and identify structural shocks 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 withing 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 significantly 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 with our estimated new-Keynesian frameworks. 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. Differently, we show that an identification based on sign restrictions recovers the qualitative patters of the macroeconomic reactions to a policy shock successfully.

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 these alternative, 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.

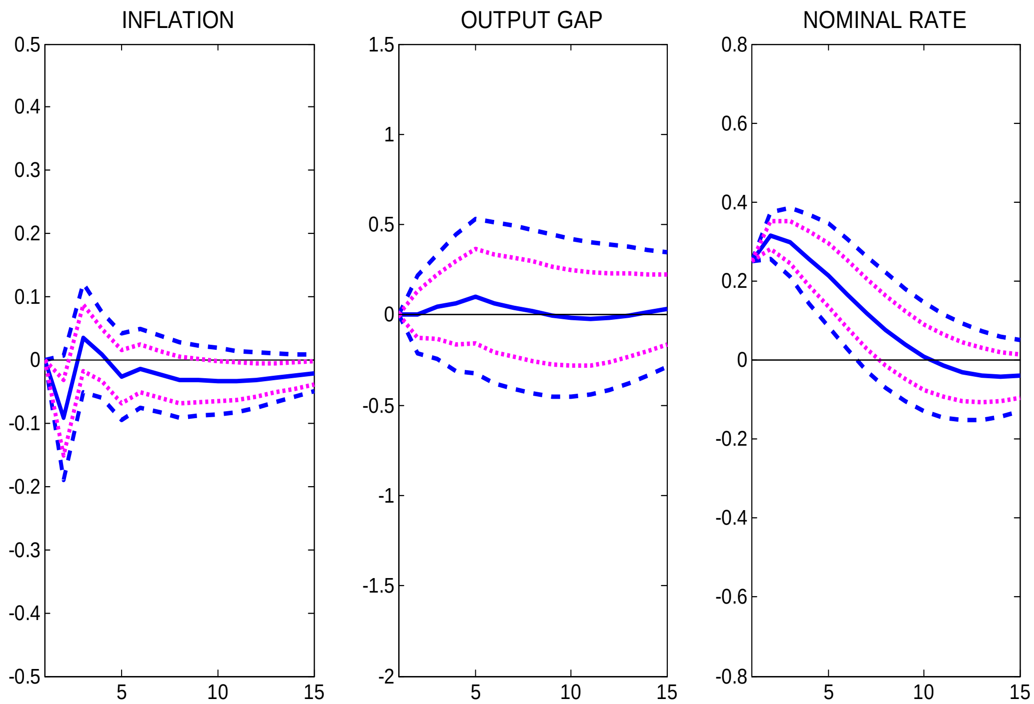

Figure 1: **CVAR impulse response functions to a monetary policy shock.** Sample: 1984:I-2008:II. Variables: Quarterly GDP inflation, CBO output gap, quarterly federal funds rate - source: FREDII. Identification of the monetary policy shock via Cholesky decomposition (lower triangular matrix, ordering: 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.

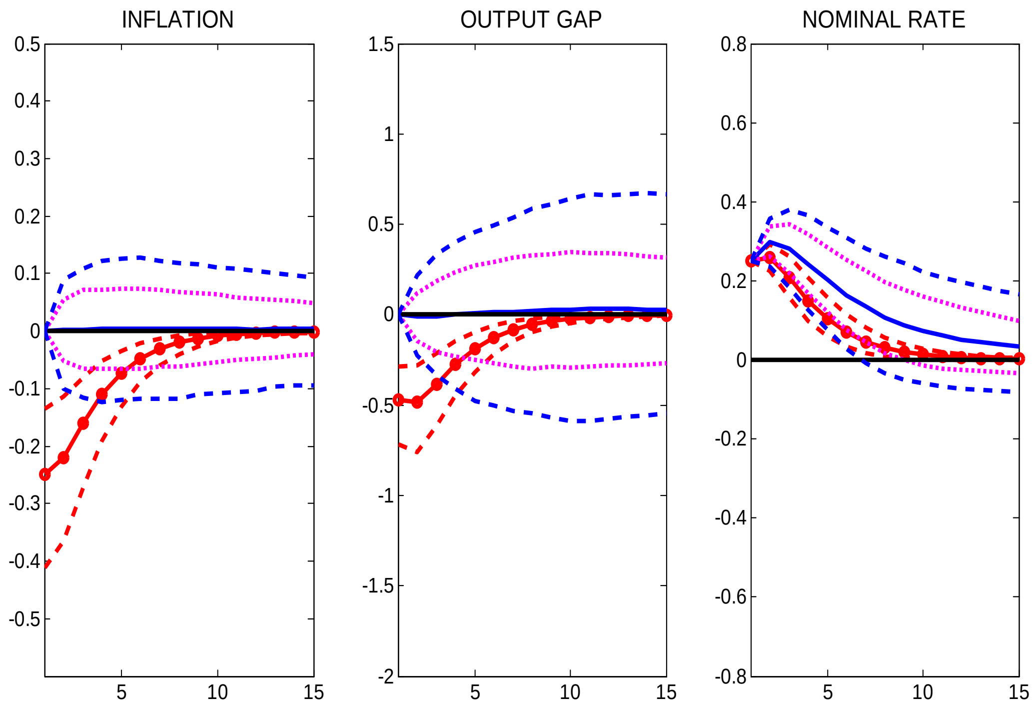

Figure 2: **Small-scale DSGE vs. CVAR impulse response functions to a monetary policy shock.** Circled red lines: DSGE Bayesian mean impulse responses; Dashed red lines: 90% credible sets. Solid blue line: CVAR mean impulse responses; Dashed blue lines: [5th,95th] percentiles; Magenta dotted lines: [16th,84th] 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.

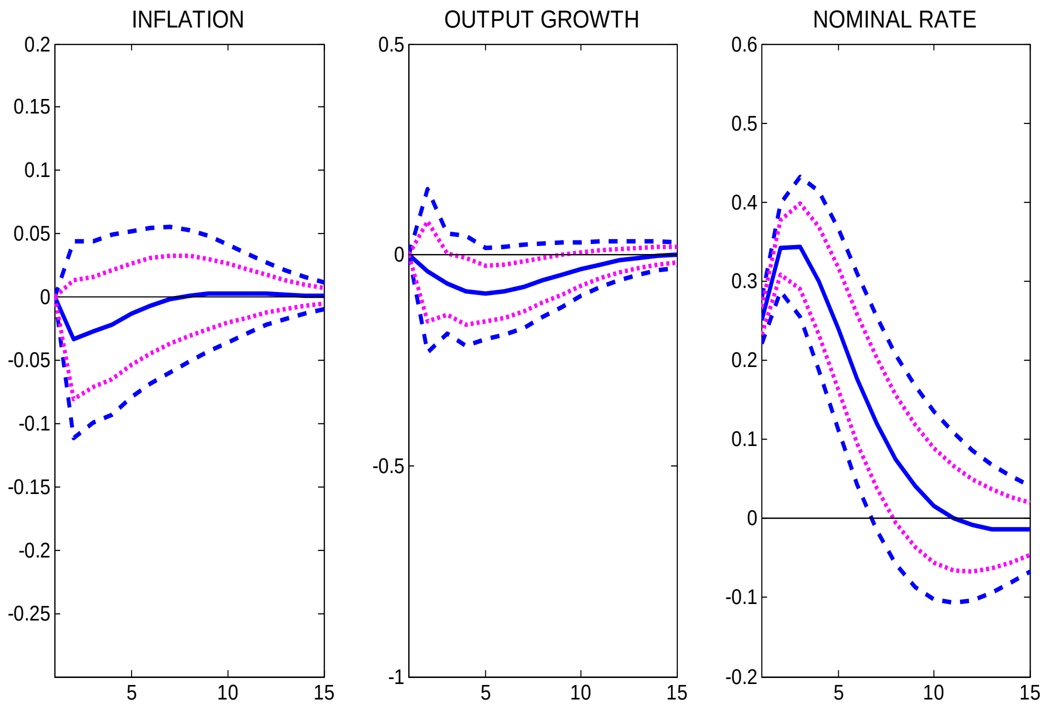

Figure 3: **CVAR impulse response functions to a monetary policy shock, 1984:I-2008:II**. Variables: Quarterly GDP inflation, quarterly output growth, quarterly federal funds rate - source: FREDII. Identification of the monetary policy shock via Cholesky decomposition (lower triangular matrix, ordering: 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.

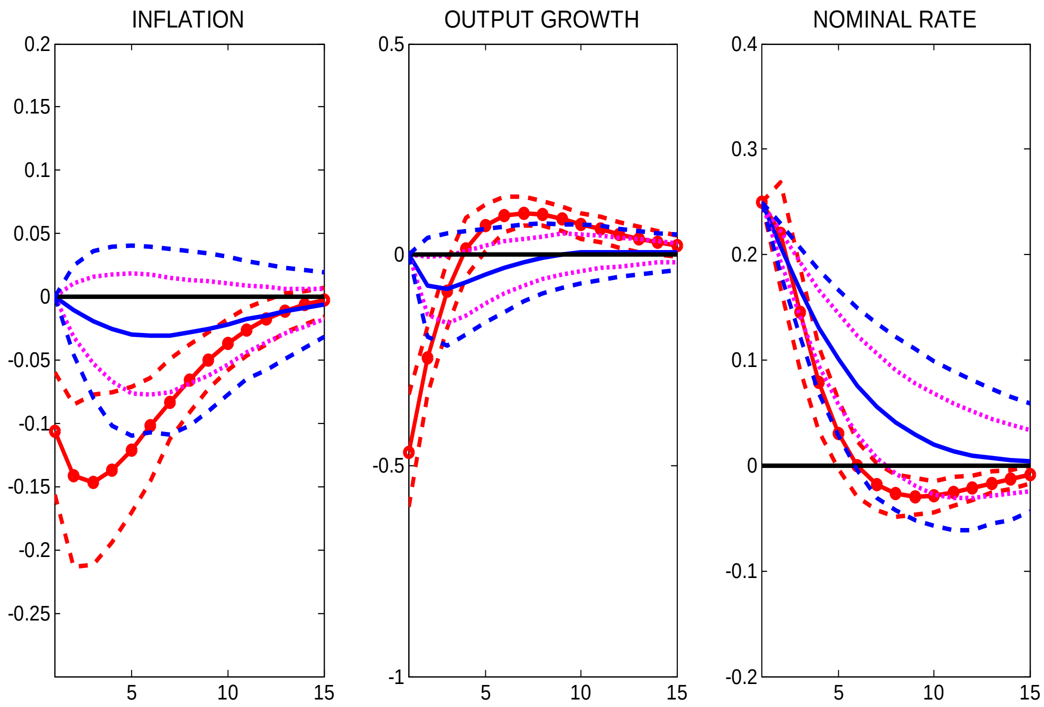

Figure 4: **DSGE à la Smets and Wouters (2007) vs. CVAR impulse response functions to a monetary policy shock.** Circled red lines: DSGE Bayesian mean impulse responses; Dashed red lines: 90% credible sets. Solid blue line: CVAR mean impulse responses; Dashed blue lines: [5th,95th] percentiles; Magenta 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.

Appendix of the paper "In Cholesky-VARs We Trust? An Empirical Investigation with U.S. Data"

1 CVAR estimated with actual U.S. data, 1954:III-2008:II

The evidence shown in the text is clearly not in line with the one provided by a variety of authors as for the effects of a monetary policy shock in the entire post-WWII sample (Christiano, Eichenbaum, and Evans (1999), Christiano, Eichenbaum, and Evans (2005)). This Section shows that the discrepancy between our results in Figures 1 and 3 in the paper, based on great moderation data, and those obtained with longer samples is mainly driven by the observations belonging to the 1970s. Figures A1 and A2 documents our CVAR evidence obtained with the sample 1954:III-2008:II with two different trivariate VARs, one featuring the CBO measure of the output gap and the other one modeling quarterly output growth. Evidently, the reactions of inflation and output are significant and quite different with respect to the mild-to-muted ones shown in Figures 1 and 3 of the paper. In particular, inflation reacts positively and significanly to a monetary policy tightening, so confirming the existence of a "price puzzle" as already documented by, among others, Boivin and Giannoni (2006) and Castelnuovo and Surico (2010). The reaction of output is in line with conventional wisdom, in that it signals a recession, whose timing and persistence is shown to depend on the business cycle indicator employed in each given VAR.

2 Bayesian estimation

To perform our Bayesian estimations we employed `DYNARE`, a set of algorithms developed by Michel Juillard and collaborators (Adjemian, Bastani, Juillard, Mihoubi, Perendia, Ratto, and Villemot (2011)). `DYNARE` is freely available at the following URL: http://www.dynare.org/.

The simulation of the target distribution is basically based on two steps.

- First, we initialized the variance-covariance matrix of the proposal distribution and employed a standard random-walk Metropolis-Hastings for the first $t \leq t_0 = 20,000$ draws. To do so, we computed the posterior mode by the "csminwel" algorithm developed by Chris Sims. The inverse of the Hessian of the target distribution evaluated at the posterior mode was used to define the variance-covariance matrix C_0 of the proposal distribution. The initial VCV matrix of the forecast errors in the Kalman filter was set to be equal to the unconditional variance of the state variables. We used the steady-state of the model to initialize the state vector in the Kalman filter.

- Second, we implemented the "Adaptive Metropolis" (AM) algorithm developed by Haario, Saksman, and Tamminen (2001) to simulate the target distribution. Haario, Saksman, and Tamminen (2001) show that their AM algorithm is more efficient than the standard Metropolis-Hastings algorithm. In a nutshell, such algorithm employs the history of the states (draws) so to 'tune' the proposal distribution suitably. In particular, the previous draws are employed to regulate the VCV of the proposal density. We then exploited the history of the states sampled up to $t > t_0$ to continuously update the VCV matrix C_t of the proposal distribution. While not being a Markovian process, the AM algorithm is shown to possess the correct ergodic properties. For technicalities, see Haario, Saksman, and Tamminen (2001).

We simulated two chains of 200,000 draws each, and discarded the first 90% as burn-in. To scale the variance-covariance matrix of the chain, we used a factor so to achieve an acceptance rate belonging to the [23%,40%] range. The stationarity of the chains was assessed via the convergence checks proposed by Brooks and Gelman (1998). The region of acceptable parameter realizations was truncated so to obtain equilibrium uniqueness under rational expectations.

3 Further results on the small-scale model

3.1 Predictive power of the estimated small-scale model

We checked the predictive power of the estimated small-scale model. It contrasts the actual series employed in our empirical exercise with the DSGE model's one step-ahead predictions. As shown by Figure A3, the model performs well along the one-step ahead forecasting dimension.

3.2 The role of the "timing discrepancy"

Are the distortions induced by the Cholesky-decomposition *quantitatively* relevant? Figure A4 displays the histograms of the distribution of the quarter-specific percentage deviations. The distributions are clearly shifted leftward with respect to the zero value, so indicating underestimation of the true effects of a monetary policy shock, or wrongly signed responses. The 90% sets suggest that these distortions are important also once sample uncertainty is accounted for. The median deviation reads -102% for inflation, and -97% for the output gap, i.e. the deviations from the true impulse responses are clearly sizeable.

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 inflation and output, the CVAR imposes a *delayed* reaction. To understand the consequences of this timing issue, let's stick to our small-scale model and consider the set of unique decision rules consistent with the rational expectation assumption and the structure of our small-scale DSGE model:[1]

$$\begin{bmatrix} \pi_t \\ y_t \\ R_t \end{bmatrix} = \mathbf{\Gamma} \begin{bmatrix} \pi_{t-1} \\ y_{t-1} \\ R_{t-1} \end{bmatrix} + \boldsymbol{B} \begin{bmatrix} \varepsilon_t^{\pi} \\ a_t \\ \varepsilon_t^{R} \end{bmatrix}, \mathbf{\Gamma} \equiv \begin{bmatrix} a_1 & f_1 & e_1 \\ a_2 & f_2 & e_2 \\ a_3 & f_3 & e_3 \end{bmatrix}, \boldsymbol{B} \equiv \begin{bmatrix} b_1 & c_1 & d_1 \\ b_2 & c_2 & d_2 \\ b_3 & c_3 & d_3 \end{bmatrix} \quad (1)$$

where $\mathbf{\Gamma}$ and $\boldsymbol{B}$ collect convolutions of the structural parameters $\boldsymbol{\xi}$ of the DSGE model. Given that the third column of $\boldsymbol{B}$ does not display, in general, zeros, the monetary policy shock ε_t^R *immediately* affects *all* the variables of the system.

The small-scale model has the following VAR(2) representation:

$$\begin{bmatrix} \pi_t \\ y_t \\ R_t \end{bmatrix} = \boldsymbol{A}_1 \begin{bmatrix} \pi_{t-1} \\ y_{t-1} \\ R_{t-1} \end{bmatrix} + \boldsymbol{A}_2 \begin{bmatrix} \pi_{t-2} \\ y_{t-2} \\ R_{t-2} \end{bmatrix} + \boldsymbol{B} \begin{bmatrix} u_t^{\pi} \\ u_t^{a} \\ u_t^{R} \end{bmatrix} \quad (2)$$

where $\boldsymbol{A}_1 = \mathbf{\Gamma} + \boldsymbol{BFB}^{-1}$ and $\boldsymbol{A}_2 = -\boldsymbol{BFB}^{-1}\mathbf{\Gamma}$. The variance-covariance matrix of $\boldsymbol{Bu}$ is given by $\boldsymbol{B\Omega B}^T$, where $\mathbf{\Omega}$ is a diagonal matrix of full rank 3 with the variances of the shocks positioned on the main diagonal. For ease of exposition (and without loss of generality), we set $\mathbf{\Omega} = \boldsymbol{I}_3$.

[1] The theoretical part of this Section heavily relies on the derivations proposed by Carlstrom, Fuerst, and Paustian (2009).

Of course, when conducting an econometric exercise, the fundamental shocks $\boldsymbol{u}_t$ are not observable, and must be inferred. To do so, the econometrician can estimate a reduced form VAR(2)

$$\begin{bmatrix} \pi_t \\ y_t \\ R_t \end{bmatrix} = \boldsymbol{A}_1 \begin{bmatrix} \pi_{t-1} \\ y_{t-1} \\ R_{t-1} \end{bmatrix} + \boldsymbol{A}_2 \begin{bmatrix} \pi_{t-2} \\ y_{t-2} \\ R_{t-2} \end{bmatrix} + \begin{bmatrix} \zeta_t^{\pi} \\ \zeta_t^{a} \\ \zeta_t^{R} \end{bmatrix},$$

where $\boldsymbol{\zeta}_t$ is a vector of *residuals* whose variance-covariance $VCV(\boldsymbol{\zeta}) = \boldsymbol{\Lambda}$ is a full (non diagonal) $[3x3]$ matrix.

To recover the unobserved structural monetary policy shock u_t^R, a researcher must impose some restrictions on the structure of the VAR, e.g. the simultaneous relationships among the variables included in the vector, the long-run impact of some economic shocks, or the sign of some conditional correlations. The most popular choice is to orthogonalize the residuals by imposing a Cholesky structure to the system, which assumes delayed effects of the 'monetary policy shock' on the variables located before the nominal interest rate in the vector $[\pi_t, y_t, R_t]^T$. This is done by computing the unique lower triangular matrix $\widetilde{\boldsymbol{B}}$ such that

$$\widetilde{\boldsymbol{B}}\boldsymbol{\varphi_t} = \boldsymbol{\zeta}, \text{ with } \widetilde{\boldsymbol{B}} = \begin{bmatrix} \widetilde{b}_1 & 0 & 0 \\ \widetilde{b}_2 & \widetilde{c}_2 & 0 \\ \widetilde{b}_3 & \widetilde{c}_3 & \widetilde{d}_3 \end{bmatrix}, \text{ and } \boldsymbol{\varphi_t} = \begin{bmatrix} \varphi_t^{\pi} \\ \varphi_t^{a} \\ \varphi_t^{R} \end{bmatrix}. \tag{3}$$

The Cholesky 'shocks' $\boldsymbol{\varphi_t}$, which are orthogonal and are assumed to have unitary variance, are then identified by computing the elements of the matrix $\widetilde{\boldsymbol{B}}$ such that

$$\widetilde{\boldsymbol{B}}\widetilde{\boldsymbol{B}}^T = \boldsymbol{\Lambda}.$$

This implies that the equivalence $\widetilde{\boldsymbol{B}}\widetilde{\boldsymbol{B}}^T = \boldsymbol{B}\boldsymbol{B}^T$ must hold. Solving the system, it is then possible to express the elements of $\widetilde{\boldsymbol{B}}$ in terms of the objects belonging to $\boldsymbol{B}$.

Given the restriction

$$\widetilde{\boldsymbol{B}}\boldsymbol{\varphi}_t = \boldsymbol{B}\boldsymbol{u}_t \tag{4}$$

imposed by eqs. (2) and (3), one may express the Cholesky-'shocks' $\boldsymbol{\varphi_t}$ in terms of the DSGE shocks $\boldsymbol{u}_t$ and the elements belonging to the matrix $\boldsymbol{B}$.

$$\boldsymbol{\varphi}_t = \Phi\boldsymbol{u}_t = \begin{bmatrix} \phi_{11} & \phi_{12} & \phi_{13} \\ \phi_{21} & \phi_{22} & \phi_{23} \\ \phi_{31} & \phi_{32} & \phi_{33} \end{bmatrix} \begin{bmatrix} u_t^{\pi} \\ u_t^{a} \\ u_t^{R} \end{bmatrix}, \tag{5}$$

where $\Phi \equiv \widetilde{\boldsymbol{B}}^{-1}\boldsymbol{B}$. Therefore, the mapping going from the true DSGE shocks to the CVAR monetary policy 'shock' reads

$$\varphi_t^R = \phi_{31}u_t^\pi + \phi_{32}u_t^a + \phi_{33}u_t^R. \tag{6}$$

The 'shock' φ_t^R is, in fact, a misspecified representation of the true monetary policy shock u_t^R. The standard Cholesky identification scheme recovers the true policy shock only under the restrictions $\phi_{31} = \phi_{32} = 0$. These would occur under $d_1 = d_2 = 0$ in the monetary impulse vector $\boldsymbol{B}[:,3]$ in eq. (1), i.e. if the structural DSGE model would feature lags in the impact of the true monetary policy shock u_t^R on inflation and output. However, these restrictions are *not* consistent with the DSGE models employed by most researchers, the model we focus on in this paper included. The calibration conditional on our estimated posterior means implies the following values for the matrices characterizing the set of decision rules (1):

$$\boldsymbol{\Gamma} = \begin{bmatrix} 0.08 & 0.03 & -0.27 \\ -0.03 & 0.25 & -0.62 \\ 0.03 & 0.02 & 0.68 \end{bmatrix}, and \boldsymbol{B} = \begin{bmatrix} 1.78 & -0.10 & -0.63 \\ -5.22 & -0.17 & -1.22 \\ 0.59 & -0.05 & 0.70 \end{bmatrix}.$$

Notably, $\boldsymbol{B}[1,3] = d_1 = -0.63$, and $\boldsymbol{B}[2,3] = d_2 = -1.22$. As a consequence, the Cholesky scheme misspecifies the monetary policy shock.

Figure A5 (top row) plots these densities. Interestingly, the cost-push shock u_t^π enters the reduced form CVAR monetary policy shock with a weight close to zero on average. Differently, the distribution of the weight ϕ_{32} assigned to the technology shock u_t^a is negative and 'significantly' different from zero, with a mean equal to -0.40. Also the density of the loading ϕ_{33} of the shock u_t^R suggests values different from zero, and displays a mean close to 6.5.

The variance decomposition analysis of the reduced form shock φ^R is depicted in Figure A5 (bottom row). The contribution of the true technology shock u_t^a is on average 31%, a figure stressing that the misspecification induced by the imposition of the Cholesky scheme is substantial. In theory, also the cost-push shock u_t^π could contribute to bias the reduced form monetary policy shock φ_t^R. In practice, however, its contribution is negligible, with a mean around 1%. The remaining volatility is due to the true monetary policy shock u_t^R.

These findings offer a rationale for the distorted CVAR responses we obtain with our MonteCarlo exercises. The stochastic element identified by the CVAR monetary policy 'shock' is in fact a convolution of the true technology shock u_t^a, which enters the reduced form φ_t^R with a *negative* weight, and the true monetary policy shock u_t^R, which enters it with a *positive* sign. A negative technology shock opens a positive output gap, which exerts a positive pressure on inflation and the policy rate. At the same time, a monetary policy shock (a policy tightening) triggers a positive reaction of the policy rate, and a negative reaction of inflation and the output gap. Then, the reduced form shock φ_t^R actually captures the *joint* effects of these two structural shocks, which basically offset each other as for inflation and output, leading to muted reactions like those depicted in Figures 3.

4 The Smets-Wouters (2007) model

The Smets and Wouters (2007) model is a Dynamic Stochastic General Equilibrium framework extremely popular in the academic and institutional circles. The model features a number of shocks and frictions, which offer a quite rich representation of the economic environment and allow for a satisfactory in-sample fit of a set of macroeconomic data (Del Negro, Schorfheide, Smets, and Wouters (2007)). Moreover, Smets and Wouters (2007) show that this model is quite competitive when contrasted with Bayesian-VARs as for forecasting exercises, in particular for the elaboration of medium-term predictions.

The Smets and Wouters (2007) model features sticky nominal price and wage settings that allow for backward-looking inflation indexation; habit formation in consumption; investment adjustment costs; variable capital utilization and fixed costs in production. The stochastic dynamics is driven by seven structural shocks, namely a total factor productivity shock, two shocks affecting the intertemporal margin (risk premium shocks and investment-specific technology shocks), two shocks affecting the intratemporal margin (wage and price mark-up shocks), and two policy shocks (exogenous spending and monetary policy shocks).

In a nutshell, the model features the following main ingredients. Households maximize a nonseparable utility function in consumption and labor over an infinite life horizon. Consumption appears in the utility function in quasi-difference form with respect to a time-varying external habit variable. Labor is differentiated by a union, so there is some monopoly power over wages, which results in explicit wage equation and allows for the introduction of sticky nominal wages à la Calvo (1983). Households rent capital services to firms and decide how much capital to accumulate given the capital adjustment costs they face. The utilization of the capital stock can be adjusted at increasing cost. Firms produce differentiated goods, decide on labor and capital inputs, and set prices conditional on the Calvo model. The Calvo model in both wage and price setting is augmented by the assumption that prices that are not reoptimized are partially indexed to past inflation rates. Prices are therefore set in function of current and expected marginal costs, but are also determined by the past inflation rate. The marginal costs depend on wages and the rental rate of capital. Similarly, wages depend on past and expected future wages and inflation. The model features, in both goods and labor markets, an aggregator that allows for a time-varying demand elasticity depending on the relative price as in Kimball (1995). This is important because the introduction of real rigidity allows us to estimate a more reasonable degree of price and wage stickiness.

The log-linearized version of the DSGE model around its steady-state growth path reads as follows:

$$
\begin{aligned}
y_t &= c_y c_t + i_y i_t + z_y z_t + \varepsilon_t^g && (7)\\
c_t &= c_1 c_{t-1} + (1-c_1)E_t c_{t+1} + c_2(l_t - E_t l_{t+1}) - c_3(r_t - E_t\pi_{t+1} + \varepsilon_t^b) && (8)\\
i_t &= i_1 i_{t-1} + (1-i_1)E_t i_{t+1} + i_2 q_t + \varepsilon_t^i && (9)\\
q_t &= q_1 E_t q_t + 1 + (1-q_1)E_t r_{t+1}^k - (r_t - E_t\pi_{t+1} + \varepsilon_t^b) && (10)\\
y_t &= \phi_p(\alpha k_t^s + (1-\alpha)l_t + \varepsilon_t^a) && (11)\\
k_t^s &= k_{t-1} + z_t && (12)\\
z_t &= z_1 r_t^k && (13)\\
k_t &= k_1 k_{t-1} + (1-k_1)i_t + k_2\varepsilon_t^i && (14)\\
\mu_t^p &= \alpha(k_t^s - l_t) + \varepsilon_t^a - w_t && (15)\\
\pi_t &= \pi_1\pi_{t-1} + \pi_2 E_t\pi_{t+1} - \pi_3\mu_t^p + \varepsilon_t^p && (16)\\
r_t^k &= -(k_t - l_t) + w_t && (17)\\
\mu_t^w &= w_t - (\sigma_l l_t + (1-\lambda/\gamma)^{-1}(c_t - \lambda/\gamma c_{t-1})) && (18)\\
w_t &= w_1 w_{t-1} + w_2(E_t w_{t+1} + E_t\pi_{t+1}) - w_2\pi_t + w_3\pi_{t-1} - w_4\mu_t^w + \varepsilon_t^w && (19)\\
r_t &= \rho r_{t-1} + (1-\rho)(r_\pi + r_Y(y_t - y_t^p)) + r_{\Delta y}[(y_t - y_t^p) - (y_{t-1} - y_{t-1}^p)] + \varepsilon_t^R && (20)\\
\varepsilon_t^x &= \rho_x\varepsilon_{t-1}^x + \eta_t^x, x = (b, i, a, R) && (21)\\
\varepsilon_t^g &= \rho_g\varepsilon_{t-1}^g + \eta_t^g + \rho_{ga}\eta_t^a && (22)\\
\varepsilon_t^z &= \rho_x\varepsilon_{t-1}^z + \eta_t^z - \chi_z\eta_{t-1}^z, z = (p, w) && (23)\\
\eta_t^j &\sim N(0, \sigma_j^2) && (24)
\end{aligned}
$$

where:

$$c_y = 1 - g_y - i_y \tag{25}$$

and g_y and i_y are the steady-state exogenous spending-output ratio and investment-output ratio, with:

$$i_y = (\gamma - 1 + \delta)k_y \tag{26}$$

where γ is the steady-state growth rate, δ is the depreciation rate of capital, k_y is the steady-state capital-output ratio; $z_y = R_*^y k_y$ is the steady-state rental rate of capital. Notice that eq. (22), the one of the stochastic process of the government spending, allows for the productivity shock to affect it. This is so because exogenous spending, in this model, includes net exports, which may be affected by domestic productivity development.

As for the consumption Euler equation (8):

$$c_1 = \frac{\lambda}{\gamma}\left(1 + \frac{\lambda}{\gamma}\right) \tag{27}$$

$$c_2 = \frac{(\sigma_c - 1)\frac{W_*^h L_*}{C_*}}{\sigma_c(1 + \frac{\lambda}{\gamma})} \tag{28}$$

$$c_3 = \frac{1 - \frac{\lambda}{\gamma}}{\left(1 + \frac{\lambda}{\gamma}\right)\sigma_c} \tag{29}$$

Current consumption is a function of past and expected future consumption, of expected growth in hours worked, of the ex ante real interest rate, and of a disturbance term ε_t^b. Under the assumption of no habits ($\lambda = 0$) and that of log-utility in consumption ($\sigma_c = 1$), $c_1 = c_2 = 0$, then the standard purely forward looking consumption equation is obtained. The disturbance term ε_t^b represents a wedge between the interest rate controlled by the central bank and the return on assets held by the households. A positive shock to this wedge increases the required return on assets held by the households. At the same time, it increases the cost of capital and it decreases the value of capital and investment (see below). This is basically a shock very similar to a net-worth shock. This disturbance is assumed to follow a standard AR(1) process.

The dynamics of investment is captured by the investment Euler equation (9), where:

$$i_1 = \frac{1}{1 + \beta\gamma^{1-\sigma_c}} \tag{30}$$

$$i_2 = \frac{1}{1 + \beta\gamma^{1-\sigma_c}\gamma^2\varphi} \tag{31}$$

where φ is the steady-state elasticity of the capital adjustment cost function, and β is the discount factor applied by households. Notice that capital adjustment costs are a function of the change in investment, rather than its level. This choice is made to introduce additional dynamics in the investment equation, which is useful to capture the hump-shaped response of investment to various shocks. In this equation, the stochastic disturbance ε_t^i represents a shock to the investment-specific technology process, and is assumed to follow a standard first-order autoregressive process.

The value-of-capital arbitrage equation (10) suggests that the current value of the capital stock q_t depends positively on its expected future value (with weight $q_1 = \beta\gamma^{-\sigma_c}(1-\delta)$), as well as the expected real rental rate on capital $E_t r_{t+1}^k$ and on the ex ante real interest rate and the risk premium disturbance.

Eq. (11) is the first one of the supply side block. It describes the aggregate production function, which maps output to capital (k_t^s) and labor services (l_t). The parameter α captures the share of capital in production, and the parameter ϕ_p is one plus the share of fixed costs in production, reflecting the presence of fixed costs in production.

Eq. (12) suggest that the newly installed capital becomes effective with a one-period delay, hence current capital services in production are a function of capital installed in the previous period k_t and the degree of capital utilization z_t. As stressed by eq. (13), the degree of capital utilization is a positive function of the rental rate of capital, $z_t = z_1 r_t^k$, where $z_1 = (1-\psi)/\psi$ and ψ is a positive function of the elasticity of the capital utilization adjustment cost function normalized to belong to the [0,1] domain.

Eq. (14) describes the accumulation of installed capital k_t, featuring the convolutions:

$$k_1 = (1-\delta)/\gamma \tag{32}$$

$$k_2 = \left[1 - \left(1 - \frac{\delta}{\gamma}\right)\right]\left(1 + \beta\gamma^{1-\sigma_c}\right)\gamma^2\varphi \tag{33}$$

Installed capital is a function not only of the flow of investment but also of the relative efficiency of these investment expenditures as captured by the investment-specific technology disturbance ε_t^i, which follows an autoregressive, stationary process.

Eq. (15) relates to the monopolistic competitive goods market. Cost minimization by firms implies that the price mark-up μ_t^p, defined as the difference between the average price and the nominal marginal cost or the negative of the real marginal cost, is equal to the difference between the marginal product of labor and the real wage w_t, with the marginal product of labor being itself a positive function of the capital-labor ratio and total factor productivity.

Profit maximization by price-setting firms gives rise to the New-Keynesian Phillips curve, i.e., eq. (16), with the convolutions being:

$$\pi_1 = \frac{\iota_p}{1+\beta\gamma^{1-\sigma_c}\iota_p}, \tag{34}$$

$$\pi_2 = \frac{\beta\gamma^{1-\sigma_c}}{1+\beta\gamma^{1-\sigma_c}\iota_p}, \tag{35}$$

$$\pi_3 = \frac{1}{1+\beta\gamma^{(1-\sigma_c)}\iota_p}\frac{(1-\beta\gamma^{1-\sigma_c}\xi_p)(1-\xi_p)}{\xi_p\left[(\phi_p-1)\varepsilon_p+1\right]}. \tag{36}$$

Notice that, in maximizing their profits, firm have to face price stickiness à la Calvo (1983). Firms that cannot reoptimize in a given period index their prices to past inflation as in Smets and Wouters (2003). In equilibrium, inflation π_t depends positively on pat and expected future inflation, negatively on the current price mark-up, and positively on a price mark-up disturbance ε_t^p. The price mark-up disturbance is assumed to follow an ARMA(1,1) process. The inclusion of the MA term is to grab high-frequency fluctuations in inflation. When the degree of price indexation $\iota_p = 0$, $\pi_1 = 0$ and eq. (16) collapses to the purely forward-looking, standard NKPC. The assumption that all prices are indexed to either lagged inflation or trend inflation ensures the verticality of the Phillips curve in the long run. The speed of adjustment to the desired mark-up depends, among others, on the degree of price stickiness ξ_p, the curvature of the Kimball goods market aggregator ε_p, and the steady-state mark up, which in equilibrium is itself related to the share of fixed costs in production $(\phi_p - 1)$ via a zero-profit condition. In particular, when all prices are flexible ($\xi_p = 0$) and the price mark-up shock is zero at all times, eq. (16) reduces to the familiar condition that the price mark-up is constant, or equivalently that there are no fluctuations in the wedge between the marginal product of labor and the real wage. Cost minimization by firms also implies that the rental rate of capital is negatively related to the capital-labor ratio and positively to the real wage (both with unitary elasticity) (see eq. (17)).

Similarly, in the monopolistically competitive labor market, the wage mark-up will be equal to the difference between the real wage and the marginal rate of substitution between working and consuming, an equivalence captured by eq. (18), where σ is the elasticity of labor supply with respect to the real wage and λ is the habit parameter in consumption. Eq. (19) shows that real wages adjust only gradually to the desired wage mark-up due to nominal wage stickiness and partial indexation, the convolutions related to this equation being:

$$w_1 = \frac{1}{1+\beta\gamma^{1-\sigma_c}} \tag{37}$$

$$w_2 = \frac{1+\beta\gamma^{1-\sigma_c}\iota_w}{1+\beta\gamma^{1-\sigma_c}} \tag{38}$$

$$w_3 = \frac{\iota_w}{1+\beta\gamma^{1-\sigma_c}} \tag{39}$$

$$w_4 = \frac{\iota_w}{1+\beta\gamma^{1-\sigma_c}} \frac{(1-\beta\gamma^{(1-\sigma_c)}\xi_w)(1-\xi_w)}{\xi_w\left[(\phi_w-1)\varepsilon_w+1\right]} \tag{40}$$

Notice that if wages are perfectly flexible ($\xi_w = 0$), the real wage is a constant mark-up over the marginal rate os substitution between consumption and leisure. When wage indexation is zero ($\iota_w = 0$), real wages do not depend on lagged inflation. Notice that, symmetrically with respect to the pricing scheme analyzed earlier, also the wage-mark up disturbance follows an ARMA(1,1) process.

The model is closed by eq. (20), which is a flexible Taylor rule postulating a systematic reaction by policymakers to current values of inflation, the output gap, and output growth. In particular, one of the objects policymakers react to is the output gap, defined as a difference between actual and potential output (in logs). Consistently with the DSGE model, potential output is defined as the level of output that would prevail under flexible prices and wages in the absence of the two mark-up shocks. Then, policymakers engineer movements in the short-run policy rate r_t, movements which happen gradually given the presence of interest rate smoothing ρ. Stochastic departures from the Taylor rate, i.e. the rate that would realize in absence of any policy rate shocks, are triggered by a stochastic AR(1) process.

Finally, eqs. (21)-(24) define the stochastic processes of the model, which features, as already pointed out, seven shocks (total factor productivity, investment specific technology, risk premium, exogenous spending, price mark-up, wage mark-up, and monetary policy).

Notice that 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.

5 Further results with Smets and Wouters (2007)

Figure A6 plots the results of our MonteCarlo exercise conditional on an estimated version of the Smets and Wouters (2007) featuring no reaction to the output gap by the Federal Reserve, which focuses on inflation and output growth only. The similarity between this Figure and Figure 4 in the text is striking, i.e., our results are not affected by this perturbation in the policy rule.

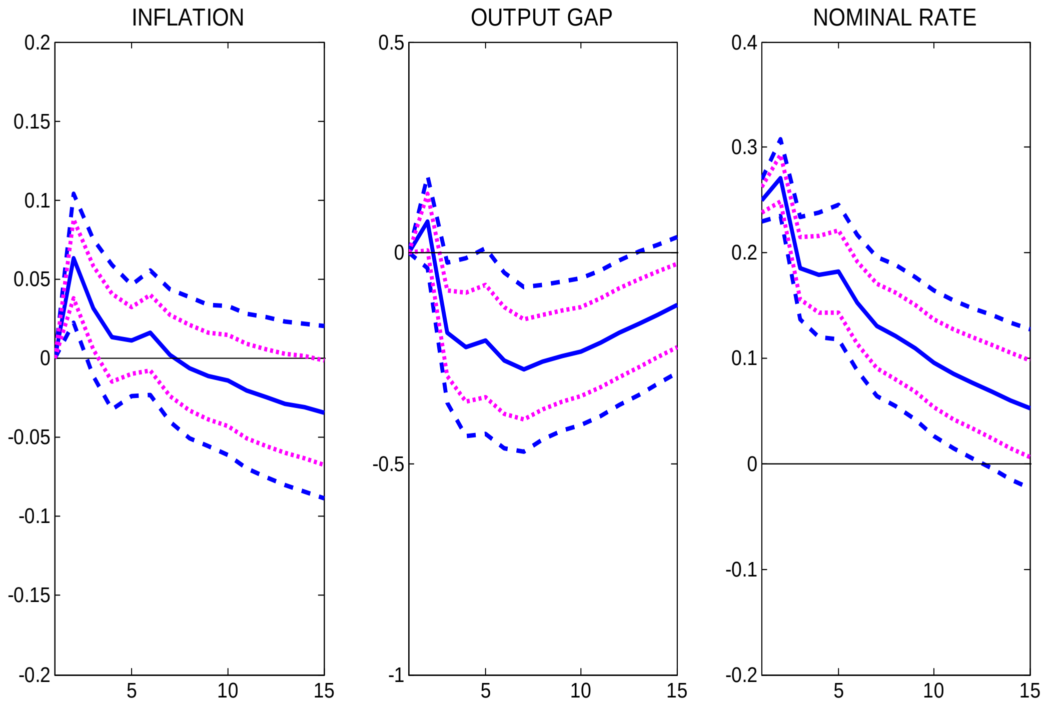

Figure A1: **CVAR impulse response functions to a monetary policy shock, 1954:III-2008:II - model with CBO output gap**. Variables: Quarterly GDP inflation, CBO output gap, quarterly federal funds rate - source: FREDII. Identification of the monetary policy shock via Cholesky decomposition (lower triangular matrix, ordering: inflation, output 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 four lags.

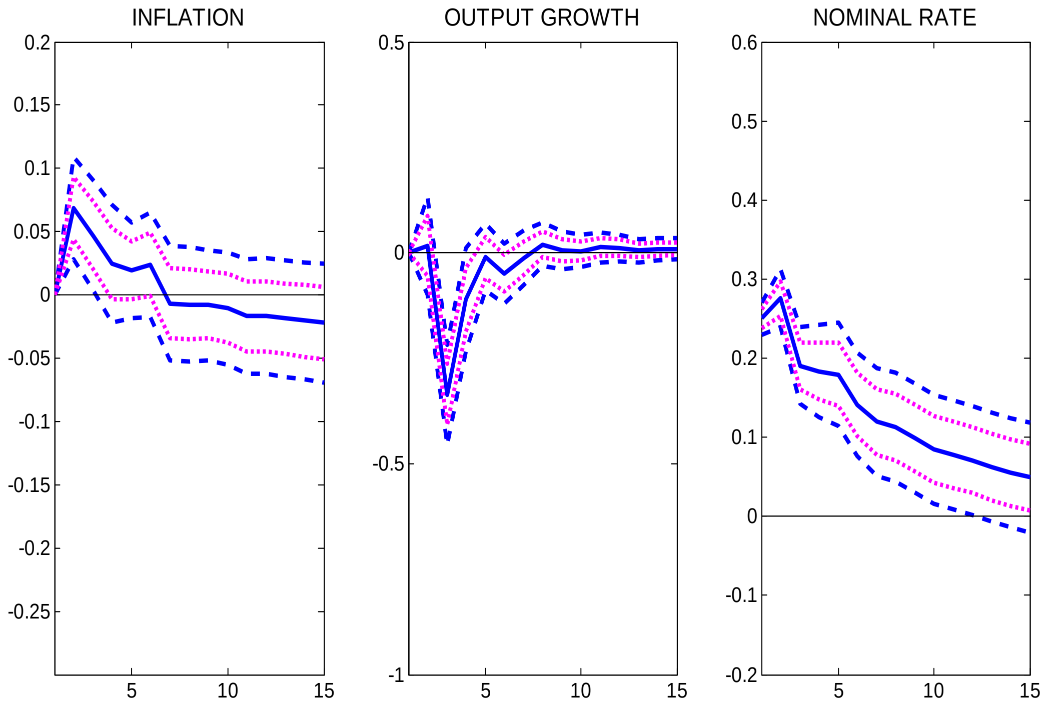

Figure A2: **CVAR impulse response functions to a monetary policy shock, 1954:III-2008:II - model with output growth**. Variables: Quarterly GDP inflation, quarterly output growth, quarterly federal funds rate - source: FREDII. Identification of the monetary policy shock via Cholesky decomposition (lower 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 four lags.

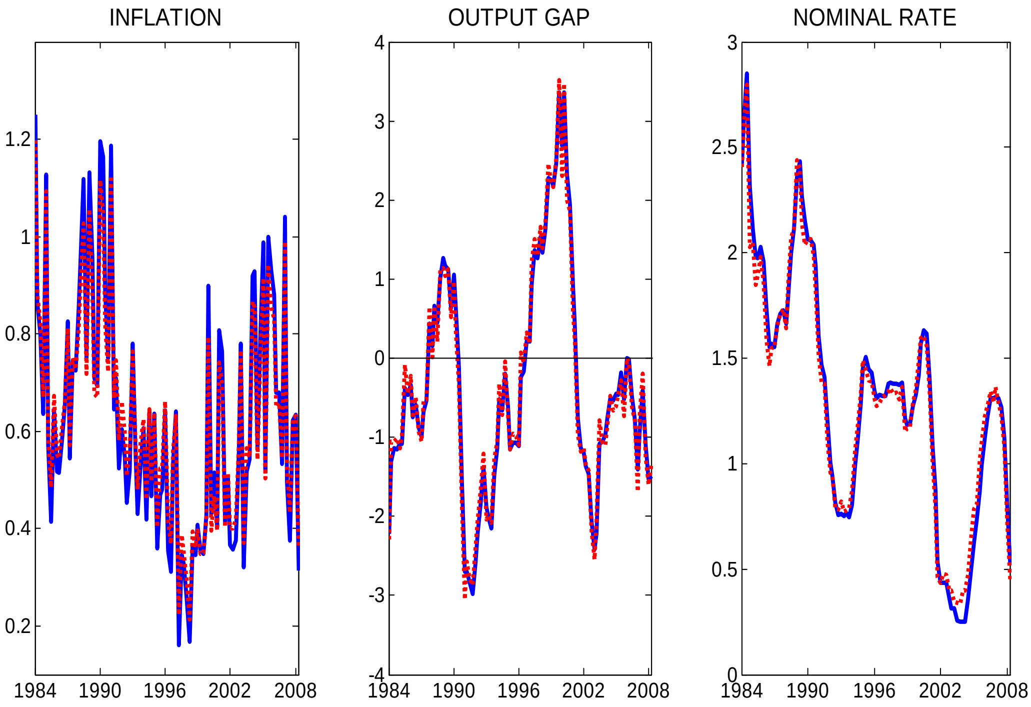

Figure A3: **Actual series vs. DSGE's one-step ahead forecasts.** Solid blue line: Actual series; Dotted red lines: DSGE's one-step-ahead predictions.

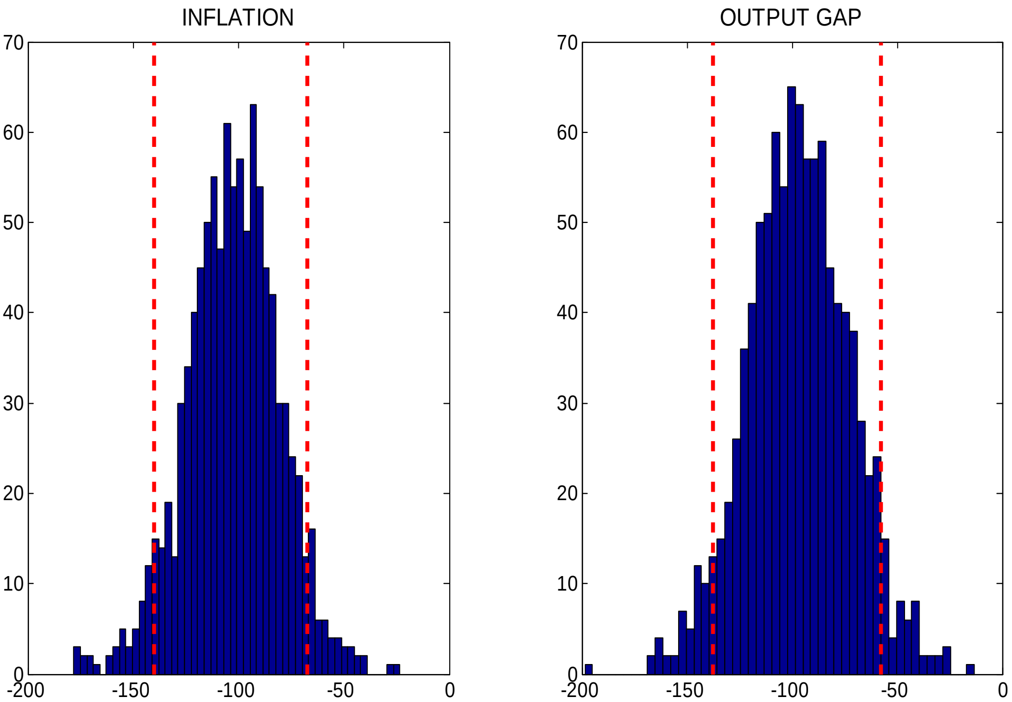

Figure A4: **Impulse response functions: Percentage deviations**. Percentage deviations of the CVAR responses with respect to the DSGE (true) responses - one quarter after the shock. Red dotted lines: [5th,95th] percentiles. Computation of the densities based on 5,000 draws of the structural parameters of the DSGE model.

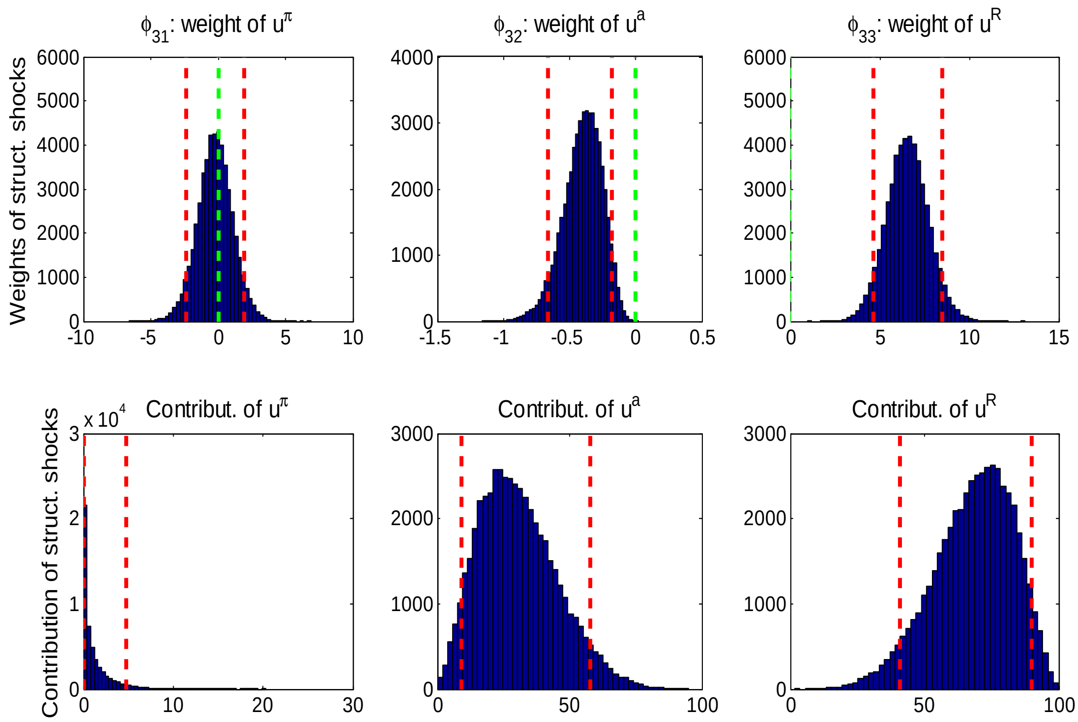

Figure A5: **CVAR monetary policy shock: Weights and contributions of the DSGE's shocks.** Distribution computed over 5,000 stochastic simulations.

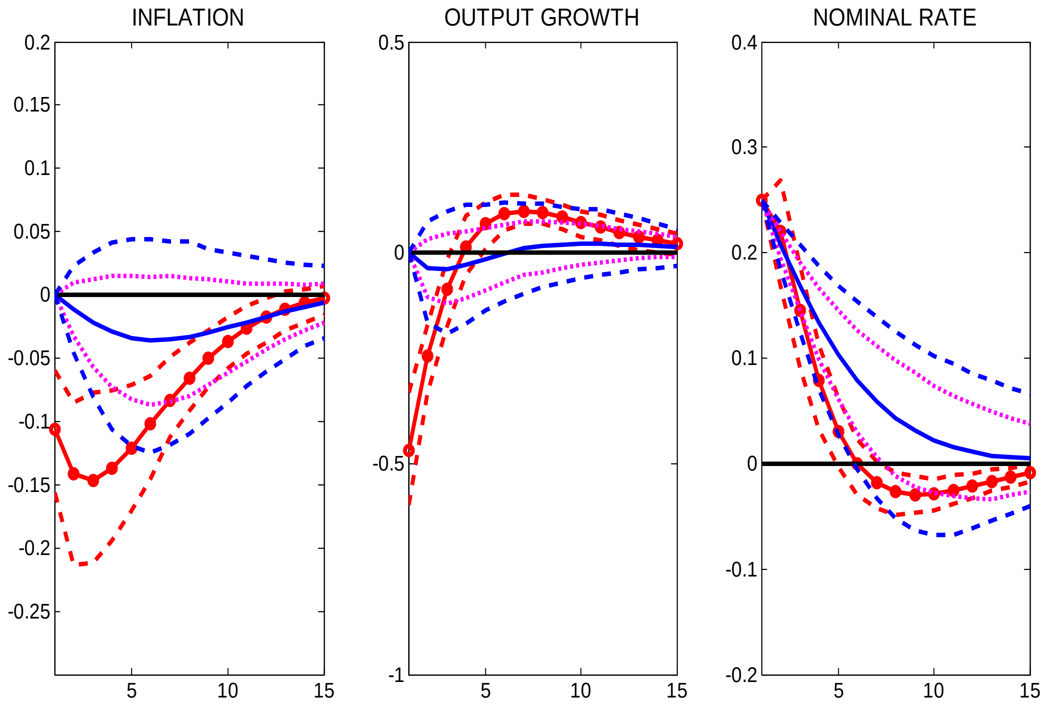

Figure A6: **DSGE à la Smets and Wouters (2007) vs. CVAR impulse response functions to a monetary policy shock - rule featuring no output gap.** Circled red lines: DSGE Bayesian mean impulse responses; Dashed red lines: 90% credible sets. Solid blue line: CVAR mean impulse responses; Dashed blue lines: [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.

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