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Abstract

Cholesky-VAR impulse responses estimated with post-1984 U.S. data predict modest macroeconomic reactions to monetary policy shocks. We interpret this evidence by employing an estimated medium-scale DSGE model of the business cycle as a Data-Generating Process in a Monte Carlo exercise in which a Cholesky-VAR econometrician is asked to estimate the effects of an unexpected, temporary increase in the policy rate. Our structural DSGE model predicts conventional macroeconomic reactions to a policy shock. In contrast, our Monte Carlo VAR results replicate our evidence obtained with actual U.S. data. Hence, modest macroeconomic effects may very well be an artifact of Cholesky-VARs. A combination of supply and demand shocks may be behind the inability of Cholesky-VARs to replicate the actual macroeconomic responses. The difference in the VAR responses obtained with Great Inflation vs. Great Moderation data may be due to instabilities in the parameters related to households' and firms' programs, more than to a more aggressive systematic monetary policy. A Monte Carlo assessment of sign restrictions as an alternative identification strategy is also proposed.

JEL classification: C3, E3

Keywords: Monetary policy shocks, Cholesky identification, VARs, Dynamic Stochastic General Equilibrium models, Monte Carlo simulations

1 Introduction

Vector Autoregression (VAR) empirical investigations dealing with U.S. data have often found modest macroeconomic effects of monetary policy shocks during the great moderation (see, among others, Bagliano and Favero (1998), Hanson (2004), Boivin and Giannoni (2006), Mojon (2008), Castelnuovo and Surico (2010), and Boivin, Kiley, and Mishkin (2010)). Figure 1 replicates this evidence.[1] The bottom panels point to a non-significant reaction of output, and to a short-lived negative response of inflation. Very different results are typically obtained when dealing with samples including the 1970s, and our dataset represents no exception. The top panels in Figure 1 show a significant and positive reaction of inflation (known as the 'price puzzle') and a negative and persistent response of output as in, e.g., Christiano, Eichenbaum, and Evans (2005).

Various interpretations have been given to such 'mild-to-muted' reactions. Hanson (2004) discusses the role possibly played by the change in the systematic policy conduct of the Federal Reserve occurred with the advent of Paul Volcker in 1979, or by a milder volatility of output during the great moderation. Boivin and Giannoni (2006) and Castelnuovo and Surico (2010) underline the role that technological progress or financial innovations may have played in easing households' consumption smoothing, or - again - the improved monetary policy management in the 1980s. Boivin, Kiley, and Mishkin (2010) also point to financial innovations and a more hawkish monetary policy conduct as possible drivers behind the much more moderate macroeconomic reactions in the post-1984 period.[2]

This paper provides an alternative interpretation to the modest reactions shown in

[1] 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. Robustness checks documented in our Appendix suggest that this qualitative message does not change if a measure of output growth is employed in our VARs. Boivin and Giannoni (2006) and Boivin, Kiley, and Mishkin (2010) confirm this evidence with Factor-Augmented VARs embedding information coming from large datasets.

[2] Mojon (2008) shows that, once one examines periods without large shifts in the level of inflation such as the great moderation, the delayed and persistent effects on inflation often attributed to monetary policy shocks tend to disappear. Other '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 may have reacted to. On this latter point, see Barakchian and Crowe (2013), who employ monthly data in their analysis. The relevance of their results at quarterly frequencies for the great moderation sample is material for future research.

Figure 1. We do so by conducting a Monte Carlo exercise in which the Smets and Wouters (2007) model, a reference Dynamic Stochastic General Equilibrium (DSGE) framework for a large number of central banks in the world, is employed as the Data-Generating Process (DGP) to produce artificial data which are then given to a Cholesky-VAR econometrician. The Cholesky-VAR econometrician is asked to identify the macroeconomic responses of output and inflation to a monetary policy shock. We show that mild-to-muted responses are very likely to arise even if, in the Smets and Wouters (2007) framework, monetary policy shocks exert recessionary and deflationary effects. Consequently, a possible interpretation of our responses plotted in Figure 1 (bottom panels) is that modest-to-muted impulse responses may very well be an artifact due to the use of Cholesky-VARs by econometricians. We show that such an artifact is due to the timing discrepancy between the actual (unknown to the econometrician) DGP and the Cholesky-VAR. In fact, while the DSGE model allows for immediate output and inflation responses to a monetary policy shock, the zero-restrictions imposed by the Cholesky-identification schemes prevent any immediate response of output and inflation to occur. This time discrepancy induces a distortion in the estimation of the policy shocks, which end up being a linear combination of genuine monetary policy shocks and all the remaining non-policy, structural shocks affecting the economic system. Therefore, Cholesky-VARs pick up a combination of demand and supply shocks for which the net effect on inflation and output induces the modest reactions shown in Figure 1 (bottom panels).

Is there any structural reason behind the instabilities in the impulse responses depicted in Figure 1? We address this question by estimating our DGP with U.S. data referring to the great inflation period. Our posterior estimates document instabilities in the parameters related to households' and firms' problems, to the systematic monetary policy conduct, and to the exogenous processes of the structural shocks in the model. Counterfactual simulations are then performed to isolate the subset of parameters which is behind our findings. Interestingly, it turns out that the most important instability is the one affecting the parameters of the private sector. However, changes in the monetary policy parameters are unlikely to distort the Cholesky-VAR macroeconomic reactions toward zero. If anything, a more aggressive policy conduct is associated with a stronger reaction of inflation and output in a Cholesky-VAR context.

Our exercise provides indirect empirical support to Smets and Wouters' (2007) model. We show that the data produced with this medium-scale model produce, conditional on Cholesky-VARs, conditional correlations which are in line with those obtained with actual U.S. data when treated with the same type of VARs. Importantly, we show that this result is robust to using the growth rate of output (instead of the theoretically-relevant output gap), to different sample-sizes of artificial data to estimate our VARs, and to different lag-structures of our VARs. To be clear, this paper does not offer any reason to reject the Cholesky-VAR approach. Indeed, the appealing 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 of this issue). Moreover, this recursive structure is consistent with structural models of the business cycle recently proposed by Christiano, Eichenbaum, and Evans (2005), Boivin and Giannoni (2006) and Altig, Christiano, Eichenbaum, and Lindé (2011).[3] This paper's goal is instead to show that it is possible to interpret the modest macroeconomic reactions suggested by a number of Cholesky-VAR exercises in the literature with the help of a non-recursive structural DSGE model widely employed by researchers in academia and in a variety of research institutions.

We then scrutinize a different identification strategy, which is sign restrictions. Sign restrictions have recently been popularized by a number of authors (see Faust (1998), Canova and de Nicoló (2002), Peersman (2005), and Uhlig (2005), among others), and represent an obviously interesting alternative because they do not necessarily involve zero exclusion restrictions. In short, sign restrictions identify a monetary policy shock via the imposition of restrictions on signs for the responses of variables of interest (output, inflation, the policy rate). Hence, the on-impact response of the variables modeled with the VAR is theoretically consistent with that of the underlying DGP. We then re-run our Monte Carlo exercise and identify monetary policy shocks by imposing restrictions for the responses of output, inflation, and the policy rate consistent with the responses of the same variables in the Smets and Wouters (2007) model. Our simulations show that, in practice, standard restrictions point to impulse responses which are correct in terms of sign and shape but not in terms of magnitude. The reason for this latter result is that monetary policy shocks explain a little share of the volatility of the variables modeled with the VAR, something which negatively affects the empirical performance of sign restrictions (Paustian (2007), Castelnuovo (2015)). Interestingly,

[3]It is unclear if the recursive structure timing assumption is favored by the data as opposed to the contemporaneous timing as in Smets and Wouters (2007) and Canova (2009), among others. Formal evidence against zero-restrictions for the reactions of output and inflation to monetary policy surprises is offered by Del Negro, Schorfheide, Smets, and Wouters (2007) (for output) and Faust, Swanson, and Wright (2004) (for 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 proposed so far. We leave this important question to future research.

we find that the Uhlig (2005) result on the imprecisely estimated response of output to a monetary policy shock could very well be due to the weak signal associated with monetary policy shocks in the U.S. economy.[4]

The paper develops as follows. Section 2 reviews some contributions dealing with similar issues. Section 3 describes the Smets and Wouters (2007) model and documents our estimation results. This model is employed as our DGP in Section 4, which sets up our Monte Carlo experiment. This Section also documents our baseline results. Section 5 attempts to isolate the key parameters behind our modest macroeconomic reactions. Section 6 investigates the ability of sign restrictions to correctly determine the response of inflation and output to a monetary policy shock. Section 7 concludes.

2 Literature review

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 Cholesky-VARs. Carlstrom, Fuerst, and Paustian (2009) propose a theoretical investigation on the consequences of the timing discrepancy between DSGE and Cholesky-VARs as for macroeconomic reactions to a monetary policy shock. They show that, depending on the chosen calibration of their DSGE models, Cholesky-VARs may provide a variety of predictions, including price and output puzzles, responses in line with the actual DSGE reactions, muted responses, and so on. These papers make a theoretical point. In contrast, the relevance of our contribution is due to its empirical content. First, we employ an *estimated, medium-scale* DSGE model of the business cycle as a DGP in our Monte Carlo exercise. Second, we draw a comparison between Cholesky-VARs' responses estimated with artificial vs. actual U.S. data. This enables us to offer a novel interpretation of the modest macroeconomic reactions to a monetary policy shock found in a number of recent contributions. Third, we identify the set of

[4]Of course, sign restrictions can and do work well in practice when it comes to identifying other shocks (e.g., technology shocks), as shown in Paustian (2007) and Canova and Paustian (2011). This is because such shocks are important drivers of the dynamics (and, therefore, the volatility) of output and inflation in the United States. In contrast, monetary policy shocks are typically found to explain a little share of the variance of these macro aggregates.

parameters responsible for the modest macroeconomic reactions detected in Figure 1. Our analysis should be seen as complementary to the ones proposed by Canova and Pina (2005) and Carlstrom, Fuerst, and Paustian (2009).

Giacomini (2013) reviews the literature on the econometric relationship between DSGE and VAR frameworks for estimation and model validation. 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 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. In contrast, as it is customary in Monte Carlo exercises, we assume our (DSGE) model to be 'true'. Then, conditional on this assumption, we conduct a Monte Carlo 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. Their paper deals, however, with the issue of identification schemes within structural VARs (for which the authors provide econometric testing), but it does not discuss structural models. Altig, Christiano, Eichenbaum, and Lindé (2011) show that, in the presence of a model featuring a lagged transmission mechanism, Cholesky-VARs perform well. Our paper complements Altig et al.'s (2011) contribution by scrutinizing the ability of a VAR to recover the actual impulse responses under the null hypothesis of a contemporaneous timing being at work. Leeper and Roush (2003) and Poilly (2010) empirically show that Cholesky-VARs may return quite different responses to monetary shocks if the stock of money is a relevant variable in the DGP. With respect to these contributions, we do not consider money or exchange rates in our analysis, and we rely on Monte Carlo simulations which enable us to judge the performance of Cholesky-VARs given our awareness of the true, underlying DGP. Finally, Istrefi and Vonnak (2014) compare the performance of Cholesky- and sign restriction-VARs to recover the effects of monetary policy shocks on exchange rates. Our exercise is instead conducted in a close-economy context with the aim of mapping Cholesky-VARs estimated on actual vs. artificial data.

Finally, and related to our sign restriction-VAR simulations, the papers closest to ours are probably Paustian (2007) and Canova and Paustian (2012). Paustian (2007) derives a closed-form solution for the identification of a shock of interest with sign restrictions in a bivariate context. Such an expression points to the relative variance of the shock of interest as a key element to achieve identification. Then, Paustian (2007) conducts a Monte Carlo simulation and finds support for his theoretical result. With respect to Paustian (2007), we employ an estimated version of the Smets and Wouters (2007) model as DGP to contrast the performance of Cholesky-VARs and sign restriction-VARs. Canova and Paustian (2011) study how robust restrictions in a class of DSGE models can be used to identify structural shocks in the data and perform model selection. Our sign restriction exercises are conducted by appealing to restrictions which are 'robust' in the sense of Canova and Paustian (2011), and which are used to show that, from an empirical standpoint, it is difficult to correctly identify the effects of monetary policy shocks even with a strategy which is theoretically consistent with the restrictions associated with the underlying DGP.

3 Our Data-Generating Process: Structure and estimation

Our Monte Carlo experiment hinges upon the Smets and Wouters (2007) model, which is the framework we will use as our DGP. This model is well known and widely employed in academic and policy circles. Then, a detailed presentation of this model is confined in our Appendix (available upon request). This Section provides a brief description of the structure of the model as well as the way in which the latent factors of the model are matched with the U.S. data.

3.1 Model description

The Smets and Wouters (2007) model features a number of nominal and real frictions which are relevant to replicate the features of the main U.S. macroeconomic series. In particular, it models 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 model features 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 this model, households maximize a nonseparable utility function in consumption and labor over an infinite time 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 an 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 as a 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, both for goods and labor markets, an aggregator that allows for a time-varying demand elasticity depending on the relative price. The model features a deterministic growth rate driven by labor-augmenting technological progress.

The nonlinear version of the model is log-linearized around its steady-state growth path. The log-linearization gives rise to a number of linear difference equations which are responsible for the evolution of the endogenous variables in the system. Such equations are the aggregate resource constraint, the Euler equation for consumption, the investment equation, the equation keeping track of the evolution of Tobin's q, the production function, the equation capturing capital services via the impact of capital utilization on physical capital, the equation relating capital utilization to the rental rate of capital, the expression for such rental rate, the law of motion of physical capital, the price and wage Phillips curves (which are loaded by, respectively, the price and wage mark-ups), and the exogenous, stochastic processes for the structural shocks hitting the economic system.

The model is closed by 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 that policymakers react to is the output gap, defined as the 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, movements which happen gradually given the presence of interest rate smoothing. Departures from the Taylor rate (i.e., the rate that would be realized in the absence of any policy rate shocks), are triggered by a stochastic AR(1) process.

Monetary policy shocks in this model work as negative demand shocks. A temporary, unexpected hike in the policy rate triggers an increase in the real ex-ante interest rate. Given the increased opportunity costs of today's consumption, this leads to an increase in consumers' incentives to postpone their consumption. At the same time, Tobin's q decreases, and this induces a reduction in productive investment and, consequently, physical capital. Overall, aggregate demand weakens, and aggregate output falls. This puts downward pressure on real wages, which fall as well. This fall, along with the one in the rental rate of capital driven by weak demand, lead to a fall in firms' marginal costs. Hence, firms work in order to adjust downward their prices and maintain them at their target level. However, because of price stickiness, just a fraction of such firms is able to do it right after the shock. Hence, some firms experience a temporarily high mark-up. On aggregate, prices (and, therefore, inflation) take a value temporarily higher than the one that would be realized in the presence of fully flexible prices (but lower than the pre-shock one). In equilibrium, a temporary monetary policy shock triggers a fall in output and inflation. Due to the persistence of the monetary policy shock as well as the endogenous transmission mechanism which also features elements of persistence (e.g., habit formation in consumption, capital adjustment costs, price and wage indexation), the effects of an unexpected increase in the policy rate are persistent as well. However, as time goes by, more and more firms manage to adjust their plans and align their control variables with their optimal values in a fully flexible prices world. In the long run, output, inflation and the policy rate go back to their pre-shock levels.

3.2 Model estimation

We estimate the Smets and Wouters (2007) model with Bayesian methods (see, among others, An and Schorfheide (2007), Canova (2007)). We work with quarterly U.S. data for the period 1984:I-2008:II, which roughly coincides with the great moderation period (McConnell and Perez-Quiros (2000)). Our sample excludes the acceleration of the financial crises, which led the Federal Reserve to implement non-conventional monetary policy moves that are difficult to model with simple policy rules like the one embedded in our framework. This sample choice enables us to control for variations in inflation dynamics which occurred in post-WWII, which might be due to different monetary policy regimes being in place (D'Agostino and Surico (2012)). In addition, it allows us to control for the possible time-varying role played by monetary aggregates in shaping U.S. output (Castelnuovo (2012a)), and for the impact of breaks in the reduced-form representation of the VARs due to variations in the U.S. systematic monetary policy conduct (Bagliano and Favero (1998), Castelnuovo and Surico (2010)). Importantly, this sample is the one employed to estimate our VAR and the impulse responses which motivate this study (see our Figure 1, bottom panel).

To estimate our model, we employ an updated version of the seven observables used in the original paper by Smets and Wouters (2007). In particular, we employ the quarterly growth rates of real GDP, real consumption, real investment, the real wage, log hours worked, the quarterly inflation rate computed with the GDP deflator, and the federal funds rate. The source of the data is the Federal Reserve Bank of St. Louis' website. Following Smets and Wouters (2007), we implement the theoretical restriction implied by the common quarterly trend growth rate of the labor-augmenting technological process on the growth rate of four variables: real GDP, real consumption, real investment, and real wages. The set of measurement equations employed for the estimation of the model is proposed in the Appendix, where the interested reader can also find further details on the estimation procedure. Finally, some of the parameters are difficult to estimate, due to identification issues which are well known for dynamic-stochastic rational-expectations models like this one (Canova and Sala (2009)). As in Smets and Wouters (2007), we then set the depreciation rate $\delta = 0.025$; the exogenous spending-GDP ratio $g_y = 0.18$; the steady-state mark-up in the labor market $\lambda_w = 1.5$; the curvature parameters of the aggregators in the goods and labor market ε_p and $\varepsilon_w = 10$.

Tables 1 and 2 report our prior densities, which are carefully discussed in the original Smets and Wouters (2007) paper, as well as some selected moments and percentiles of our posterior densities. Our results are in line with most of the literature focusing on the estimation of DSGE models with great moderation data for the U.S. economy. In particular, we find a strong systematic policy reaction to inflation, a mild reaction to the model-consistent output gap, and a slightly stronger response 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, even when considering the MA(1) component of these last two shocks.

4 DSGE vs. Cholesky-VARs: A Monte Carlo exercise

We now turn to our Monte Carlo exercise. We aim at comparing the DSGE-consistent impulse responses to those produced with a VAR in which the monetary policy shock is identified by a Cholesky decomposition of the variance-covariance matrix of the reduced-form residuals. As is typically done in this literature, 'slow-moving' variables such as inflation and output are ordered before the policy rate. This implies that (i) monetary policy shocks cannot hit inflation and output in the very short run (contemporaneously); and (ii) shocks to inflation and output enter the interest rate equation contemporaneously.

4.1 The Monte Carlo algorithm

We calibrate the Smets and Wouters (2007) model with our estimated posterior means. Then, we produce the DSGE model-consistent (DGP-consistent) responses of inflation and output to a monetary policy shock (normalized to have a 25 basis-point on-impact effect on the federal funds rate). Hence, for $k = 1$ to K:

1. we produce our artificial dataset $\boldsymbol{x}^k_{[3xT]}$ (where $\boldsymbol{x}^k_{[3xT]}$ stands for the matrix containing information on three artificial time series of dimension T) with our DGP by sampling realizations from the mutually independent densities of all our structural shocks in each given period. A burn-in of 400 observations is considered;

2. we estimate a reduced-form VAR(3) and identify the monetary policy shock by imposing a Cholesky-structure on the short-run mapping going from the reduced-form residuals to the structural shocks of the VAR (the ordering being inflation, output gap, nominal rate);[5]

[5] The choice of using three lags in our baseline exercise is due to our willingness to keep our Monte

3. we compute the impulse responses suggested by our Cholesky-VAR to a normalized monetary policy shock;

4. we store the results obtained at step 3) in a $[3xHxK]$ $\boldsymbol{Cholesky-VAR\_IRFs}$ matrix, the content of which will be subsequently used to plot the Cholesky-VAR mean responses and the surrounding uncertainty.

We run this algorithm by setting the number of repetitions as $K = 5,000$, the horizon of the impulse response functions as $H = 15$, and the length of the artificial data sample as $T = 98$. 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 Cholesky-VAR for which 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.

4.2 Monte Carlo results

Empirical evidence. Figure 2 contrasts the impulse responses obtained with the DSGE model with those generated with our Cholesky-VARs. The estimated DSGE model predicts 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 suggests that U.S. monetary policy shocks do affect inflation and output.

A dramatically different picture arises when turning to our Cholesky-VARs. On average, our Cholesky-VARs return modest and insignificant inflation and output responses to a monetary policy shock. The similarity between the Cholesky-VAR responses shown in Figure 2 and those reported in Figure 1 (bottom panels) is substantial. In both cases, a monetary policy 'shock' identified with the Cholesky-scheme induces modest inflation and output responses. Consequently, our Monte Carlo evidence suggests that

Carlo VARs as close as possible to the ones estimated with actual U.S. data via which we produced the impulse responses proposed in Figure 1. Our results are robust to perturbations of the number of lags of our Monte Carlo VARs (discussion in Section 4.2). Notice that if we were interested in getting results in population we could skip the Monte Carlo experiments. In fact, we could compute autocovariances from our linear DSGE models. Then, we could compute VAR population coefficients. From these we could compute impulse responses and the composition of the VAR monetary policy shock in terms of the structural shocks. However, the evidence in Figure 1 is obtained by appealing to short-samples and with unrestricted VARs. Our Monte Carlo exercises allow us to replicate what a Cholesky-VAR econometrician typically does when handling time-series to investigate the macroeconomic effects of a monetary policy shock.

the mild-to-muted reactions reported in Figure 1 may very well be an *artifact* due to the imposition of the (wrong) Cholesky identification scheme. To reiterate, *mild-to-muted Cholesky-VAR responses* to a (misspecified) monetary policy shock are in principle *fully consistent* with a monetary policy actually *able to affect* the macroeconomic environment.

Our Monte Carlo exercise uncovers the difficulties by Cholesky-VARs to replicate the actual impulse responses to a monetary policy shock when a framework like Smets and Wouters (2007) is used as a DGP. It is well known that, if the underlying framework does not feature a recursive structure, a Cholesky-identification scheme might not deliver correct impulse responses (Canova and Pina (2005), Carlstrom, Fuerst, and Paustian (2009)). Our exercise adds knowledge to the existing literature by employing a DGP which has been shown to be empirically informative in a variety of contexts, including forecasting horse-races involving several Bayesian-VARs (Smets and Wouters (2007)) and the assessment of the macroeconomic effects of some key macroeconomic shocks (Del Negro, Schorfheide, Smets, and Wouters (2007)).

Interpretation. As stressed in the Introduction, if a recursive scheme is wrongly imposed on data generated by a non-recursive DSGE model, the impulse responses may end up being severely distorted because of the inability of the Cholesky-VAR framework to correctly discriminate between the macroeconomic effects of monetary vs. non-monetary policy shocks. To unveil the contribution of each structural disturbance to the amalgamated Cholesky-VAR 'shock', we compute the correlation between our VAR monetary policy shocks and the structural shocks in our DSGE framework. Figure 3 plots the distributions of such correlations. As is evident, a number of shocks enter the linear combination forming the Cholesky-VAR monetary policy 'shock'. Notice the role played by the TFP shock. The loading of the TFP shock is, on average, negative. Given the nature of this shock, which is a supply shock, the response of inflation is dampened by negative realizations of TFP disturbances. On the other hand, a negative TFP shock should reinforce the negative reaction of output in the short run. However, three 'demand' shocks (risk-premium, Government spending, and investment-specific shocks), correlate positively with the monetary policy shock, so contrasting the effects of a monetary policy shock on inflation and output. The correlation between the Cholesky-VAR and the DSGE monetary policy shocks is positive and large, as expected. As far as our mark-up shocks are concerned, the price mark-up shock correlates negatively with the Cholesky-VAR monetary policy shock, therefore contrasting the effect of a genuine monetary policy tightening on output and reinforcing that on inflation. In contrast, the wage mark-up shock displays a weak correlation with the Cholesky-VAR monetary policy shock (on average). The Cholesky-VAR monetary policy 'shock' represents a linear combination of a variety of structural shocks, their joint effect being that of dampening inflation and output responses to a genuine monetary policy shock.

Robustness checks. Our exercise assumes our econometrician is endowed with the correct knowledge of the DSGE-consistent output gap. We share this assumption with Carlstrom, Fuerst, and Paustian (2009), who work with a three-equation DSGE framework and document the performance of Cholesky-VARs under different scenarios. This assumption is, however, unpalatable, because the true potential output process is hardly known by the econometrician. We then relax this assumption by substituting the output gap with the output growth rate. In particular, we (i) re-estimate our VARs with actual U.S. data in order to understand if the modest responses in Figure 1 (bottom panels) stay the same when this alternative vector is considered; and (ii) re-do our Monte Carlo simulations accordingly. Our results turn out to be robust to the employment of the output growth rate.

Another issue relates to our approximation of the dynamics of inflation, output, and the policy rate. Our VAR(3) approximation may not be adequate to capture the VARMA structure of the Smets and Wouters (2007) model. We then re-run our Monte Carlo exercise by employing a VAR(10). If some relevant dynamics are not well-captured by our VAR(3), the impulse responses conditional on VAR(10)-models should differ from our baseline responses. Clearly, also the small-sample bias may be an issue, because 98 observations are not necessarily enough to correctly characterize the dynamics of the macroeconomic series under consideration. Again, if the small-sample bias is a serious issue for our exercise, our impulse responses should look quite different conditional on VARs estimated with much larger samples. We then re-run our Monte Carlo simulations by endowing the econometrician with 1,000 observations for each artificial dataset. Our results turn out to be robust to these additional robustness checks. All there robustness checks are documented in our Appendix, which is available upon request.

5 Behind the scenes: Possible drivers at work

The results of our Monte Carlo exercise presented in Figure 2 mimic the conditional correlations obtained with actual U.S. data and shown in Figure 1 (bottom panels). Hence, the Smets and Wouters model can be fruitfully employed to identify the reasons behind our 'mild-to-muted' macroeconomic reactions presented in Figure 1. As pointed out in our Introduction, different explanations have been proposed for such reactions. They mainly relate to changes in the policy parameters, variations in the parameters strictly related to the private sector, and the volatility of more moderate structural shocks.

To shed light on the relative importance of the different subset of parameters in this model, we partition the vector Δ of structural parameters of our DSGE model as follows: $\Delta = [\Delta^{POLICY}, \Delta^{PS}, \Delta^{ARMA}, \Delta^{SHOCKS}]$, where the vector $\Delta^{POLICY} = [r_\pi, \rho, r_y, r_{\Delta y}]$ collects the policy parameters; the set $\Delta^{PS} = [\varphi, \sigma_c, h, \xi_w, \sigma_l, \xi_p, \iota_w, \iota_p, \psi, \Phi]$ refers to the parameters for the private sector; $\Delta^{ARMA} = [\rho_a, \rho_b, \rho_g, \rho_I, \rho_r, \rho_p, \rho_w, \mu_p, \mu_w, \rho_{ga}]$ collects the autoregressive parameters for the ARMA processes of the model; and $\Delta^{SHOCKS} = [\sigma_a, \sigma_b, \sigma_g, \sigma_I, \sigma_r, \sigma_p, \sigma_w]$ takes the volatilities of the shock into account. Different combinations of these sets of parameters are intended to represent different scenarios.

To calibrate the changes in the structural parameters, we re-estimate the Smets and Wouters (2007) model with the sample 1960Q1-1979Q3 to capture the macroeconomic dynamics of the 1960s and 1970s. Following most of the literature, we do not model the late 1970s/early 1980s in order to not contaminate our estimates with observations referring to the 'Volcker experiment', which is the phase during which the Federal Reserve targeted non-borrowed reserves. Tables 1 and 2 document our estimates. In line with some previous contributions (Clarida, Galí, and Gertler (2000), Justiniano and Primiceri (2008)), we find evidence in favor of weaker systematic policy conduct and larger shocks in the 1960s and 1970s. The parameters for the private sector also turn out to be different in the two subsamples.

We then run counterfactual scenarios featuring the following combinations of parameters:

- *Monetary policy.* This scenario plants the policy regime of the 1960s and 1970s in the post-1984 era by combining sets of parameters as follows: $\Delta^{MONETARY\_POLICY} = [\Delta^{POLICY}_{66Q1-79Q3}, \Delta^{PS}_{84Q1-08Q2}, \Delta^{ARMA}_{84Q1-08Q2}, \Delta^{SHOCKS}_{84Q1-08Q2}]$, where $\Delta^{i}_{66Q1-79Q3}$ stands for a 'set of parameters i conditional on the posterior means obtained with 1966Q1-1979Q3 data', and $\Delta^{i}_{84Q1-08Q2}$ stands for a 'set of parameters i conditional on the posterior means obtained with 1984Q1-2008Q2 data'. If the aggressive systematic monetary policy which took place in 1984Q1-2008Q2 is the explanation for the modest macroeconomic reactions to a monetary policy shock, this 'monetary policy' scenario should be associated with macroeconomic responses larger than those shown in Figure 2. Notice that, following most of the literature, here we admit a single break in the policy parameters. A richer description of the U.S. monetary policy regimes in the post-WWII period can be found in Sims and Zha (2006), Canova (2009), and Bianchi (2013).

- *Private sector.* This scenario is simulated by combining the following subset of parameters: $\Delta^{PRIVATE\_SECTOR} = [\Delta^{POLICY}_{84Q1-08Q2}, \Delta^{PS}_{66Q1-79Q3}, \Delta^{ARMA}_{84Q1-08Q2}, \Delta^{SHOCKS}_{84Q1-08Q2}]$. One explanation proposed in the literature for the modest macroeconomic effects documented in the Introduction refers to a number of financial innovations which may have eased households' consumption smoothing since the early 1980s. Smets and Wouters' (2007) model is ill-suited to structurally capture the impact of such innovations, as the model does not embed financial constraints or the possibility of a credit crunch. If this explanation is correct, the data associated with the 1970s should be described via a different model (possibly encompassing the Smets and Wouters (2007) framework), in which these financial frictions are explicitly modeled. However, if financial innovations have led to changes in households' and firms' behavior, such changes should be reflected in the estimates of the parameters for the private sector (which are likely to be reduced-form coefficients, according to this interpretation).

- *ARMA processes.* This scenario is characterized by the following sets of parameters: $\Delta^{ARMA\_PROCESSES} = [\Delta^{POLICY}_{84Q1-08Q2}, \Delta^{PS}_{84Q1-08Q2}, \Delta^{ARMA}_{66Q1-79Q3}, \Delta^{SHOCKS}_{84Q1-08Q2}]$. This scenario is simulated to capture the role played by a set of autoregressive parameters in shaping the macroeconomic responses to a monetary policy shock in Cholesky-VARs like the ones we are focusing on. Carlstrom, Fuerst, and Paustian (2009) provide closed-form solutions for the three-equation DSGE model case which suggest that the impact of the calibration of the autoregressive parameters for the structural shocks of the economic system may be substantial.

- *Shocks.* Structural shocks (in particular, their variance) are another possible explanation for the moderate responses obtained with our Cholesky VARs. We investigate to what extent the advent of the great moderation may have implied modest macroeconomic reactions in the VAR analysis by simulating the following scenario: $\Delta^{SHOCKS} = [\Delta^{POLICY}_{84Q1-08Q2}, \Delta^{PS}_{84Q1-08Q2}, \Delta^{ARMA}_{84Q1-08Q2}, \Delta^{SHOCKS}_{66Q1-79Q3}]$.

Figure 4 plots the dynamic responses to a monetary policy shock in our baseline case and in our four counterfactual scenarios. The change in structural parameters appears to be the most influential driver behind our Cholesky-VAR responses. According to our simulations, if the parameters for the private sector had not changed when moving from the great inflation period to the great moderation phase, a Cholesky-VAR analysis would have been more likely to predict a recession after a monetary policy shock. Qualitatively, we obtain a similar results for inflation. However, the quantitative impact of the change in private sector parameters on the response of inflation turns out to be much milder. In contrast, the persistence coefficients of the ARMA processes and the magnitude of the structural shocks seem to play a negligible role for our responses. Finally, according to our simulations, a weaker monetary policy would have induced, *ceteris paribus*, a milder reaction by our Cholesky-VARs. Hence, our simulations show that (i) the change in parameters associated with households' and firms' problems may be one of the factors pushing the responses of output and, to a lesser extent, inflation toward zero; and (ii) the aggressive monetary policy implemented by the Federal Reserve since the early 1980s is unlikely to be a major factor behind such modest reactions.

A discussion on the exercise conducted above is in order. Our estimates of the Taylor rule response to inflation in the 1966Q1-1979Q3 sample imply fairly aggressive behavior by the Federal Reserve. While being consistent with the subsample estimate proposed by Smets and Wouters (2007), this result stands in contrast with the evidence in favor of a passive monetary policy in the pre-Volcker period proposed by, among others, Clarida, Galí, and Gertler (2000), Lubik and Schorfheide (2004), Boivin and Giannoni (2006), and Mavroeidis (2010). Most likely, this is due to the fact that, following most of the extant empirical DSGE literature, our estimation assumes a unique rational expectations equilibrium to be in place (i.e., it does not allow for indeterminacy).[6] Indeterminacy is typically associated with inefficient fluctuations driven by sunspot shocks, which might be one of the drivers of the larger fluctuations of the 1970s. Hence, admitting for a passive monetary policy could unveil a larger role played by the switch in systematic monetary policy which possibly occurred at the end of the 1970s. Moreover, Castelnuovo and Surico (2010) conduct Monte Carlo simulations with small-scale, three equation-AD/AS models à la Clarida, Galí, and Gertler (2000) and show that passive monetary policy is an important element in replicating a VAR regularity associated with the 1970s (i.e., the price puzzle). In spite of this evidence, we decided not to

[6]Exceptions of system-estimations allowing for indeterminacy are Lubik and Schorfheide (2004), Boivin and Giannoni (2006), Justiniano and Primiceri (2008), Benati and Surico (2009), and Castelnuovo and Fanelli (2015).

model indeterminacy for the following reasons. First, the empirical evidence pointing to indeterminacy/passive monetary policy becomes much less clear when one focuses on medium scale models. For instance, Justiniano and Primiceri (2008) estimate several versions of a medium-scale model à la Smets and Wouters (2007), one of which allows for a switch from a passive to an active policy conduct in the first part of the 1980s. They find that the specification where policy remains always active outperforms the one in which the policy break is modeled. Second, the results obtained with simulations conducted with DSGE models under indeterminacy may heavily depend on the way in which the macro-modeler picks a single equilibrium out of the many (possibly infinite) equilibria that are consistent with rational expectations (Castelnuovo (2012b)). Third, the uni-equational, Taylor rule-based evidence on passive monetary policy is not robust to the employment of real-time data, which are likely to better represent the information set available to policymakers when setting monetary policy (Orphanides (2001), Coibion and Gorodnichenko (2011), and Coibion and Gorodnichenko (2012)).[7] However, we admit that our simulations could be overly conservative regarding the role that systematic monetary policy may have played in affecting the dynamics of the U.S. economic system in response to a monetary policy shock.

Another caveat associated with our counterfactual exercises is the (commonly entertained) assumption of mutual independence of the estimated parameters. This practice, however, neglects the empirical fact that estimated parameters in DSGE models are often non-independent. Hence, the set of parameters characterizing our counterfactual scenarios might be associated with a very low likelihood (Canova (2009)).

Finally, our model features a Ricardian fiscal policy. Leeper (1991) has shown that, in the context of modern new-Keynesian frameworks, different fiscal policies may very well have a different impact on the equilibrium values of inflation and output. Bianchi (2012) finds support for the role of fiscal policy as one of the factors behind the run-up of U.S. inflation during the 1970s. This might explain why Smets and Wouters' (2007) model is better-equipped to replicate the Cholesky-VAR correlations during the Great Moderation than those conditional on data of the 1960s and 1970s (results are shown in the Appendix). We leave the investigation of this conjecture to future research.

[7] To be precise, real-time data are likely to be more informative for the information set available to policymakers when making their policy decisions. It may very well be that, in spite of an intended aggressive monetary policy behavior, such behavior was, in fact, passive. This might be due to imprecise estimates about the state of the economy available to policymakers when making their decisions, or misspecification of the reference transmission mechanism. See Orphanides (2001) for a discussion on this issue.

6 An alternative identification strategy: Sign restrictions

Our analysis has focused on Cholesky-VARs so far. Cholesky-VARs impose exclusion restrictions which are inconsistent with the contemporaneous non-zero responses of output and inflation to a monetary policy shock in our DGP. A way to overcome this inconsistency is to appeal to an alternative identification strategy: sign restrictions. In a nutshell, the idea is to impose sign restrictions on a set of moments generated with the VAR (e.g., a set of impulse responses to a given shock). In our application, we estimate the reduced-form VAR coefficients $\boldsymbol{A}(L)$ and covariance matrix $\boldsymbol{\Lambda}$ from the data via OLS.[8] Following Canova and de Nicoló (2002), we orthogonalize the VAR residuals via an eigenvalue-eigenvector decomposition such that $\boldsymbol{\Lambda} = \boldsymbol{P}\boldsymbol{H}\boldsymbol{P}^T$, where $\boldsymbol{P}$ is the matrix of eigenvectors and $\boldsymbol{H}$ is the diagonal matrix of eigenvalues. The non-uniqueness of the MA representation of the VAR is exploited to provide a set of alternative proposals for the shock(s) of interest via the employment of three Givens rotation matrixes $\boldsymbol{Q}_{ij}(\omega)$, where i and j indicate that rows i and j are rotated by an angle $\omega \in (0, \pi/2)$, $\boldsymbol{R} = \boldsymbol{Q}_{12}(\omega_1)\boldsymbol{Q}_{13}(\omega_2)\boldsymbol{Q}_{23}(\omega_3)$, and $\boldsymbol{R}\boldsymbol{R}^T = \boldsymbol{I}_3$. The 'impulse' matrix loading the VAR with candidate 'shocks' is therefore given by $\widetilde{\boldsymbol{B}}(\boldsymbol{\omega}) = \boldsymbol{P}\boldsymbol{D}^{1/2}\boldsymbol{R}(\boldsymbol{\omega})$. If the impulse responses to the 'candidate' shock satisfy all the required restrictions, then the draw of the vector $\boldsymbol{\omega}$ and the corresponding responses are retained. Otherwise, they are discarded. In so doing, we assign equal and strictly positive weight to the draws we retain (those that meet our restrictions), and assign zero prior weight to those that violate our constraints.[9] A non-exhaustive list of recent applications of the sign-restriction strategy to identify structural shocks includes Faust (1998), Canova and de Nicoló (2002), Peersman (2005), and Uhlig (2005). Rubio-Ramírez, Waggoner, and Zha (2010) and Arias, Rubio-Ramírez, and Waggoner (2014) propose algorithms to efficiently compute rotations of the impulse matrix. Such an algorithm also works well when the number of variables in the vector is large and several restrictions are imposed to identify more than one structural shock. Canova and Paustian (2011) propose an algorithm which derives a set of 'robust' restrictions from a class of structural DSGE

[8] We employ a frequentist approach to obtain prior-free results. The key question of this paper relates to the ability of an agnostic identification procedure imposing no constraints on output to correctly recover the response of such a variable to a monetary policy shock. The employment of Monte Carlo simulations enables us to control for omitted factors which might otherwise render our results difficult to interpret.

[9] In contrast, one could set up a penalty function to penalize violations and reward large and correct responses. For an in-depth discussion, see Uhlig (2005).

models that one may exploit to identify the shock(s) of interest with VARs. Fry and Pagan (2011) critically review the estimation of structural VARs with sign restrictions.

We identify the monetary policy shock by imposing 'textbook' constraints on the impulse responses of inflation and the policy rate to a monetary policy shock. In particular, it is worth noticing that, conditional on the Smets and Wouters (2007) model, the only shock which is able to generate a short-run negative correlation between inflation and the policy rate is exactly the monetary policy shock. This correlation is consistent not only with the specific model we employ in this paper (more on this later), but also with a *class* of differently calibrated small-scale models (Canova and Paustian (2011)). Hence, we impose a positive (negative) response of the policy rate and a negative (positive) response of inflation and output to an unexpected temporary hike (drop) in the short-term rate.[10]

Figure 5 displays the impulse responses to a monetary policy shock identified with these restrictions.[11] A few comments are in order. On top of getting the on-impact response of all three variables right (given the signs we imposed, this is an assumption by the econometrician, not a result driven by the data), sign restrictions deliver responses which are correctly persistent and their shapes are close to the actual responses. However, the magnitude of the VAR responses is clearly disappointing, with an incredibly deep recession and a strong disinflation associated with the monetary policy shock. While this result is clearly specific to the DGP chosen for this exercise, it is of interest to compare these responses to those one could obtain if the DGP featured more volatile monetary policy shocks. Then, we conduct two further simulations by first doubling the standard deviation of the monetary policy shock, and then quintupling it. To ease comparison, we also plot these two sets of responses in Figure 5. The performance of the sign restriction-VAR is found to dramatically improve in the presence of more volatile monetary policy shocks (all else being equal). Indeed, the VAR responses conditional on the scenario characterized by the quintupled standard deviation of the monetary

[10] Experiments conducted by imposing additional restrictions to identify extra structural shocks delivered the same dynamics of the modeled variables. Notice that, given the linearity of the model, we can identify the effects of a policy hike by flipping the signs of the response of the policy rate, inflation, and the output gap if we obtain a draw such that the policy rate reacts negatively, inflation reacts positively, and the output gap reacts positively on impact.

[11] Such responses are often reported as pointwise median or mean values computed across all the retained rotations. But, as pointed out by Fry and Pagan (2011), each retained rotation is a model, and indeed sign restrictions identify sets of models. Hence, pointwise median or mean values are unlikely to be associated with a single existing model. Following Fry and Pagan (2011), we then report the impulse responses conditional on a single model, which is the one associated to the responses which are found to be the closest to the pointwise median responses.

policy shock track the actual responses very closely.

Wrapping up, sign restriction-VARs are found to record a performance which depends on the size of the monetary policy shock. Such size is important because it regulates, all else being equal, the relative contribution of monetary policy shocks in explaining the volatility of the macroeconomic variables modeled with the DSGE model and the VAR. Paustian (2007) analytically shows that sign restrictions 'work well' when the shock the econometrician aims at identifying is 'powerful enough' (i.e., it explains a large-enough share of the volatility of the variable whose impulse response one is investigating). Canova and Paustian (2011) and Castelnuovo (2015) propose simulations that quantitatively confirm the relevance of Paustian's (2007) result. Conditional on our estimated DGP, the share of unconditional variance of the output gap, inflation, and the policy rate explained by the monetary policy shock is just, respectively, 3%, 6%, and 6%. Hence, the estimated model suggests that the effects of a monetary policy shock are difficult to identify with sign restrictions. Things change when the standard deviation of the monetary policy shock is (counterfactually) doubled, an operation conducted by keeping the remaining calibration of the model fixed. In this case, the contribution of the policy shock increases to 10% for the output gap, and 20% for inflation and the policy rate. In a world in which the standard deviation of the policy shock in our DGP is assumed to be five times larger than what the data suggest, such shares increases to 40% for the output gap, and 62% for inflation and the policy rate. But these alternative calibrations for the standard deviation of the monetary policy shock in our DSGE model fall clearly outside the estimated 90% credible set for such parameters (i.e., they are obviously counterfactual). Hence, our conclusion is that, if one is willing to use a version of the Smets and Wouters (2007) model estimated over the 1984-2008 period as DGP, sign restrictions *per se* do not necessarily represent an improvement over the Cholesky-VAR model.[12]

The findings illustrated and discussed above relate to the paper by Uhlig (2005), in which he also finds modest real effects of monetary policy shocks when identifying such shocks with sign restrictions. Notice that Uhlig (2005) does not impose any sign on the response of output in order to be 'agnostic' with regards to the real effects of monetary policy shocks. In contrast, in our exercise we assume a short-run negative response of output to an unexpected monetary policy tightening. Our Appendix shows

[12] Of course, one could play with extra restrictions such as magnitude restrictions on the values of the impulse responses to some of the modeled variables, restrictions on the VAR parameters, and so on. See Arias, Caldara, and Rubio-Ramírez (2014) and Castelnuovo (2015) for investigations along these lines.

that when relaxing such a constraint we basically replicate Uhlig's (2005) result even in presence of a DGP à la Smets and Wouters (2007). In particular, the response of output, which is positive conditional on the estimated DGP, changes sign and gets much closer to the actual response when we counterfactually inflate the standard deviation of the monetary policy shock. Hence, one possible interpretation of Uhlig's result is, again, the inability of sign restrictions to correctly identify a monetary policy shock when such a shock explains a small share of the variance of the variable whose impulse response one is investigating.

7 Conclusions

This paper shows that evidence of modest macroeconomic responses to monetary policy shocks produced with Cholesky-VARs that are estimated with U.S. great moderation data may be consistent with substantial macroeconomic effects triggered by unexpected moves in the policy nominal interest rate. We make this point by proceeding in two steps. First, we estimate a medium-scale new-Keynesian DSGE model with Bayesian techniques, and verify that it predicts a persistent reaction of inflation and output to an unexpected monetary policy tightening. Then, we set up a Monte Carlo experiment in which we feed Cholesky-VARs with artificial data generated by our estimated new-Keynesian framework. We show that Cholesky-VARs return, on average, modest responses of inflation and output. Our result is due to the mismatch existing between the timing implied by the Cholesky-identification scheme and the one suggested by our non-recursive Data-Generating Process. As a consequence, various shocks are shown to contaminate the actual effects of monetary policy shocks on output and inflation. The calibration of parameters related to households' and firms' structural equations is shown to be possibly more important than that of the monetary policy rule in understanding our results. An investigation conducted with an alternative identification strategy based on signs imposed on the on-impact responses of output, inflation, and the nominal interest rate to a monetary policy shock (i.e., sign restrictions) reveal both the potential gains coming from this alternative strategy and the associated empirical pitfalls.

What 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 Fernández-Villaverde, Rubio-Ramírez, Sargent, and Watson (2007) state (p. 1025), '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 the evidence on the macroeconomic responses to a monetary policy shock identified with a standard recursive scheme should be interpreted with great care. DSGE models are likely to be misspecified. But, as shown by Del Negro and Schorfheide (2004) and Del Negro, Schorfheide, Smets, and Wouters (2007), 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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			1966Q1-1979Q3	1984Q1-2008Q2
Param.	*Interpretation*	*Priors*	*Post.Means* [5h,95th]	*Post.Means* [5h,95th]
φ	Capital adj. cost elast.	$\mathcal{N}(4, 1.5)$	3.92 [2.17,5.51]	6.06 [4.22,7.96]
σ_c	Risk aversion	$\mathcal{N}(1.5, 0.375)$	1.46 [1.13,1.76]	1.39 [1.16,1.62]
h	Habit formation	$\mathcal{B}(0.7, 0.1)$	0.63 [0.51,0.75]	0.63 [0.50,0.75]
ξ_w	Wage stickiness	$\mathcal{B}(0.5, 0.1)$	0.68 [0.57,0.80]	0.64 [0.49,0.79]
σ_l	Elast. lab. supply	$\mathcal{N}(2, 0.75)$	1.39 [0.38,2.31]	1.76 [0.78,2.74]
ξ_p	Price stickiness	$\mathcal{B}(0.5, 0.1)$	0.65 [0.56,0.74]	0.71 [0.62,0.80]
ι_w	Wage indexation	$\mathcal{B}(0.5, 0.15)$	0.58 [0.38,0.78]	0.52 [0.28,0.76]
ι_p	Price indexation	$\mathcal{B}(0.5, 0.15)$	0.41 [0.17,0.63]	0.40 [0.20,0.59]
ψ	Capacity utiliz. elast.	$\mathcal{B}(0.5, 0.15)$	0.53 [0.32,0.75]	0.69 [0.54,0.85]
$\Phi - 1$	Fixed prod. costs, share	$\mathcal{N}(0.25, 0.125)$	0.34 [0.20,0.48]	0.44 [0.30,0.57]
r_π	T. Rule, inflation	$\mathcal{N}(1.5, 0.25)$	1.51 [1.22,1.79]	2.10 [1.78,2.43]
ρ	T. Rule, inertia	$\mathcal{B}(0.75, 0.10)$	0.78 [0.71,0.85]	0.83 [0.80,0.87]
r_y	T. Rule, output gap	$\mathcal{N}(0.125, 0.05)$	0.15 [0.10,0.21]	0.05 [0.02,0.09]
$r_{\Delta y}$	T. Rule, output growth	$\mathcal{N}(0.125, 0.05)$	0.18 [0.13,0.24]	0.16 [0.11,0.20]
$\overline{\pi}$	St. state inflation rate	$\mathcal{G}(0.625, 0.10)$	0.72 [0.57,0.88]	0.64 [0.55,0.73]
$100(\beta^{-1} - 1)$	St. state interest rate	$\mathcal{G}(0.25, 0.10)$	0.25 [0.10,0.41]	0.25 [0.10,0.40]
$\bar{l}$	St. state hours worked	$\mathcal{N}(0, 2)$	-0.28 [-2.11,1.57]	0.87 [-0.84,2.57]
$\overline{\gamma}$	Trend growth rate	$\mathcal{N}(0.4, 0.1)$	0.34 [0.25,0.43]	0.42 [0.37,0.47]
α	Share of capital in prod.	$\mathcal{N}(0.3, 0.05)$	0.29 [0.22,0.37]	0.32 [0.25,0.39]

Table 1: **Bayesian estimates of the Smets and Wouters (2007) DSGE model - Structural Parameters.** Prior densities: Figures indicate the (mean,st.dev.) of each prior distribution. The letters 'N', 'B', 'G' stand for 'Normal', 'Beta', 'Gamma' densities. Posterior densities: Figures reported indicate the posterior mean and the [5th,95th] percentile of the estimated densities. Details on the estimation procedure are provided in the text.

			1966Q1-1979Q3	1984Q1-2008Q2
Param.	*Interpretation*	*Priors*	*Posterior Means* [5h,95th]	*Posterior Means* [5h,95th]
σ_a	TFP shock, st.dev.	$\mathcal{IG}(0.1, 2)$	0.61 [0.51,0.71]	0.41 [0.36,0.46]
σ_b	Risk-premium shock, st.dev.	$\mathcal{IG}(0.1, 2)$	0.25 [0.18,0.31]	0.16 [0.10,0.21]
σ_g	Gov. spending shock, st.dev.	$\mathcal{IG}(0.1, 2)$	0.59 [0.50,0.68]	0.41 [0.36,0.46]
σ_I	Invest.-specific tech. shock, st.dev.	$\mathcal{IG}(0.1, 2)$	0.57 [0.41,0.72]	0.35 [0.27,0.42]
σ_r	Mon. policy shock, st.dev.	$\mathcal{IG}(0.1, 2)$	0.20 [0.17,0.23]	0.12 [0.10,0.14]
σ_p	Price mark-up shock, st.dev.	$\mathcal{IG}(0.1, 2)$	0.22 [0.16,0.27]	0.10 [0.08,0.12]
σ_w	Wage mark-up shock, st.dev.	$\mathcal{IG}(0.1, 2)$	0.19 [0.14,0.25]	0.29 [0.23,0.35]
ρ_a	TFP shock, AR(1) coeff.	$\mathcal{B}(0.5, 0.2)$	0.98 [0.94,0.99]	0.95 [0.92,0.97]
ρ_b	Risk-premium shock, AR(1) coeff.	$\mathcal{B}(0.5, 0.2)$	0.32 [0.10,0.53]	0.32 [0.04,0.62]
ρ_g	Gov. sp. shock, AR(1) coeff.	$\mathcal{B}(0.5, 0.2)$	0.90 [0.85,0.95]	0.95 [0.93,0.97]
ρ_I	Inv.-spec. tech. shock, AR(1) coeff.	$\mathcal{B}(0.5, 0.2)$	0.64 [0.49,0.80]	0.74 [0.65,0.85]
ρ_r	Mon. pol. shock, AR(1) coeff.	$\mathcal{B}(0.5, 0.2)$	0.21 [0.06,0.35]	0.30 [0.15,0.44]
ρ_p	Price mark-up shock., AR(1) coeff.	$\mathcal{B}(0.5, 0.2)$	0.63 [0.34,0.90]	0.89 [0.81,0.98]
ρ_w	Wage mark-up shock, AR(1) coeff.	$\mathcal{B}(0.5, 0.2)$	0.93 [0.88,0.99]	0.93 [0.88,0.98]
μ_p	Price mark-up shock, MA(1) coeff.	$\mathcal{B}(0.5, 0.2)$	0.50 [0.24,0.78]	0.66 [0.47,0.85]
μ_w	Wage mark-up shock, MA(1) coeff.	$\mathcal{B}(0.5, 0.2)$	0.77 [0.57,0.95]	0.71 [0.53,0.88]
ρ_{ga}	Gov.spend.-TFP shocks, correl.	$\mathcal{B}(0.5, 0.2)$	0.61 [0.42,0.81]	0.44 [0.28,0.61]

Table 2: **Bayesian estimates of the Smets and Wouters (2007) DSGE model - Shock processes.** Prior densities: Figures indicate the (mean,st.dev.) of each prior distribution. The letters 'IG', 'B' stand for 'Inverse Gamma', 'Beta' densities. Posterior densities: Figures reported indicate the posterior mean and the [5th,95th] percentile of the estimated densities. Details on the estimation procedure are provided in the text.

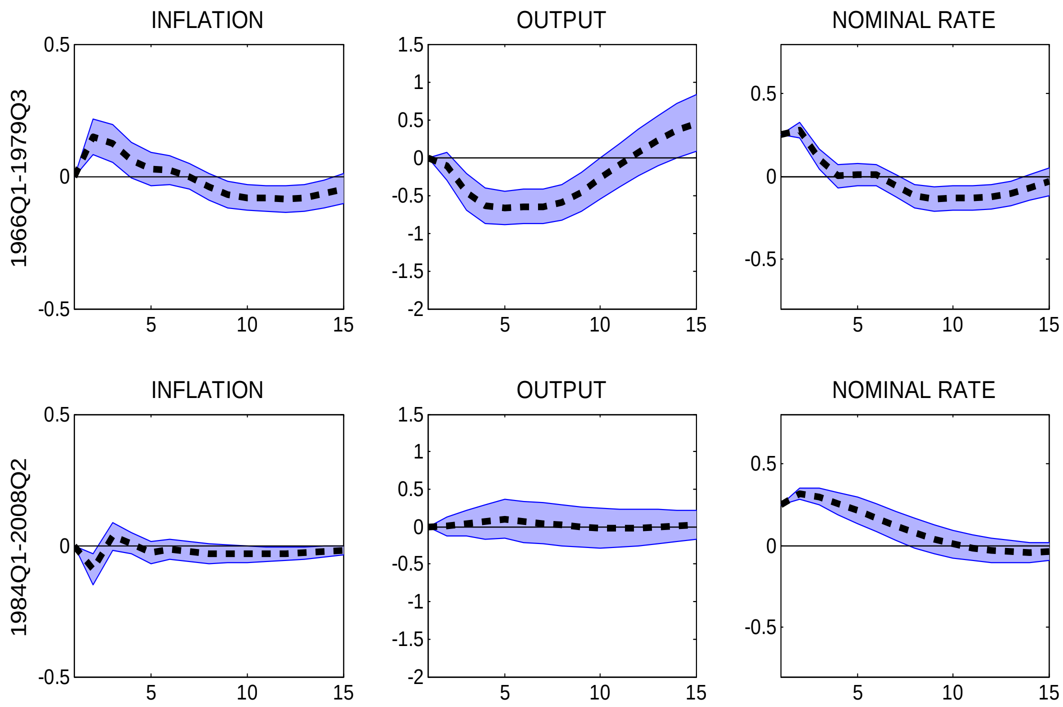

Figure 1: **VAR impulse response functions to a monetary policy shock. Actual U.S. data, Great Inflation vs. Great Moderation.** Variables: Quarterly GDP inflation, output gap, quarterly federal funds rate - source: FREDII. Identification of the monetary policy shock via Cholesky decomposition (lower triangular matrix, ordering: inflation, output gap, federal funds rate). Dashed black lines: Mean responses. Shaded areas: [16th,84th] percentiles (bootstrapped, 500 repetitions). VAR estimated with a constant, a linear trend, and three lags.

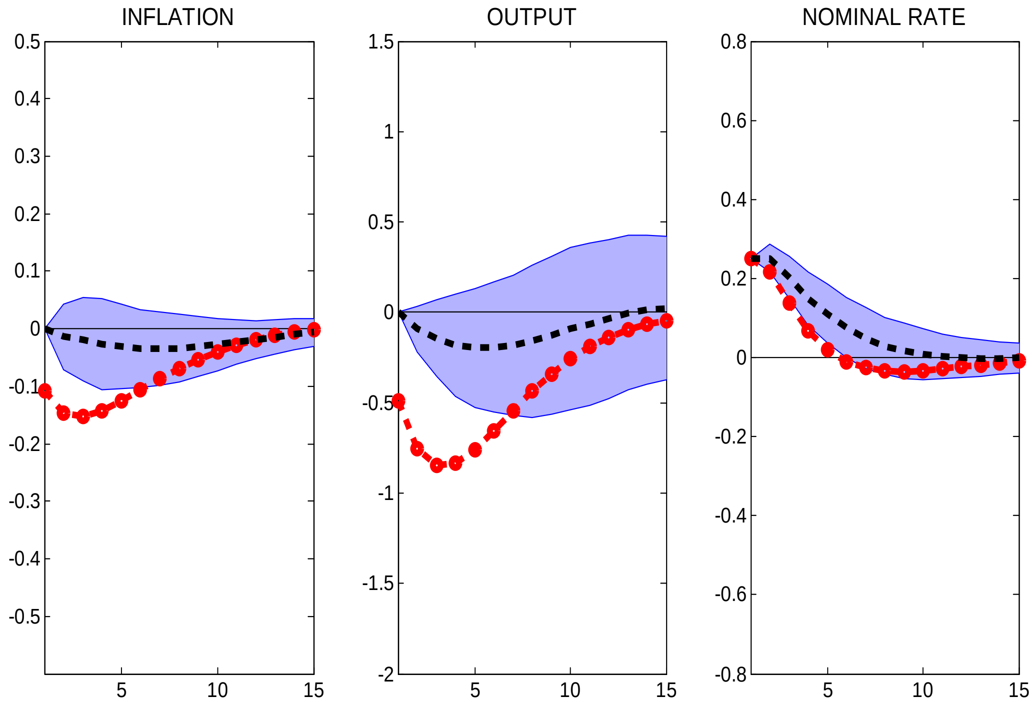

Figure 2: **Monte Carlo exercise, DSGE- vs. Cholesky-VAR-impulse response functions to a monetary policy shock.** Dashed-circled red lines: DSGE Bayesian mean impulse responses referring to the 1984Q1-2008Q2 period. Dashed black lines: Cholesky-VAR mean responses. Shaded areas: Cholesky-VAR responses, [16th,84th] percentiles. Simulations based on 1,000 repetitions of our Monte Carlo algorithm. Identification of the monetary policy shock via Cholesky decomposition (lower triangular matrix, ordering: inflation, output gap, nominal rate). VAR estimated with a constant and three lags.

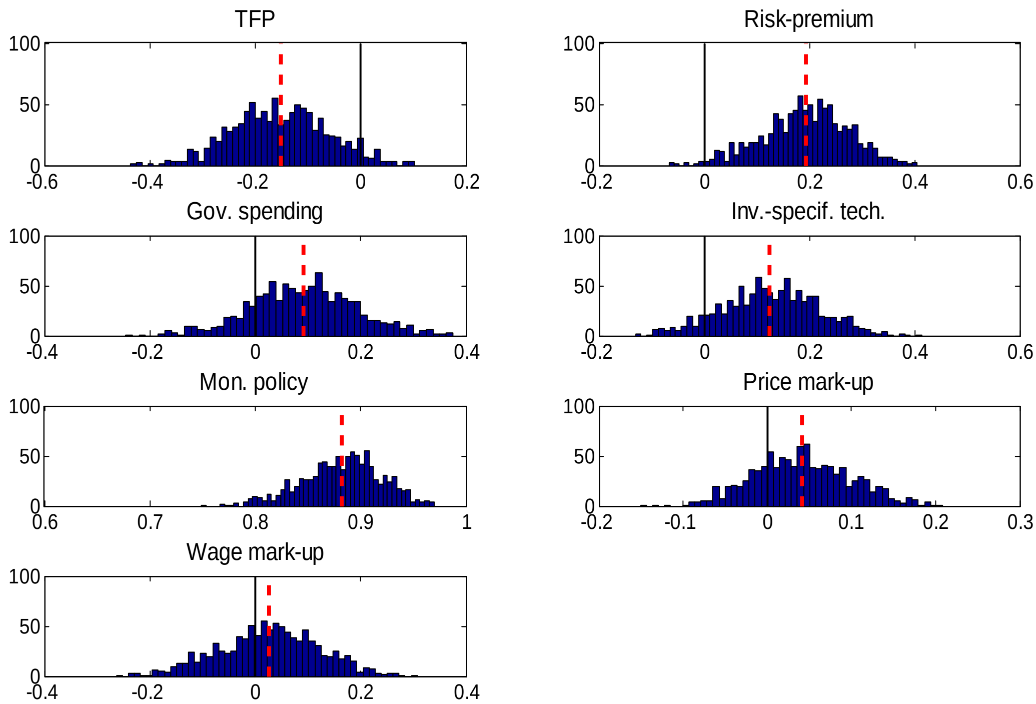

Figure 3: **Monte Carlo exercise, correlations involving Cholesky-VAR monetary policy shocks and DSGE shocks.** Empirical distributions based on 1,000 simulations. Calibration of the DSGE model referring to the 1984Q1-2008Q2 period.

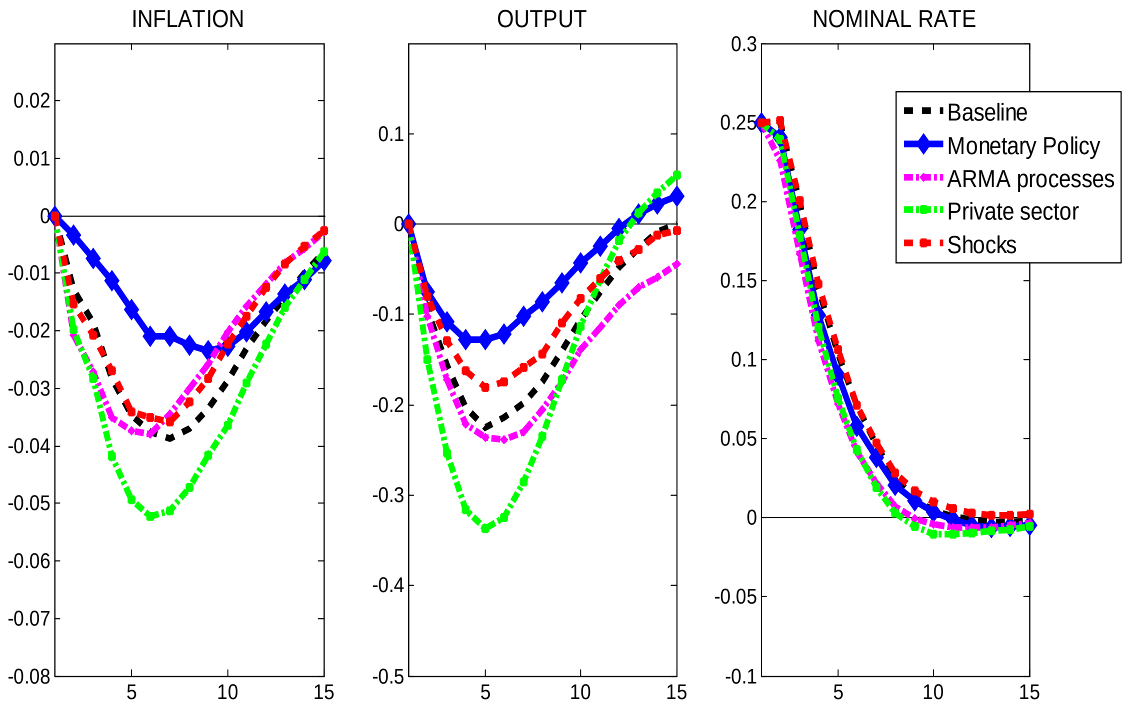

Figure 4: **Monte Carlo exercise, Cholesky-VAR impulse response functions: 'Counterfactual' exercises.** Black-dashes lines: Baseline exercise. Other scenarios: Counterfactual exercises in which we calibrated a subset of parameters with our 1966Q1-1979Q3 posterior means while keeping the remaining parameters calibrated as suggested by our baseline 1984Q1-2008Q2 posterior means. Scenarios described in the text. Simulations based on 1,000 repetitions of our Monte Carlo algorithm. Identification of the monetary policy shock via Cholesky decomposition (lower triangular matrix, ordering: inflation, output gap, nominal rate). VAR estimated with three lags.

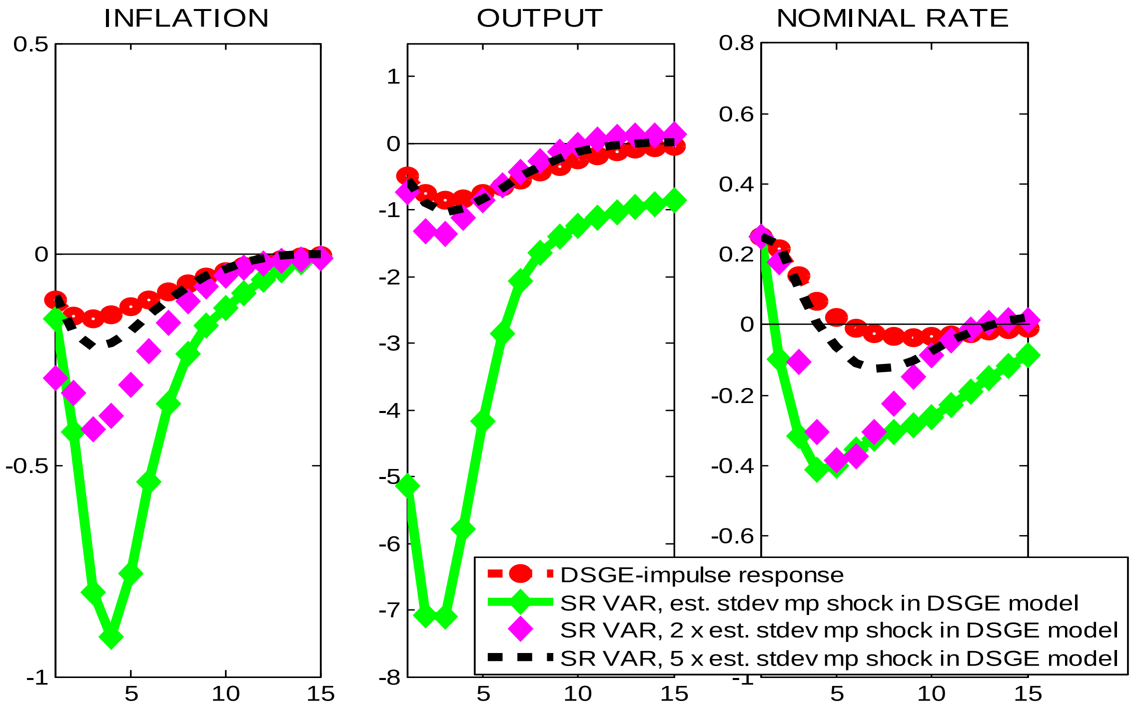

Figure 5: **Monte Carlo exercise, DSGE- vs. Sign Restriction-VAR-impulse response functions to a monetary policy shock.** Dashed-circled red lines: DSGE Bayesian mean impulse responses referring to the 1984Q1-2008Q2 period. Other lines: Sign Restriction-VAR impulse responses à la Fry and Pagan (2011) computed over 1,000 retained draws and conditional on different calibrations of the standard deviation of the monetary policy shock in the Smets and Wouters (2007) model used as DGP (estimated standard deviation, twice the estimated standard deviation, and five times the estimated standard deviation). VARs estimated with a constant and three lags.