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Delayed Overshooting Puzzle in Structural Vector Autoregression Models*

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ABSTRACT: Some authors argue that the delayed overshooting puzzle often found in the literature is an artifact of improper identification of monetary policy shocks, like Cholesky ordering. We investigate this claim by estimating the dynamic effect of monetary policy shocks on exchange rate using various identification schemes, where the data is generated by a small open economy DSGE model. We find that, on large sample, Cholesky type of restrictions perform comparably with model-consistent sign restrictions approach and do not produce the puzzle artificially. On short samples, however, Cholesky restrictions produce a more uncertain estimate for the shape of the exchange rate than sign restrictions.

KEYWORDS: Monetary Policy; Exchange Rate; DSGE; Vector Autoregressions; Cholesky Decomposition; Sign restrictions.

JEL CLASS.: E52, F41, C32

RÉSUMÉ: Certains auteurs estiment que l'énigme de sur-ajustement retardé souvent trouvée dans la littérature serait un artefact dû à une identification incorrecte des chocs de politique monétaire, comme, par exemple, la factorisation de Cholesky. Nous évaluons cette hypothèse en estimant l'effet dynamique de chocs de politique monétaire sur le taux de change en utilisant différentes stratégies d'identification, et sur un échantillon de données générées par un modèle DSGE d'une petite économie ouverte. Sur grand échantillon, nous trouvons que les restrictions à la Cholesky permettent une estimation similaire à celle obtenue par restriction de signe cohérente avec le modèle, et ne produisent pas ce paradoxe de façon artificielle. Sur de plus petits échantillons, en revanche, les restrictions de Cholesky produisent une estimation plus incertaine de la réponse du taux de change que la méthode des restrictions de signe.

MOTS-CLÉS: Politique monétaire; taux de change; DSGE; VAR structurel; factorisation de Cholesky; restrictions de signe.

CLASSIFICATION JEL: E52, F41, C32

Non-Technical Summary

In a speech in 2006, "Monetary Policy Today: Sixteen Questions and about Twelve Answers", Alan Blinder writes: "In the usual story of the role of exchange rates in monetary transmission, a country that raises its interest rates experiences a currency appreciation. The theory of uncovered interest parity explains that this happens in order to induce (rational) expectations that the currency will subsequently depreciate back to its original (real) exchange rate. Thus a tightening of monetary policy is supposed to lead to a quick appreciation followed by depreciation. Nice and logical. But, empirically, it does not happen. How, then, does monetary policy influence exchange rates? A good question. And until it gets a good answer, central bankers are operating in a dense fog. So this issue also ranks high on the research agenda..." One of the empirical results that Blinder refers to is coined as the delayed overshooting puzzle, the hump-shaped impulse response of exchange rate to a monetary policy shock found in VARs. This has been a robust finding in the literature of the exchange rate for different countries, for many years.

In this paper we work under a controlled experiment with a known data generating process and ask if this puzzle is an artifact of the estimation methodology. This investigation is motivated by the recent consensus in the literature that most of the puzzles or the non-sense results are due to improper identification strategies used in structural VAR (SVARs) analyses. This is often the case with recursive SVARs. Studies that employ this identification strategy usually assume that the exchange rate does not react to a monetary policy shock or that monetary policy does not take into account exchange rate surprises, depending on the ordering (the latter is more common). When an identification strategy that allows for the simultaneous reaction between monetary policy and exchange rate is used, the results are either inconclusive or puzzles seem to disappear.

We estimate VAR models using two sets of artificial observations from a small open economy DSGE model estimated on Swedish data and identify monetary policy shocks with two classical identification strategies used in the literature, Cholesky ordering and sign restrictions. With respect to our data generating process, the first identification imposes improper restrictions while the second imposes model-consistent sign restrictions. Our results show that when data are abundant, Cholesky-type restrictions perform comparably with sign restrictions. This is true irrespective of whether the data generating process exhibits delayed overshooting or not. Thus, we do not find evidence that zero restrictions imposed by the Cholesky scheme would lead to artificial exchange rate puzzle on long samples. On the other hand, its performance deteriorates when the number of observations is reduced to typical levels available in applied research. Given sampling uncertainty, the shape of the exchange rate response is estimated with higher uncertainty when using Cholesky ordering. Identification with sign restrictions is more robust to the sampling uncertainty despite the fact that imposing sign restrictions

typically brings additional (identification) uncertainty.

We also confirm a previous finding in the literature that SVARs performance is poor when trying to identify effects of shocks with low variance, as it is the case for monetary policy shocks. Supported by this weakness, a SVAR-identified monetary policy shock is often then a linear combination of many structural shocks that fulfill the restrictions imposed by a particular identification strategy. Adding this to sampling and estimation uncertainty increases the likelihood of SVARs to produce puzzling results.

1 Introduction

It is a common finding in exchange rate econometrics that the implications of theoretical exchange rate models are not supported by the data. One of the most famous empirical puzzles is the delayed overshooting puzzle, the hump-shaped impulse response of exchange rate to a monetary policy shock found in VARs, first introduced in Eichenbaum and Evans (1995) and supported, among others by Peersman and Smets (2003) and Scholl and Uhlig (2008). However, recent work in the literature suggests that most of the puzzles or the non-sense results are due to improper identification strategies used in VAR analyses. This is often the case with recursive SVARs. Studies that employ this identification strategy usually assume that the exchange rate does not react to a monetary policy shock or that monetary policy does not take in account exchange rate surprises, depending on the ordering (the latter is more common). When a more careful and plausible identification strategy is used, the results are either inconclusive or puzzles seem to disappear (see, among others, Kim and Roubini (2000), Faust and Rogers (2003), Bjornland (2009) and Vonnak (2010)).

These results emerge from identified-VAR analyses which try to recover a monetary policy shock from a set of observed variables for which the data generating process and the underlying shocks are not known. Practice shows that identification strategies are then considered proper if they deliver what are perceived as reasonable results. In this paper we investigate if the delayed overshooting puzzle (DOP hereafter) is an artifact of the identification assumptions in structural VAR models. We do this under a controlled experiment, where the data generating process (DGP) in the economy, the underlying structural shocks and their effects on the variables are known to the researcher. This experiment involves two versions of a small open economy DSGE model: one that does not produce delayed overshooting after a surprise monetary policy shock and another one that produces such an effect on the exchange rate. The experiment is designed as follows: first, artificial data are simulated from the DSGE model, which are then fed into SVARs with various identification schemes. The estimated responses to the monetary policy shock are then compared with the true responses. We conduct this experiment on long sample in order to get an idea about the asymptotic behavior of the SVAR estimators, but also on short sample which is a more realistic situation in applied research.

The ability of the SVAR model in recovering the true underlying monetary policy shock in the simulated data is tested for the following identification strategies: identification using Cholesky decomposition and identification with DSGE model - consistent sign restrictions only. With respect to our data generating process, the Cholesky decomposition imposes improper restrictions as they exclude contemporaneous reaction of certain variables to specific shocks contrary to the structure of the DGP¹. On the other hand, sign-identified VARs allow for simultaneous reactions, but restrict the sign

¹The inconsistency between DSGE models and the assumptions implied by the Cholesky identification are already discussed in the literature, see for example, Canova (2005) and Carlstrom and Fuerst and Paustian (2009).

of the reaction, in line with the DGP. Through this experiment, we are able to see if identification with Cholesky is insufficient to isolate the monetary policy shock and if improper restrictions it imposes lead to artificial puzzles, especially for the exchange rate. Furthermore, we can test how results from structural VARs change if we move from improper to model-consistent restrictions.

We generate the artificial data using a medium scale DSGE model for a small economy, estimated for Sweden, by Adolfson et al. (2008). This is not an incidental choice. First of all, the model as in Adolfson et al. (2008) represents a new Keynesian small open economy model in the tradition of Christiano et al. (2005) and Smets and Wouters (2003), widely used in central banks for policy analysis and forecasting. This model is in the same spirit as the pool of DSGE models developed for small open economies and has well-documented empirical properties. Furthermore, this model provides an example of how DSGE models are built to replicate some of the puzzles found in empirical work. More specifically, Adolfson et al. (2008) present a modified uncovered interest rate parity (UIP) condition that allows for a negative correlation between the risk premium and the expected change in the exchange rate. Under this modification the DSGE model delivers the hump-shaped impulse response of exchange rate to a monetary policy shock found in VARs. Adolfson et al. (2008) offer two versions of their model, one estimated with the standard UIP condition and one with the modified UIP condition. For our purpose, we use both versions of the model.

It is important to stress that in this paper our primary focus is on identifying the monetary policy shock and particularly its effect on the exchange rate, using the tools of Eichenbaum and Evans (1995), Kim and Roubini (2000), Faust and Rogers (2003) or Bjornland (2009). That is, under our experiment we do not give to the researcher more information than was available to the ones in the above mentioned studies. Furthermore, under this experiment we do not build the VAR representation of the DSGE model². Instead, we work with only a subset of the variables from the DSGE model, which are then fed to VARs, as it is usually done in empirical work when the DGP is not known. On a next step, when discussing the results, we use information from the DGPs in order to shed light on what could help or mislead the SVARs to correctly identify the true effects of the structural shocks.

For the SVAR estimation we use the most common variables employed in such analysis for a small open economy: production, consumer prices and interest rate for both foreign and domestic economy and the nominal exchange rate. Foreign economy variables are treated as exogenous. Results show that identified VARs are able to discriminate between different data generating processes and that the identification schemes with DSGE model-consistent restrictions are more successful in recovering the effects of the underlying monetary policy shock. Interestingly, when data are abundant, these results show that Cholesky-type restrictions, although not always successful in recovering the true monetary

²See Fernandez-Villaverde et al. (2007). We are aware of the issues raised when trying to approximate the data generating process behind a DSGE model with a finite lag VAR with omitted variables, therefore we do not expect a perfect match of the responses from the estimated SVARs and the DSGE model.

shock, do not produce delayed overshooting in case the data generating process does not contain it. In case when the delayed overshooting is a feature of the DGP, all the identification schemes under investigation recover it correctly. This is in contrast to what the recent literature seemed to converge to, assigning the DOP to improper restrictions in the Cholesky identification. However, on small sample, all the identification strategies under discussion tend to suggest a short delayed response of the exchange rate when there is no delay in the DGP or a less delayed response than the true response.

In overall, from our experiment we learn that SVARs do not suggest wrongly a DOP when the data is abundant but when improper identification is combined with sample uncertainty. Also we observe that the SVARs perform poorly in recovering correctly the true responses of the monetary policy shock in general. We investigated several potential reasons for this result. In line with Paustian (2007) and Canova and Paustian (2011), we find that SVARs performance is poor when identifying shocks with relatively weak signal. When we counterfactually increase the size of the monetary policy shock we observe a big improvement in the ability of the SVARs to recover the true responses. In addition, the SVAR-identified monetary policy shock is a linear combination of structural shocks, under both identification schemes. We find that the SVAR-monetary policy shock contains most prominently loadings of the true monetary policy shock followed by a domestic markup shock and a risk premium shock. The domestic markup shock especially appears to “trick” the Cholesky to produce a price puzzle and to “help” it overcome the delayed overshooting puzzle. This issue is more notable when the signal of the monetary policy shock is relatively weak to other shocks.

In addition, we test the same identification strategies with real data from the Swedish economy. Estimates on actual data are closer to the small sample simulation results with no delayed overshooting from Adolfson et al. (2008). As it was the case with the simulated data, the result regarding the response of the exchange rate is robust to both identification schemes used, casting some doubt on whether it is justified to apply a modified UIP in the Adolfson et al. (2008).

Controlled experiments in similar fashion have been recently used in the literature mostly to assess the effects of the monetary policy on output in the case of closed economies³. For the open economy case, Jaaskela and Jennings (2011) have used a DSGE model estimated for Australia to assess the ability of VARs in identifying monetary policy shocks, finding that recursive identification produces puzzles while identification with sign restrictions does not, provided that a sufficient number of shocks are identified. Relative to their work, we focus on the exchange rate and contribute to the literature by performing the experiment with data generated from an empirically plausible DSGE model with a richer structure. With this model we have the advantage to test the performance of VAR identification strategies for both, the case when the DGP features delayed overshooting of the exchange rate and the case when it does not. Furthermore, our approach compares favorably as it provides insights into how

³See, among others, Carlstrom et al. (2009), Liu and Theodoridis (2012) and Castelnuovo (2013b).

identification strategies perform given asymptotic and small sample properties of the SVAR estimates. In contrast to Jaaskela and Jennings (2011) we use minimal restrictions to identify the effect of the monetary policy shock on the exchange rate, a situation that is more plausible to occur in practice, when the researcher may not have a priori valid assumptions to identify many of the shocks present in the data.

The paper proceeds as follows: Section 1 describes shortly the Adolfson et al. (2008) model that we use as a data generating process. Section 2 presents the steps of the controlled experiment, the results from the VAR analysis under different identification schemes, in long and small samples and a discussion on them. Section 3 concludes.

2 The model as DGP

To generate the data we use a medium scale DSGE model as in Adolfson et al. (2008). This model represents a new Keynesian small open economy model in the tradition of Christiano et al. (2005) and Smets and Wouters (2003), widely used in central banks for policy analysis and forecasting. This model is in full operational use at the Sveriges Bank. The model features a number of nominal and real frictions such as sticky prices and sticky wages, incomplete exchange rate pass-through, habit persistence and investment adjustment cost. Adolfson et al. (2008) is a version of the model developed in Adolfson et al. (2007) with the main difference being in the specification of the uncovered interest rate parity (UIP) condition. The first version of the model employs a standard UIP condition while the latest version of the model employs a modified UIP condition. The modified version allows for a negative correlation between the risk premium and the expected change in the exchange rate. Under this modification, the DSGE model delivers a hump-shaped response of the exchange rate to a monetary policy shock, thus replicating the delayed overshooting of the exchange rate commonly found in empirical works.

The domestic economy in Adolfson et al. (2008)⁴ is comprised of households, four types of firms working on the domestic and the external sector, a government and a monetary authority. Households choose consumption, work effort and holding of real balances to maximize their utility. They consume domestic and imported goods, own physical capital, domestic and foreign assets. They supply labor monopolistically and set their wages subject to *à la Calvo* rigidity. On the production side, there are four different firms: domestic goods firms, importing consumption, importing investment, and exporting firms. They all produce a differentiated good and have monopolistic power when setting prices. Price setting in all these sectors is subject to *à la Calvo* rigidity.

Monetary authority is assumed to follow a Taylor rule, with short term interest rate reacting to deviations of CPI inflation from the inflation target, to output gap and to the real exchange rate. The

⁴For a detailed description of the model see Appendix A and also please refer to Adolfson et al. (2008).

foreign economy is modeled exogenously, resembling the typical situation of a small open economy with no power in the world market. Both foreign economy and fiscal policy are modeled and estimated exogenously as VAR processes out of the DSGE model. The model is rich in terms of the exogenous shocks, containing 21 of them. Among these shocks we mention here preference shocks to consumption and labor efforts, technology shocks (a unit root and a stationary technology shock), four sector specific mark-up shocks, a monetary policy shock, shocks related with fiscal policy, a risk premium shock and shocks coming from the foreign economy.

Adolfson et al. (2008) estimate both versions of the model on Swedish data using Bayesian techniques for the period 1980Q1-2004Q4. When exploring the consequences of the modification for empirical coherence their results show a preference for the model featuring the modified UIP condition. The modification seems to be important especially for the model's forecasting performance.

2.1 Standard and modified UIP condition

Let's denote the model with standard UIP condition as M1 and the model with modified UIP condition as M2. In Adolfson et al. (2008), households hold cash, domestic bank deposits and foreign bonds. Saving in domestic deposits pays a gross nominal rate of R_t while saving in foreign bonds pays a gross interest rate of R_t^* adjusted for a risk premium of holding the foreign bonds⁵.

In M1, the risk premium is a function of net foreign asset position of the domestic households, $a_t \equiv (S_t B_t)/P_t z_t$. In M2, risk premium specification contains an additional term, namely the expected change in the exchange rate, $\frac{E_t S_{t+1}}{S_{t-1}}$, with S_t and B_t being the nominal exchange rate and bond holding, respectively. Adolfson et al. (2008) base and justify this modification on the empirical findings of strong negative correlations between risk premium and expected change in the exchange rates (the so-called the forward premium puzzle). In M2 risk premium is given by:

$$\Phi(a_t, S_t, \tilde{\phi}_t) = \exp(-\tilde{\phi}_a(a_t - \bar{a}) - \tilde{\phi}_s(\frac{E_t S_{t+1}}{S_t} \frac{S_t}{S_{t-1}} - 1) + \tilde{\phi}_t) \quad (1)$$

where $\tilde{\phi}_t$ is the risk premium shock. In M1, risk premium has a similar specification with $\tilde{\phi}_s = 0$.

Given the specification of the risk premium in M2, combining the first order condition for cash and for foreign bond holdings from the maximization problem of the households, we derive the following modified UIP condition (in log-linearized form):

$$\hat{R}_t - \hat{R}_t^* = (1 - \tilde{\phi}_s)E_t \Delta \hat{S}_{t+1} - \tilde{\phi}_s \Delta \hat{S}_t - \tilde{\phi}_a \hat{a}_t + \tilde{\phi}_t \quad (2)$$

while the standard UIP condition in M1 is of the form:

⁵The introduction of risk premium is needed to ensure a well-defined steady state for open economy models (see Schmitt-Groh and Uribe (2003)) for more details.

$$\hat{R}_t - \hat{R}_t^* = E_t \Delta \hat{S}_{t+1} - \tilde{\phi}_a \hat{a}_t + \tilde{\phi}_t \quad (3)$$

Figure 1 illustrates the effect of an exogenous monetary policy shock to output, consumer prices, nominal interest rate and nominal exchange rate, from M1 and M2. In both cases, the contractionary shock induces a gradual decline of output and prices and an appreciation of the exchange rate on impact. Under the model with standard UIP condition, the sharp appreciation of the exchange rate is followed by a gradual depreciation. However, under the modified UIP, the exchange rate reaches its peak appreciation only after several quarters, resembling the delayed overshooting found in empirical works. This change is due to the larger persistence that the modified UIP condition assumes. The negative correlation between risk premium and the expected change of the exchange rate, induces the households to require a lower risk-adjusted R_t^* on the foreign bond holdings. But since the returns on domestic and foreign assets should equalize, a larger appreciation of exchange rate is required.

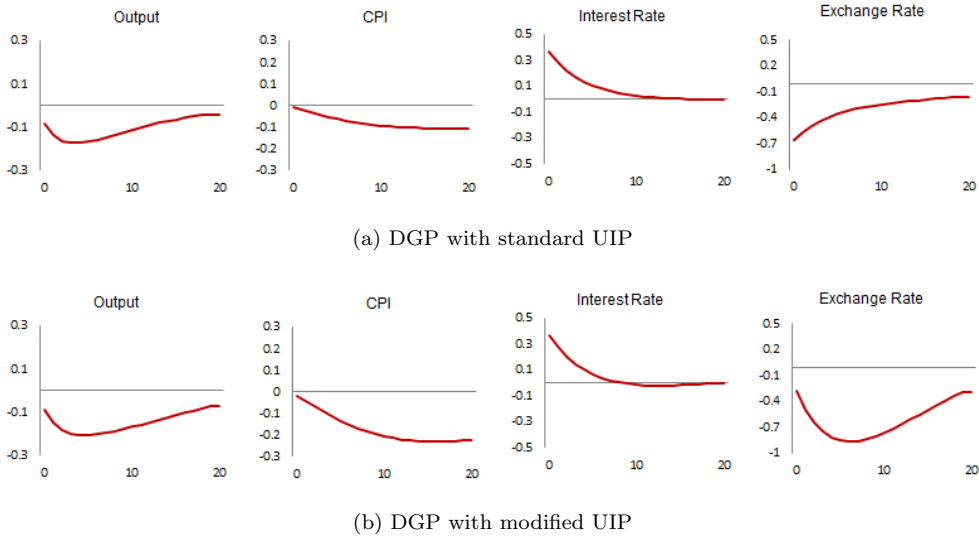


Figure 1: IRFs to a monetary policy shock

Notes: The line in red denotes the median impulse response from the estimated DSGE model as in Adolfson et al. (2008). A decrease in exchange rate implies appreciation. Horizontal axis is lag horizon in quarters. In vertical axis, the log deviations from the steady state.

3 VAR analysis

To examine the effect of monetary policy in real life problems using identified-VAR modeling, the econometrician needs to recover the monetary policy shock from the variables she can observe. In this case the econometrician does not know the true data generating process but uses, in the best case, a priori justifiable identifying restrictions to be able to recover such shocks. Practice has shown

that the identification strategy in VARs is considered to be proper when they are able to deliver what are perceived as reasonable results (see for example Uhlig (2005) for a discussion). Instead, when using simulated data from a DSGE model, we know the data generating process in the economy, the underlying structural shocks and their effect on the variables. As the econometricians with actual data do, we can select a list of variables from this economy, estimate VAR models on them and investigate if the VAR identifying restrictions used in the literature are able to recover the true underlying structural shocks.

In the following we run exactly this experiment⁶. First we solve the estimated DSGE model using calibrated parameters and the posterior median of estimated parameters as in Adolfson et al. (2008). Then we simulate the model and with a selected subset from the simulated variables we estimate structural VARs and investigate their performance in recovering the monetary policy shock and its effect on the exchange rate. Their performance is judged by comparing VAR responses with the true responses from the DSGE model. The set of the DSGE model variables chosen for the VAR estimation comprises the following variables: foreign output, y^* , foreign consumer prices, p^* , foreign interest rate, i^* , domestic output, y , domestic consumer prices, p , domestic nominal interest rate, i , and the nominal exchange rate, s . This set represents the most common set of variables used in the empirical literature of monetary policy analysis for small open economies. We estimate the VARs using variables in levels and adopt a Bayesian approach when making statistical inference. The VAR coefficients are drawn from a normal-inverse-Wishart distribution with flat prior, following Uhlig (2005).⁷ The frequency of the data is quarterly.

In reality the econometrician does not have the luxury to conduct VAR analysis using long series of data for different reasons. In cases when long samples of data exist, most probably there are regime changes or structural breaks and estimating a VAR on these data would be in contradiction with the principle that VARs should be estimated on sample periods that define reasonably constant parameter regimes. With simulated data from a DSGE model we do not face this problem so we are able to approximate arbitrarily closely the asymptotic distribution of our estimators. For the long sample VAR analysis we use a simulated sample of 10,000 periods. Since VAR analyses are typically made on much shorter samples, we investigate the small sample properties of our estimators as well. For the small sample SVAR analysis, we use 500 sets of simulated data, each having 150 observations.

⁶Examples of such experiments are found in the literature, see among others Canova and Pina (2005) and Carlstrom et al. (2009).

⁷We have a constant in the model. Stock et al. (1990) show that in the presence of non-stationarity OLS estimates are super-consistent. Sims (1998) shows that Bayesian inference with a standard Normal-Wishart prior is not affected by non-stationarity.

3.1 Identification of monetary policy shock

The reduced-form VAR model has the following representation:

$$A(L)z_t = \epsilon_t \tag{4}$$

where z_t is a n -dimensional vector of variables, $A(L) \equiv I + A_1L - A_2L^2 - \dots - A_pL^p$ is the autoregressive lag order polynomial and ϵ_t represents the reduced-form errors with covariance matrix, \sum_{ϵ} . $A(L)$, ϵ_t and \sum_{ϵ} can be estimated consistently with standard estimation methods. However we are interested in recovering the corresponding structural shocks, u_t , which relate to the reduced-form errors through the following relationship: $u_t = S\epsilon_t$, with $E(u_t u_t') \equiv \sum_u = I_n$. To this purpose we need to recover the elements of S . Noting that $\epsilon_t = S^{-1}u_t$, we can see that $\sum_{\epsilon} = S^{-1}E(u_t u_t')S^{-1'} = S^{-1}S^{-1'}$. Since \sum_{ϵ} is known (estimated), we can uniquely identify on maximum $n(n+1)/2$ from n^2 elements of S^{-1} . For exact identification we should impose additional $n(n-1)/2$ restrictions on S^{-1} .

The literature of structural vector autoregressions has continuously evolved in finding valid identifying restrictions which vary from short and long run zero restrictions to sign restrictions⁸. In the following, we revisit some of the identifying assumptions used in the literature and investigate their ability in recovering the true underlying monetary policy shock in the simulated data, and especially its effect on the exchange rate. More specifically, we will investigate the following identification strategies⁹: identification using the Cholesky decomposition and identification with sign restrictions only. With respect to our data generating processes, Cholesky decomposition imposes improper restrictions as they exclude contemporaneous reaction of certain variables to specific shocks contrary to the structure of the DGP. On the other hand, sign-identified VARs allow for simultaneous reactions but restrict the sign of the reaction. Through this experiment we are able to see if Cholesky is insufficient to isolate the monetary policy shock and if improper restrictions it imposes lead to unexpected results, especially for the exchange rate. Furthermore, we can test (at least with respect to our DGPs) how results from structural VARs change if we move from improper to model-consistent restrictions.

Let $z_t = (y_t, p_t, i_t, s_t)$ be the block of the endogenous variables for the small open economy and $z_t^* = (y_t^*, p_t^*, i_t^*)$ the block of the exogenous variables. In line with the DGP, a block exogeneity is imposed on the foreign variables, so the domestic variables cannot affect them. With Cholesky identification we impose a recursive ordering of the shocks and assume that a monetary policy and an exchange rate shock will not affect contemporaneously the GDP and CPI. Furthermore, an exchange

⁸For a recent survey on structural VARs and identification assumptions refer to Kilian (2011).

⁹Another identification strategy that has proved successful in avoiding the delayed overshooting of the exchange rate is the one proposed in Bjornland (2009), where a combination of zero short- and long - run restrictions is used for the identification of the monetary policy shock. This identification uses the assumption that the long run effect of monetary policy in the real exchange rate is neutral. Technically, to use this identification strategy requires the real exchange rate to be non-stationary in levels. In the DSGE model we use, the PPP holds and the real exchange rate is stationary in levels. That is, under our experiment we cannot test the identification strategy as in Bjornland (2009).

rate shock is assumed to not affect interest rates as well (see for example Eichenbaum and Evans (1995)). Being ordered last, an exchange rate shock will have a delayed effect on output, prices and interest rates. Such a slow exchange rate pass-through to macroeconomy can be justified by the presence of nominal rigidities in the economy. These (identifying) assumptions provide enough restrictions for the Cholesky decomposition to be applied. This case belongs to the class of fully identified VAR models where all shocks are identified even though we are interested only in the identification of the monetary policy shock. Under the Cholesky identification the relationship between the reduced-form errors and structural shocks is presented as below¹⁰:

$$\begin{pmatrix} \epsilon_t^y \\ \epsilon_t^p \\ \epsilon_t^i \\ \epsilon_t^s \end{pmatrix} = \begin{bmatrix} x & 0 & 0 & 0 \\ x & x & 0 & 0 \\ x & x & x & 0 \\ x & x & x & x \end{bmatrix} \times \begin{pmatrix} u_t^1 \\ u_t^2 \\ u_t^{MP} \\ u_t^4 \end{pmatrix}$$

The pure sign restriction identification relaxes the contemporaneous rigidity on GDP and CPI to the monetary policy shock.¹¹ It also allows for contemporaneous responses between monetary policy and the exchange rate. This is based on the idea that exchange rate developments are important for monetary policy decisions in small open economies (see McCallum (1994), Cushman and Zha (1997)). The direction of the restrictions are justified by the general postulation that a monetary contraction is associated with an increase in interest rates, a drop on output and prices and an exchange rate appreciation on impact. Using the sign restrictions, the relationship between the reduced form errors and structural shocks is presented as below:

$$\begin{pmatrix} \epsilon_t^y \\ \epsilon_t^p \\ \epsilon_t^i \\ \epsilon_t^s \end{pmatrix} = \begin{bmatrix} x & x & - & x \\ x & x & - & x \\ x & x & + & x \\ x & x & - & x \end{bmatrix} \times \begin{pmatrix} u_t^1 \\ u_t^2 \\ u_t^{MP} \\ u_t^4 \end{pmatrix}$$

By restricting the sign of the exchange rate response on the impact period only, by construction, we avoid exchange rate puzzles which are immediate depreciations of the exchange rate after a monetary policy shock. Still, after this period the exchange rate can react freely in whatever direction, keep appreciating or depreciate.¹² The sign restrictions we set above are also model consistent and unique to the monetary policy shock in the DGP. If we leave one of the variables above not sign restricted

¹⁰This discussion does not relate to our specific DGP but it is a general one on the mapping between reduced-form errors and structural errors when the DGP is not known, as it is usually the case in practice.

¹¹With real data, delayed CPI and GDP reaction to monetary policy may not be a bad assumption for monthly data.

¹²The restriction on impact might slightly influence the shape of the response as well but not importantly, since for the period after the shock there is no restriction.

than we would face the multiple shocks problem as discussed in Fry and Pagan (2011). In Adolfson et al (2008) there are four other shocks that have the same sign impact as the monetary policy shock on interest rate and exchange rate. These shocks are the markup shocks for imported consumption and investment, the asymmetric technology shock and the consumption preference shock. However, in comparison with these shocks, the monetary policy shock differs on the sign impact on output and consumer prices. Therefore, by imposing a sign restriction on the four variables in the VAR one should, *a priori*, discriminate the monetary policy shock between the other shocks discussed above.¹³

In the following, we show the evidence from the estimated VARs for each of the identification scheme in parallel. First, we summarize and interpret the evidence from structural VARs mostly based on the posterior medians of impulse responses and the corresponding 16th and 84th percentiles. Sims and Zha (1999) argued that using medians and confidence bands by simply connecting the percentile values for each horizon can be misleading about the true nature of the uncertainty. This is particularly true when drawing conclusions on the shape of the impulse response. A similar point was made by Fry and Pagan (2011) and Inoue and Kilian (2011). To account for this discussion, for robustness, we also show the “median target” response as suggested by Fry and Pagan (2011), for SVARs with sign restriction and in long sample. Furthermore, since we are primarily interested in the dynamics of the exchange rate response, we construct an additional statistics to capture the most important characteristic of it, namely the peak location of the exchange rate appreciation. Peak location refers to the distribution of the quarter with the highest appreciation of the exchange rate (distribution of the peak location).

3.2 Results and Discussion

In the following, we show the evidence from the estimated VARs first for the long sample and then for the Monte Carlo exercise with small samples, for each identification scheme and DGP, in parallel. All SVARs are estimated with an optimal lag length selected with Bayesian Information Criteria (BIC). The information on the selected lag is given below each figure. The baseline results are then followed by a detailed discussion which takes into account several features of the DGPs in hand, with the aim to understand better the SVAR performance.

3.2.1 Long sample

Figure 2 reports the responses of output, consumer prices, interest rates and the nominal exchange rate to a SVAR identified monetary policy shock. The black line depicts the median response, the magenta line depicts the Fry-Pagan response and the true responses from the DSGE model are shown in red. The size of the monetary policy shock is normalized such that the impact response of the nominal

¹³In Adolfson et al (2008), the true response of consumer prices on impact to the monetary policy shock is negative and small but not equal to zero.

interest rate is the same as the true response from the respective DGP, to ensure that the differences between the true and the estimated responses are not due to biases in estimating the size of the shock.

In general, we do not observe a poor match between the responses from Cholesky identification and the true ones regardless of the data generating process. With data from model M1 the response of prices displays the familiar “price puzzle” and an overestimated and more persistent reaction of the interest rate. Although with differences in the magnitudes, the response of output is recovered correctly. With data from model M1, the response is slightly overestimated and with data from model M2 is muted in comparison with the respective true responses. Regarding the exchange rate, even though its response appears less persistent than the true one, its direction is correctly recovered for both data generating processes. More specifically, in panel (a) of Figure 2 we observe an appreciation of the exchange rate on impact, followed by a continued depreciation thereafter, in line with what the underlying model with standard UIP condition predicts (although a much faster depreciation). In panel (c) of Figure 2, the appreciation of the exchange rate continues for more quarters following the shock, in line with the delayed overshooting the underlying model with modified UIP condition predicts.

On the other hand, compared to the recursive identification, the usage of sign restrictions introduces identification uncertainty in our SVAR analysis, because the econometrician identifies a set of impulse responses satisfying the restrictions. This is visible from the wide credible intervals around the median response (see panel (b) and (d) of Figure 2).¹⁴ We also observe another common finding in the literature, namely, the consistent overestimation of the impact of the identified shock with sign restrictions. The “price puzzle” is avoided only for some quarters. With respect to the exchange rate, although the response is overestimated in comparison with the true response, the appreciation shows up immediate in panel (b) and delayed in panel (d), in line with what the underlying DGPs suggest.

In panel (b) and (d) of Figure 2 we observe as well that the Fry-Pagan response is not a superior measure to the simple median response (magenta line is most of the time on top of the black line).¹⁵ Still, in order to provide a more accurate assessment for the shape of the response of the exchange rate we investigate this response through quarters by calculating the posterior distribution of the quarter where exchange rate response is the lowest (the timing of the peak appreciation). Figure 3 shows the posterior distribution of peak locations, per each identification strategy and each DGP.

Results show that in case the DGP is the model without delayed overshooting (M1), all peak appreciations suggested by the SVAR with Cholesky are recovered correctly, in the first quarter. With sign restrictions, although the uncertainty of the responses has increased in comparison with Cholesky, results regarding the posterior distribution of the peak appreciation remain robust; about 80 percent of peak appreciations occur immediately in the first quarter. The peak appreciation corresponding

¹⁴Note that the scale of the y-axis in the graphs with sign restrictions is larger, not comparable with those for Cholesky. Therefore, the wider credible bands with sign restrictions might not be immediately obvious.

¹⁵Similar result is found in the literature as well, see Canova and Paustian (2011) for an example.

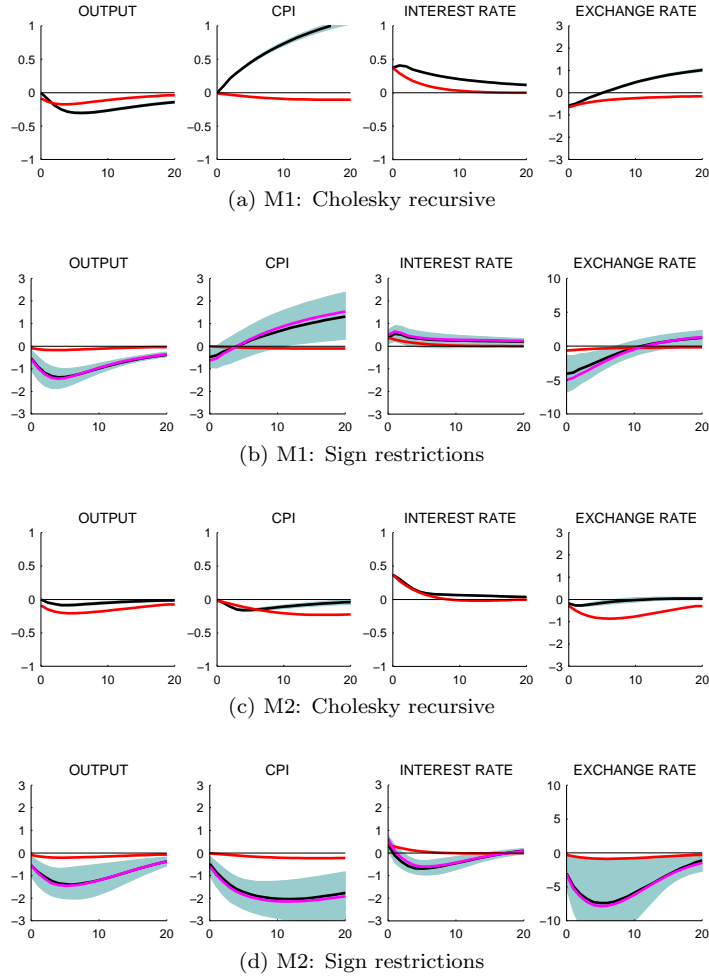


Figure 2: Long Sample: IRFs to a monetary policy shock identified with SVARs

Notes: The solid line in black denotes median impulse response from the estimated VAR(4), large sample size ($T=10,000$), and the shaded area the corresponding 68 percent credible interval. The line in magenta denotes the Fry-Pagan median-target response. The line in red denotes the median impulse response from the estimated DSGE model as in Adolfson et al. (2008). A decrease in exchange rate implies appreciation. All sign restrictions are set only for one period (the shock impact period). Horizontal axis is lag horizon in quarters.

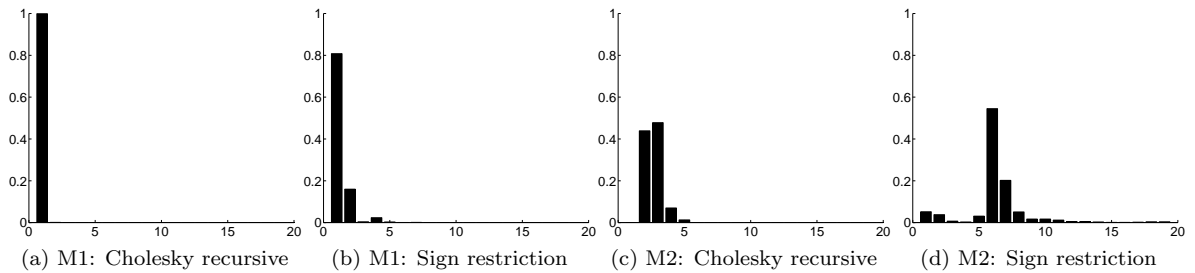


Figure 3: Long Sample: Posterior distribution of peak appreciation of exchange rate to a monetary policy shock

Notes: Bars in black show the results from the estimated VAR(4), large sample size ($T=10,000$). The true peak appreciation is in the first quarter for M1 and in the sixth quarter for M2.

to the Fry-Pagan response also occurs in the first quarter. When the DGP is the model with delayed overshooting, Cholesky identification predicts that around 80 percent of the peak appreciations occur in the second and third quarter, identifying correctly a delayed appreciation of the exchange rate although a less persistent one when compared with the true response where peak appreciations occur in the sixth and the seventh quarter. With sign restrictions, the distribution of the peak is centered on the sixth quarter (the peak of the Fry-Pagan response also occurs in the sixth quarter).

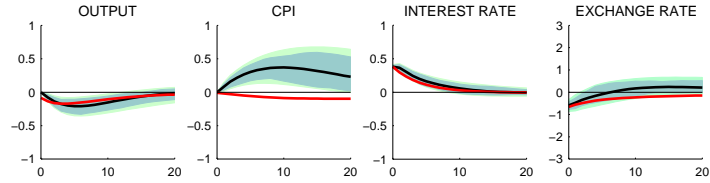
To sum up, from the analysis with long samples (with low estimation uncertainty) we observe that even though Cholesky recursive assumes improper restrictions with respect to both DGPs we use, it captures correctly the exchange rate response to a Cholesky-identified “monetary policy shock”. The delayed overshooting of the exchange rate shows up correctly when SVARs are estimated with data from the DSGE model with delayed overshooting. In the case of the theoretical model without this feature, the Cholesky identification does not produce the puzzle artificially. This is an interesting result taking into account that the recent literature¹⁶ has been converging to the conclusion that the DOP may be an artifact of improper restrictions one imposes with Cholesky identification.

3.2.2 Monte Carlo Exercise

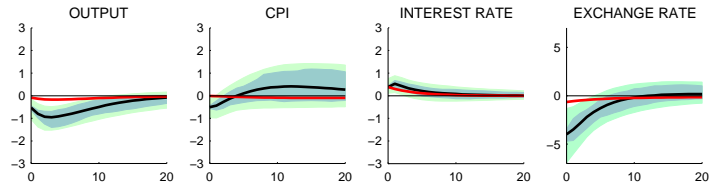
In order to investigate the small sample properties of our SVAR estimations, we repeat the experiment using 500 samples of simulated data with 150 observations. For each sample of data we estimate the same SVARs as discussed previously, with the aim to identify the underlying monetary policy shock and its effect on the exchange rate. Figure 4 presents the responses for data simulated from model M1 and M2 in parallel. In this figure, the line in black denotes the median response across all samples and the light blue area its 68 percent credible interval. In addition, we show the posterior distribution of the sample medians, in darker blue, with the aim to see how the estimation uncertainty evolves within samples.

As one might expect, given sampling uncertainty, the light blue area shows higher uncertainty about the response of the variables to the identified monetary policy shock, for both identification strategies. On large sample, the median responses were precisely estimated under Cholesky while on small sample there is more uncertainty around its estimates. On the other hand, the credible set of all-sample median with sign restrictions continues to be relatively wide as before. On small sample, responses with sign restrictions “suffer” from both sample and identification uncertainty but as Canova and Paustian (2011) as well show, under this type of identification, sample uncertainty is small to identification uncertainty. The distribution of sample medians are almost as wide as the all-sample distribution which means that bulk of the estimation uncertainty stems from the choice (or the realization) of the particular sample and not from the uncertainty within one sample. Put differently, when we estimate the impulse

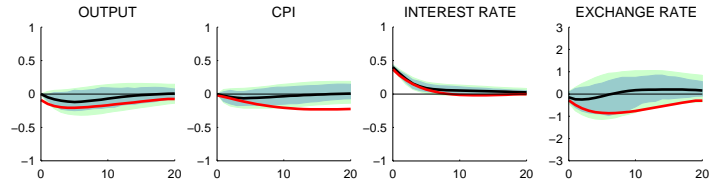
¹⁶See Cushman and Zha (1997), Kim and Roubini (2000), Bjornland (2009) and Bjornland and Halvorsen (2010).



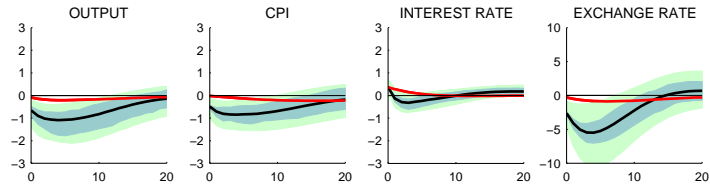
(a) M1: Cholesky



(b) M1: Sign restrictions



(c) M2: Cholesky



(d) M2: Sign restrictions

Figure 4: MC: IRFs to a monetary policy shock identified with SVARs

Notes: The solid line in black denotes the point wise median of impulse responses from all samples. The shaded area in light blue indicates the 16th and 84th percentile region. The deep blue area corresponds to the 68 percent distribution of the point wise sample medians. The VAR for each sample is estimated with $T=150$ and lag selected with BIC and the chosen lag is 3. The line in red denotes the true impulse response from the DSGE model. A decrease in exchange rate implies appreciation. All sign restrictions are set only for one period (the shock impact period). Horizontal axis is lag horizon in quarters.

responses on a small sample, the estimates will be unreliable even if we have narrow confidence bands.

Qualitatively, the results are comparable with those from the long sample analysis. The “price puzzle” is present for the identification scheme with zero restrictions. The true DSGE responses are either contained within the credible interval or lie close to its bounds. Estimates under Cholesky identification are particularly uncertain when data are from model M2 with delayed overshooting, as except from the interest rate, the distributions of the responses lie quite symmetrically around the horizontal axis. This uncertainty is reflected even in the distribution of the peak appreciation of the exchange rate. With data from model M1, about 80 percent of peak appreciations appear in the first and second quarter while in the other case the distribution is dispersed between the first and the tenth quarter.

On the other hand, in SVARs with sign restrictions, the small-sample results do not seem to differ significantly in magnitudes from those with long sample. The responses are overestimated compared to the true ones; however there is an improvement on the persistence of the responses. With respect to the exchange rate, respectively, about 65 percent of the sample median peak appreciations correspond to the first quarter in case of the DGP without delayed overshooting. In the other case, the distribution of the peaks falls around the fifth quarter. When compared with the true distribution and with the results from the long sample analysis we observe that in short sample the distribution moves away from the true value. In case the delayed overshooting is not an effect of the monetary policy shock, the SVARs suggest a delayed peak response (at least one quarter after the shock) and in case delayed overshooting is there, the SVARs suggest a less delayed one.

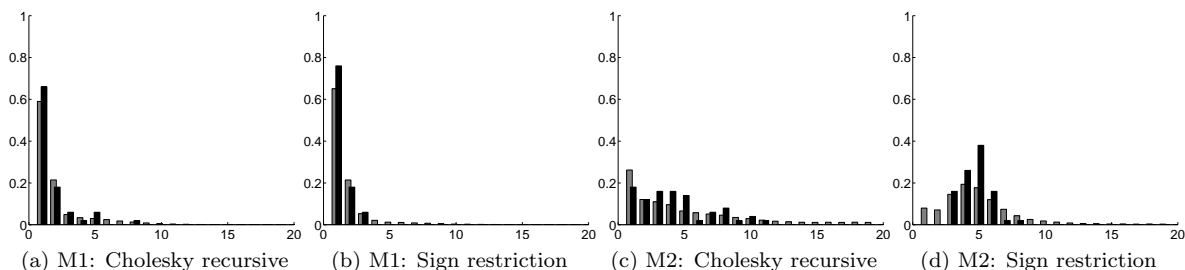


Figure 5: MC: Peak appreciation of exchange rate to a monetary policy shock

Notes: Results from the estimated VAR(3), small sample size, T=150. Bars in gray indicate across all samples peak appreciations. Bars in black correspond to sample medians peak appreciation. The true peak appreciation for M1 is in the first quarter and for M2 in the sixth quarter.

This exercise shows that a combination of improper identification (as Cholesky is with respect to our DGPs) and sampling uncertainty tend to wrongly suggest a delayed overshooting. In large sample, with minimized sample uncertainty, improper identification did not lead to a false puzzle.¹⁷ Still in

¹⁷We performed the same experiment with two other small open economy DSGE models as data generating processes, the model of Medina and Soto (2007) estimated for Chile and the model of Jaaskela and Jennings (2011) estimated for

reality, the econometrician would not be able to avoid estimation uncertainty given the available length of the series or other potential problems.

3.2.3 Discussion of results

Using the tools of Eichenbaum and Evans (1995), Kim and Roubini (2000), Faust and Rogers (2003) or Bjornland (2009) we checked whether the delayed overshooting puzzle is an artifact of the identification scheme that wrongly imposes zero restrictions. In our experiment we found that this is not the case when the data is abundant but when the improper identification is combined with sample uncertainty. However, even though our identification strategies performed relatively well in capturing the timing of the exchange rate appreciation to a monetary policy shock, they performed less well in recovering correctly the true responses of the monetary policy shock in general. There can be different reasons for this, related either to the mapping of the DSGE model into a VAR or to the disability of SVARs to correctly identify the effect of shocks with weak signal. Given that we are under a controlled experiment, we have the luxury to explore these issues in order to get more insights on the above results. More specifically, with long sample SVARs we discuss below issues related to the “truncation bias”, to the relative strength of the monetary policy shock and to the possibility of the identified shock to be a linear combination of other shocks. In addition, we explore some of these issues when combined with sample uncertainty under the MC exercise. The respective results are organized in Figure 6, showing the SVAR-responses to the identified monetary policy shock, and in Table 1 and 2, showing the share of peak appreciations of the exchange rate in the first quarter, with data from model M1.

Chari et al. (2005), Christiano, Eichenbaum, and Vigfusson (2006) and Ravenna (2007), among others, have discussed extensively the potential problems arising when mapping DSGE models into VAR models¹⁸. In the context of our analysis some of the issues mean that, apart from the sample bias and the identification bias discussed above, mapping a DSGE into a VAR model might suffer also from the so called “truncation bias”. This bias might arise when a finite ordered VAR is chosen to approximate the true dynamics implied by the DSGE model, a VARMA process. However, Ravenna (2007), among others, has shown that, for a fixed number of lags, the truncation bias is larger at longer horizons. Since in this paper we are interested on the exchange rate response on shorter horizons, the truncation bias should be a minor issue. This is indeed the case. On large sample, we estimate the SVAR models with the lag length suggested by BIC, 2, and 50, and observe that SVAR responses are comparable across these lags, per each identification strategy, for all horizons.¹⁹ As expected, the share of peak appreciations of the exchange rate to a monetary policy shock delivers the same message. With

Australia. In these models, the theoretical monetary policy shock does not produce a delayed overshooting response of the exchange rate. The SVAR results are comparable with the results as discussed above. We do not include the graphs to this version of the paper but they are available upon request.

¹⁸For a recent overview of the discussions in the literature please refer to Giacomini (2013).

¹⁹To make the paper self-contained we do not present these figures. They are available upon request.

data from model M1, Cholesky peak appreciations belong correctly to the first quarter, for all lags. A slight improvement is observed in the case of sign restrictions, where the share of peak appreciations in the first quarter increases with the number of lags, from 77 percent to 82 percent (see Table 1, 1st column).

Table 1: Long sample: Share of peak appreciations in the 1st quarter, model M1.

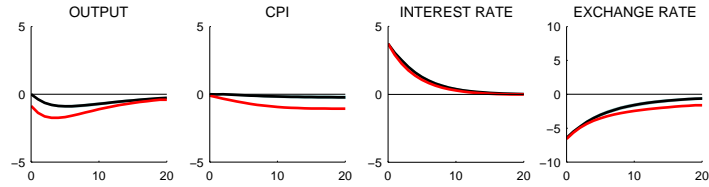
	Baseline	Large MP shock	Muted RP shock	Muted DM shock
Cholesky				
VAR(2)	100	100	100	98.9
BIC	99.9	100	100	98.2
VAR(50)	100	100	100	98.7
Sign restrictions				
VAR(2)	77.0	91.0	84.0	76.8
BIC	78.5	90.0	97.4	78.4
VAR(50)	82.0	89.1	97.0	51.0

Notes: Large MP shock (data from model M1 with monetary policy shock increased ten fold). Muted RP shock (data from model M1 with muted risk premium shock). Muted DM shock (data from model M1 with muted domestic markup shock). BIC refers to the optimal lag selected with Bayesian Information Criteria which is 4.

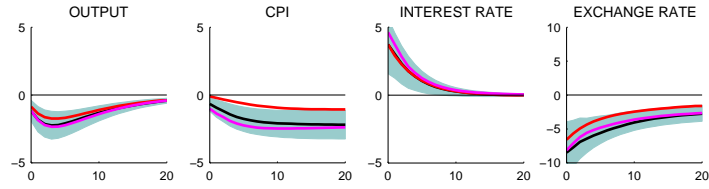
In addition, especially with respect to the sign restrictions, Paustian (2007) and Canova and Paustian (2011) argue that the size of the shock to be identified should be sufficiently large for a successful SVAR identification. We confirm this finding for both identification schemes and show that the SVARs do perform better in recovering correctly the true impulse responses of a monetary policy shock when the “signal” of this shock is relatively strong. In Adolfson (2008) the estimated strength of the monetary policy shock is relatively small and its contribution in explaining the forecast error variance (FVED) of our variables, y_t, p_t, i_t, s_t , is 4.26, 0.10, 9.04 and 3.42 percent, respectively. Increasing the strength of the monetary policy shock tenfold would make it comparable in strength with 12 of the 21 shocks in the model. Panel (a) and (b) of Figure 6 show the responses from the SVARs with data from model M1, where the DSGE monetary policy shock is counterfactually increased tenfold. We observe that, in comparison to the baseline model, the price puzzle disappears for all periods, for both identification schemes. With respect to the response of the exchange rate, there is a big improvement towards the correct persistence and also on the share of peak appreciations under sign restrictions (see 2nd column of Table 1). Also one can observe that the overestimation of the responses under sign restrictions is severely reduced.²⁰

Furthermore, another suggestion from related literature is that the SVAR-estimated monetary policy shock might be a combination of several structural shocks. With data from a simple New-Keynesian model, Carlstrom et al. (2009) have shown that the Cholesky-identified monetary policy shock is in fact a linear combination of the underlying true structural shocks such that output and prices are unaffected on impact. On the other hand, Castelnuovo (2013a) has a related argument for SVARs with

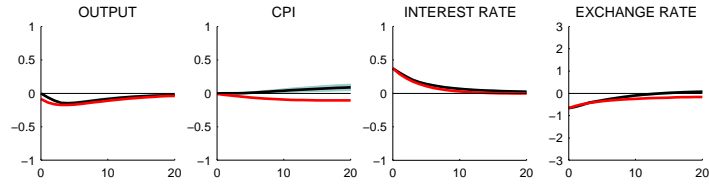
²⁰In Appendix B, Figure 2 shows the same result when SVARs are estimated on data from model M2.



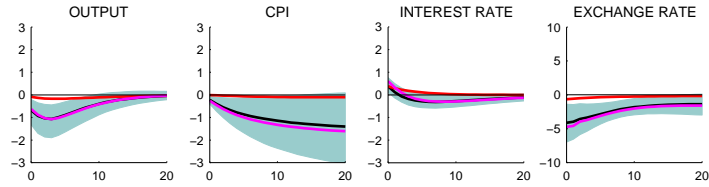
(a) Large MP shock: Cholesky



(b) Large MP shock: Sign restrictions



(c) Muted DM shock: Cholesky



(d) Muted DM shock: Sign restrictions

Figure 6: Counterfactual DGP: IRFs to a monetary policy shock identified with SVARs

Notes: The solid line in black denotes median impulse response from the estimated VAR(4), large sample size ($T=10,000$), and the shaded area the corresponding 68 percent credible interval. The line in magenta denotes the Fry-Pagan median-target response. The line in red denotes the median impulse response from the estimated DSGE model as in Adolfson et al. (2008). A decrease in exchange rate implies appreciation. All sign restrictions are set only for one period (the shock impact period). Horizontal axis is lag horizon in quarters.

sign restrictions; when the signal associated to the monetary policy shock is weak, the responses are not only due to the monetary policy shock but to a combination of shocks. Our results above indeed showed that when the monetary policy is relatively strong, both identification strategies lead to a better performance in replicating the true responses but still some differences remained unexplained, namely, the responses of output and prices continue to be slightly muted under Cholesky and overestimated under sign restrictions, in comparison to the true ones.

To understand if the “combination of shocks” issue affects our results, we regress the SVAR monetary policy disturbances on each of the DSGE structural disturbances (21 of them). With data from model M1 we find that, in case of Cholesky identification, the median R^2 is highest for the DSGE monetary policy shock (0.76) followed by the domestic markup shock (0.03). In case of sign restrictions, the median R^2 is again highest for the DSGE monetary policy shock (0.50) followed by the domestic markup shock (0.09) and the risk premium shock (0.08). According to the DGP in hand, the domestic markup shock contributes to 65 percent of the FVED of prices and 6.7 percent of the FVED of the exchange rate. On the other hand, the risk premium shock explains 54 percent of the FVED of the exchange rate. This observation suggests that if the SVAR-monetary policy shock is a linear combination of different DSGE shocks, then the loadings of the domestic markup shock and of the risk premium shock should be important (after the loadings of the true monetary policy shock), especially for the responses of prices and that of the exchange rate. To check for this hypothesis, we mute these shocks in the DSGE model and estimate the SVARs with the respective counterfactual DGPs.

We show the responses obtained with data from model M1 with muted domestic markup shock in panel (c) and (d) of Figure 6. In comparison to the baseline model, the Cholesky responses of output, interest rate and exchange rate replicate the true responses very well. The “price puzzle” is severely reduced in comparison to the baseline model, for all quarters. Improvement is observed for the estimation with sign restrictions as well, although not as much as under Cholesky. This result suggests that indeed, the SVAR-identified monetary policy shock contains loadings of the domestic markup shock that induce an increase in prices and a faster depreciation of the exchange rate. This shock appears to “trick” the Cholesky to produce a price puzzle and to “help” it overcome the DOP puzzle.²¹ When this shock is muted, the share of peak appreciations of exchange rate in the first quarter drops slightly to 97 percent. Muting the risk premium shock appears to be more important for the responses of the exchange rate under the identification with sign restriction as it increases considerably the share of peak appreciations in the first quarter (see Table 1, 3rd and 4th column).

All the above observations apply to the SVARs under the MC exercise as well (see Figure 1 in Appendix B and Table 2 below). In addition to the scenarios discussed above, in Table 2 we show how

²¹In the DSGE model, in response to a positive domestic markup shock, output drops on impact, prices increase followed with a rise in interest rate and exchange rate depreciation.

the share of peak appreciations changes when increasing the sample size. We observe that increasing the sample size up to 550 observations improves the ability of the SVARs with Cholesky but still only making it relatively comparable to SVARs with sign restrictions (see Table 2, 1st column). On the other hand, the ability of SVARs with sign restrictions does not clearly increase with the number of the observations, a result in line with the findings of Canova and Paustian (2011). With respect to the counterfactual analysis, we see again that a stronger signal for the monetary policy shock helps the SVARs toward a correct identification of the effect of the monetary policy shock. Furthermore, when the domestic markup shock is muted the SVARs perform worse in identifying the correct quarter with peak appreciation.

Table 2: MC simulation: Share of peak appreciations in the 1st quarter, model M1.

	Baseline	Large MP shock	Muted RP shock	Muted DM shock
Cholesky				
T=150	57.6	82.4	73.9	47.5
T=350	58.2	89.8	80.0	47.9
T=550	68.8	88.0	89.0	58.7
Sign restrictions				
T=150	64.9	74.4	61.6	56.7
T=350	59.8	71.8	59.8	59.3
T=550	66.8	69.1	64.0	60.2

Notes: Large MP shock (data from model M1 with monetary policy shock increased tenfold). Muted RP shock (data from model M1 with muted risk premium shock). Muted DM shock (data from model M1 with muted domestic markup shock). For each sample size the lag is selected with BIC.

To summarize, we find that under our experiment, there are two reasons that hinder the ability of SVARs to recover the true responses even in large sample. First, the signal of the monetary policy shock in the DGP is too weak relative to other shocks. Second, the SVAR-identified monetary policy shock is a linear combination of structural shocks, under both identification schemes. Both issues seem to be correlated, with the latter being more notable when the signal of the monetary policy shock is relatively weaker. Adding then on top sampling uncertainty increases the likelihood of SVARs to produce puzzling results.

3.3 VAR with actual data - Sweden

In the following we investigate what do the two identification schemes suggest for the effect of the monetary policy shock when using real data for Sweden (using the same model and techniques as with the artificial data in the analysis above). Since the DSGE model of Adolfson et al. (2008) is estimated with data for the Swedish economy, we will compare the results from model M1 or M2 and see which one is more likely in the light of our SVAR estimates with actual data. We use the same model and techniques as with the artificial data in the analysis above. We estimate SVARs on quarterly data for 1993Q1 - 2014Q1, with $z_t^{Sweden} = (GDP_t, CPI_t, R_t, S_t)$ being the vector of endogenous variables and

$z_t^* = (GDP_t^*, CPI_t^*, R_t^*)$ the block of exogenous variables. We have chosen this period to avoid the break in monetary policy regime that this country experienced in the beginning of 1993, switching from a fixed exchange rate regime to an inflation targeting regime.²² A detailed description of the data we use is provided in Appendix B. Regarding the foreign economy, in line with the data used for estimation in Adolfson et al. (2008), we use foreign weighted variables across Sweden’s largest trading partners. For the exchange rate, S_t , we use the Effective Total Competitiveness Weights nominal exchange rate. All variables are in logs except for interest rates.

As before, the VAR coefficients are drawn from a normal-inverse-Wishart distribution with flat prior, following Uhlig (2005). In line with Sweden being a small open economy we impose a block ergogeneity on the foreign variables. The lag length selected with BIC is 3. Figure 7 shows the median response (black line) and its 68 percent credible interval from the SVARs with data for Sweden along with the median response from the MC exercise with data from model M1 (green line) and M2 (red line), for the same sample size. The size of the monetary policy shock is normalized such that the impact response of the nominal interest rate is 100 basis point, at an annualized rate.

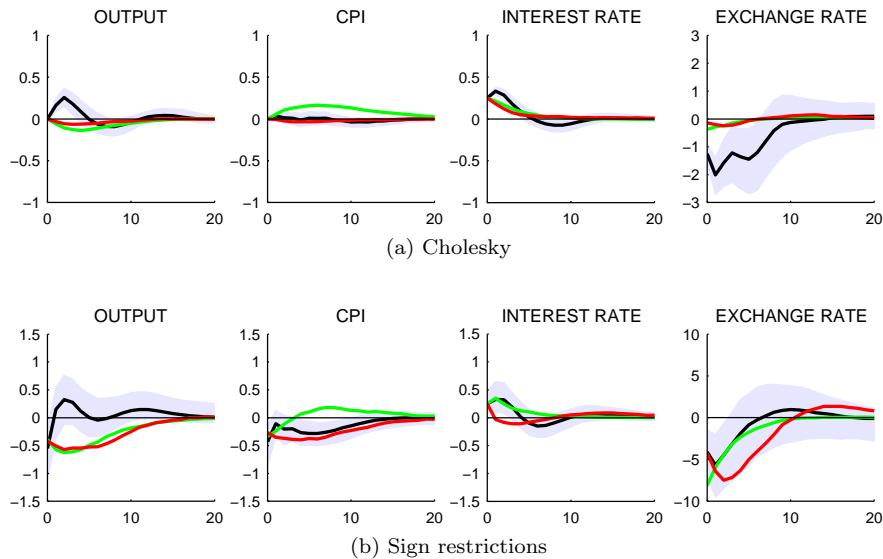


Figure 7: Sweden: Impulse Responses to a monetary policy shock

Notes: Results from the estimated SVAR(3) with data for Sweden, 1993Q1 - 2014Q1, and respective sample size SVARs with simulated data from model M1 and M2. The black line is the median and the shaded area its 68 percent credible interval from the SVARs with data for Sweden. The median response from the MC exercise with data from model M1 is in green and for model M2 in red. A decrease in exchange rate implies appreciation. Horizontal axis is lag horizon in quarters.

In general, for both identification strategies, we observe a muted response of output and in the case of sign restrictions this response is statistically not different from zero. Under Cholesky one observes

²²We admit that the best would be to use the same sample as in Adolfson et al. (2008), starting in 1984, but because of the regime change (which we cannot control for in the type of VAR as used in the simulation exercise) we start in 1993. In order to have enough observations, the end of the sample is extended until the latest available data.

not only a ‘price puzzle’ but even an ‘output puzzle’ for the first quarters.²³ On the other hand, the exchange rate reacts strongly to the monetary policy shock. Under both identification schemes, a monetary contraction is followed by an immediate appreciation of the exchange rate and the peak appreciation appears to be distributed around the second quarter (see Figure 8).

SVAR responses with real data show as well little success in resembling the expected theoretical responses to a monetary policy shock. In the case of Cholesky identification, the SVAR responses resemble rather a demand shock, or more precisely an anticipated demand shock to which the monetary authority responds by raising interest rate before output and prices rise. The sign restriction seems to do a better job which is not surprising because of the restrictions. Still, the response of output becomes insignificant right after the first period and fluctuates around zero thereafter. The response of prices is more significant and has the expected sign in the course of the second year after the shock.

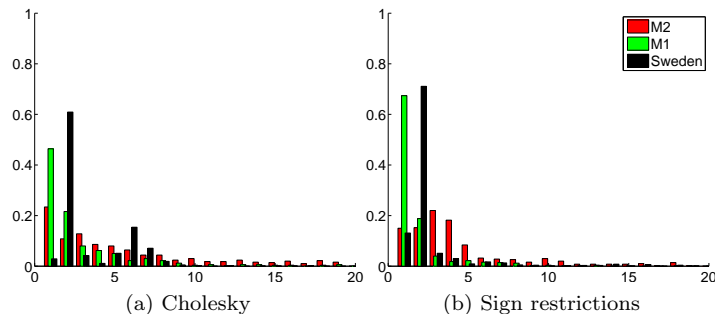


Figure 8: Sweden: Peak appreciation of exchange rate to a monetary policy shock

Notes: Results from the estimated SVAR(3) with data for Sweden, 1993Q1 - 2014Q1, and respective sample size SVARs with data from model M1 and M2. Horizontal axis is lag horizon in quarters.

The difference between the SVAR responses with actual and simulated data may come from different sources. First possibility is that the SVARs could not capture the exogenous shifts in monetary policy on this particular sample.²⁴ Other reasons could be the ones that we studied above; the strength of the monetary policy shock signal might be even weaker than the one estimated in Adolfson et al. (2008) and other stronger shocks contaminate the responses of the SVAR. Finally, the differences might relate to the restrictions imposed in the theoretical model as well. These features are inherited by the pseudo data and may be well captured by our SVARs. In this case for instance the significant negative output response to the monetary policy shocks may come from the model’s features even if it is not in the data.

In order to infer the actual shape of the exchange rate response to the monetary policy shock, we compare the distribution of the peak response for Sweden to those we obtained using the simulated data

²³Using a similar estimation strategy with quarterly data for Sweden (1999-2005), Adolfson et al. (2007) find as well a muted response of output to a identified monetary policy shock, with uncertainty bands including an output puzzle. Carlstrom et al. (2009) have argued that such a puzzle can arise when using identification schemes with zero restrictions.

²⁴Note again that the sample used here is different from that was used in Adolfson et al. (2008) to estimate the DSGE model.

(see Figure 8). The peak distribution for Sweden (black bars) seems to be in-between the distributions for model M1 (green) and M2 (red), suggesting that this statistic cannot distinguish between the two models. To make a more formal inference, we calculate the probabilities of model M2 being the true DGP, conditioned on the median of the estimated distribution of the peak response ("peak median" henceforth). From the simulated distribution of peak medians, we can calculate the probability of M2 being the true model, which is $p_2(i)/(p_1(i) + p_2(i))$, with $p_x(i)$ being the number of simulated samples where the peak median is in the i th quarter with model x ($x=1,2$). We assign a 0.5 prior probability to both models and exclude the possibility that the true DGP is different from both models.

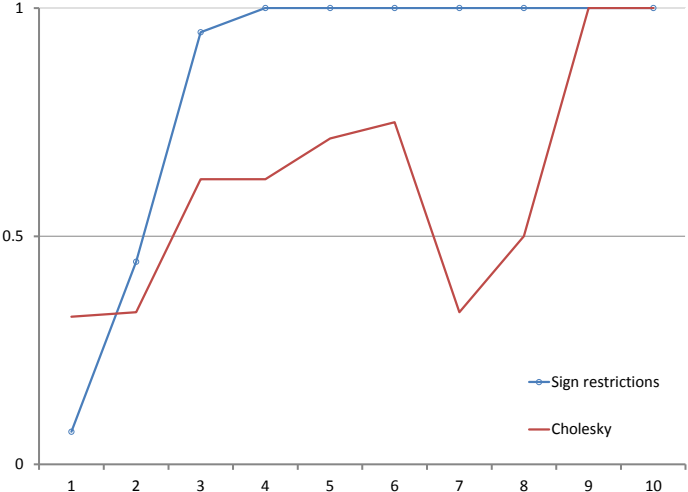


Figure 9: Likelihood of model M2, observing a particular peak median

Notes: Probability of M2 being the true model, $p_2(i)/(p_1(i) + p_2(i))$, with $p_x(i)$ being the number of simulated samples where the peak median is in the i th quarter with model x ($x=1,2$). Results from SVAR estimated on same sample size as the Swedish data. Horizontal axis is lag horizon in quarters.

Figure 9 shows the probabilities of model M2 for both identification scheme. The most obvious difference between the two approaches is that with sign restrictions it is much easier to choose between the two candidate DGPs, because the probabilities are closer to 0 or 1. If, for instance, our estimated peak median is 1 (first quarter), the probability of model M2 with sign restrictions is only 0.07, while with Cholesky it is 0.32. If the peak median is 3, the probability of model M2 is 0.95 with sign restrictions and 0.63 with Cholesky. With Swedish data, we estimated a peak median on the second quarter with both identification strategies, for which model M2 probabilities are 0.33 and 0.44 (Cholesky and sign restriction, respectively), meaning that these results favor slightly model M1, but M2 is almost equally likely.

4 Conclusion

Recent work in the literature suggests that delayed overshooting puzzle or other non-sense results are due to improper identification strategies used in VAR analyses. In this paper, we investigate to what extent various identification restrictions in structural VARs are able to recover the true effect of monetary policy on the exchange rate under a controlled experiment. We estimate VAR models using two sets of artificial observations from an open economy DSGE model of Adolfson et al. (2008) and identify monetary policy shocks by Cholesky ordering and by sign restrictions only. With respect to our data generating process, the Cholesky decomposition imposes improper restrictions while sign-identified VARs allow for model-consistent sign restrictions.

Our results show that when data are abundant, Cholesky-type restrictions perform comparably with schemes relying on sign restrictions when identifying the peak appreciation of the exchange rate. This is true irrespective of whether the data generating process exhibits delayed overshooting or not. Thus, we do not find evidence that zero restrictions imposed by the Cholesky scheme would lead to artificial exchange rate puzzle on long samples, even if they can be challenged by intuition. On the other hand, its performance deteriorates when the number of observations is reduced to typical levels available in applied research. Given sampling uncertainty, the shape of the exchange rate response is estimated with higher uncertainty when using Cholesky ordering. Identification schemes replacing zero restrictions by sign restrictions are more robust to the sampling uncertainty despite the fact that imposing sign restrictions typically brings additional (identification) uncertainty. We also confirm a previous finding in the literature that SVARs performance is poor when trying to identify effects of shocks with weak signal, as it is the case in practice for monetary policy shocks. Supported by this weakness, a SVAR-identified monetary policy shock is often then a linear combination of many structural shocks. Adding sampling uncertainty to these problems increases the likelihood of SVARs to produce puzzling results.

Overall, we believe our exercise adds new contribution to the empirical literature of exchange rate dynamics and the exchange rate channel of monetary transmission mechanism. It provides a test on the claims of recent literature that delayed overshooting of the exchange rate is an outcome of improper identification restrictions in VARs. For econometricians it provides more insight into the performance of different identification strategies. Results suggest that improper restrictions set with the recursive identification scheme may be, but are not necessarily the source for puzzling outcomes. Nevertheless, given estimation uncertainty, identification of monetary policy shocks with Cholesky decomposition, under a traditional VAR estimation, should be avoided.

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A The model as DGP: Adolfson et al. (2008)

The domestic economy in Adolfson et al. (2008) is comprised of households, four types of firms working on the domestic and the external sector, a government and a monetary authority. The foreign economy is modeled exogenously, resembling the typical situation of a small open economy with no power in the world market. In the following we present a description of the main characteristics, decisions and constraints of each of the agents in this economy for the version of the model with standard UIP.

A.1 Households

A continuum of households consumes domestic and imported goods, own capital and financial assets in the form of domestic deposits, foreign bonds and cash balances. A representative household j , maximizes the following utility function:

$$E_0^j \sum_{t=0}^{\infty} \beta^t \left[\zeta_t^c (C_{j,t} - bC_{j,t-1}) - \zeta_t^h A_L \frac{(h_{j,t})^{1+\sigma_L}}{1+\sigma_L} + A_q \frac{\left(\frac{Q_{j,t}}{z_t P_t}\right)^{1-\sigma_q}}{1-\sigma_q} \right] \quad (\text{A.1})$$

subject to the budget constraint given by:

$$\begin{aligned} & M_{j,t+1} + S_t B_{j,t+1}^* + P_t^c C_{j,t} (1 + \tau_t^c) + P_t^i I_{j,t} + P_t K_{j,t} \\ &= R_{t-1} (M_{j,t} - Q_{j,t}) + Q_{j,t} + (1 - \tau_t^k) \Pi_t + (1 - \tau_t^y) \frac{W_{j,t}}{1 + \tau_t^w} h_{j,t} \\ &+ (1 - \tau_t^k) R_t^k \bar{K}_{j,t} + R_{t-1}^* \Phi\left(\frac{A_{t-1}}{z_{t-1}}, \tilde{\phi}_{t-1}\right) S_t B_{j,t}^* \\ &- \tau_t^k \left[(R_{t-1} - 1) (M_{j,t} - Q_{j,t}) + \left(R_{t-1}^* \Phi\left(\frac{A_{t-1}}{z_{t-1}}, \tilde{\phi}_{t-1}\right) - 1 \right) S_t B_{j,t}^* + B_{j,t}^* (S_t - S_{t-1}) \right] \\ &+ TR_t + D_{j,t} \end{aligned} \quad (\text{A.2})$$

with $C_{j,t}$, $h_{j,t}$ and $\frac{Q_{j,t}}{z_t P_t}$ being the level of consumption, the work effort and real balances for the household j , respectively. Consumption and work efforts are subject to consumption preference shocks, ζ_t^c and labor supply preference shocks, ζ_t^h . Households supply differentiated labor service and set their wage, subject to \acute{a} la Calvo rigidity. Under such rigidity, households index their wages to past CPI inflation, current inflation target and to technology growth. Households earn R_{t-1} interest rate on domestic deposits and a risk-adjusted pre-tax gross rate interest, $R_{t-1}^* \Phi\left(\frac{A_{t-1}}{z_{t-1}}, \tilde{\phi}_{t-1}\right)$, in foreign bond holdings.

Aggregate consumption and total investment are given by a CES index of domestically produced and imported goods, as below:

$$C_t = \left[(1 - \omega_c)^{1/\eta_c} (C_t^d)^{(\eta_c-1)/\eta_c} + \omega_c^{1/\eta_c} (C_t^m)^{(\eta_c-1)/\eta_c} \right]^{\eta_c/\eta_c-1} \quad (\text{A.3})$$

$$I_t = \left[(1 - \omega_i)^{1/\eta_i} (I_t^d)^{(\eta_i-1)/\eta_i} + \omega_i^{1/\eta_i} (I_t^m)^{(\eta_i-1)/\eta_i} \right]^{\eta_i/\eta_i-1} \quad (\text{A.4})$$

where C_t^d , C_t^m and I_t^d , I_t^m are domestic and imported consumption and investment goods, respectively. The share of imports on consumption and investment is given by ω_c and ω_i , while η_c , η_i represent the elasticity of substitution between domestic and foreign goods.

Households own physical capital, K_t , and decide how much of it to rent to the domestic firms given costs of adjusting the investment rate, $\bar{S}(I_t, I_{t-1})$. The law of motion for \bar{K}_t is given by:

$$K_{t+1} = (1 - \delta) K_t + \Upsilon_t (1 - \bar{S}(I_t, I_{t-1})) \quad (\text{A.5})$$

with Υ_t being a stationary investment-specific technology shock.

Each household solves its maximization problem with respect to consumption, cash holdings, physical capital, investments and foreign bond holdings. Combining the f.o.c for cash and for foreign bond holdings (after log-linearization) we get the following uncovered interest parity condition:

$$\hat{R}_t - \hat{R}_t^* = E_t \Delta \hat{S}_{t+1} - \tilde{\phi}_a \hat{a}_t + \tilde{\phi}_t \quad (\text{A.6})$$

A.2 Firms

There are four different firms: domestic goods firms, importing consumption, importing investment, and exporting firms. They all produce a differentiated good and have monopolistic power when setting prices. Price setting is subject to *à la* Calvo rigidities. The four final goods are each a CES (constant elasticity of substitution) composite of respective differentiated goods. For the final domestic good, the CES composite would be as follows:

$$Y_t = \left[1_0^{\int} (Y_{i,t})^{1/\lambda_t^d} di \right]^{\lambda_t^d}, 1 \leq \lambda_t^d < \infty, \quad (\text{A.7})$$

with λ_t^d being the time-varying flexible-price markup shock. The demand for firm *i*'s differentiated product follows:

$$Y_{i,t} = \left(\frac{P_{i,t}^d}{P_{i,t}^d} \right)^{\frac{-\lambda_t^d}{\lambda_t^d - 1}} Y_t \quad (\text{A.8})$$

Domestic goods firms produce their differentiated goods using capital and labor inputs under the following technology:

$$Y_{i,t} = z_t^{1-\alpha} \epsilon_t K_{i,t}^\alpha H_{i,t}^{1-\alpha} - z_t \phi \quad (\text{A.9})$$

where $K_{i,t}$ is the capital stock, $H_{i,t}$ the homogeneous labor hired by the *ith* firm, z_t a unit-root technology shock common to the domestic and foreign economies and ϵ_t is a domestic covariance stationary technology shock. Cost minimization for an intermediate firm *i* is given by:

$$MC_t^d = \frac{1}{(1-\alpha)^{1-\alpha}} \frac{1}{\alpha^\alpha} (R_t^k)^\alpha [W_t(1 + \nu(R_{t-1} - 1))]^{1-\alpha} \frac{1}{(z_t)^{1-\alpha}} \frac{1}{\epsilon_t} \quad (\text{A.10})$$

where R_t^k is gross nominal rental rate of capital, R_{t-1} the gross nominal interest rate, and ν the fraction of the wage bill that a firm *i* finances in advance through loans to financial intermediary.

The price setting problem is the same for all the four firms, resulting in four specific Phillips curve equations determining inflation for domestic consumption, import consumption, import investment and export sectors. As an example we give the price setting decision of the domestic goods firms. Each of these firms is subject to price stickiness *à la* Calvo where with a probability, $(1 - \xi_d)$, the firm can reoptimize its price, $P_t^{d,new}$, in any period. Firms that cannot reoptimize every period, index the new price to the current inflation target and to last period's inflation rate. Optimization problem of an individual firm is then:

$$\max_{P_t^{d,new}} E_t \sum_{s=0}^{\infty} (\beta \xi_d)^s v_{t+s} ((\pi_t \pi_{t+1} \dots \pi_{t+s-1})^{\kappa_d} (\bar{\pi}_{t+1}^c \bar{\pi}_{t+2}^c \dots \bar{\pi}_{t+s}^c)^{1-\kappa_d} P_t^{d,new}) Y_{i,t+s} - MC_{i,t+s}^d (Y_{i,t+s} + z_t \phi) \quad (\text{A.11})$$

The log-linearized Phillips curve for the domestic good producing firm is:

$$\begin{aligned}
(\hat{\pi}_t^d - \hat{\pi}_t^c) &= \frac{\beta}{1 + \kappa_d \beta} (E_t \hat{\pi}_{t+1}^d - \rho_\pi \hat{\pi}_t^c) + \frac{\kappa_d}{1 + \kappa_d \beta} (\hat{\pi}_{t-1}^d - \hat{\pi}_t^c) \\
&\quad - \frac{\kappa_d \beta (1 - \rho_\pi)}{1 + \kappa_d \beta} \hat{\pi}_t^c + \frac{(1 - \xi_d)(1 - \beta \xi_d)}{\xi_d (1 + \beta \kappa_d)} (\hat{m}c_t^d - \hat{\lambda}_t^d)
\end{aligned} \tag{A.12}$$

Firms that produce differentiated importing consumption and investment goods, use a brand naming technology to convert the homogeneous goods bought in the world market at price P_t^* , with marginal cost $S_t P_t^*$. The exporting goods firms use this technology to convert homogeneous domestic goods in differentiated goods to sell in the world market, with marginal cost P_t^d/S_t . Since these firms face nominal rigidities a la Calvo when setting prices in local currency, a short run incomplete exchange rate pass-through is induced.

A.3 Monetary Policy

Central bank is assumed to follow an instrument rule, where short term interest rate is adjusted in response to deviations of CPI inflation from the time-varying inflation target, the output gap (measured as actual minus trend output), the real exchange rate, \hat{x}_t , and the interest rate set in the previous period. In log-linearized form, the rule is given as:

$$\hat{R}_t = \rho_{R,t} \hat{R}_{t-1} + (1 - \rho_{R,t}) [\hat{\pi}_t^c + r_{\pi,t} (\hat{\pi}_{t-1}^c - \hat{\pi}_t^c) + r_{y,t} \hat{y}_{t-1} + r_{x,t} \hat{x}_{t-1}] + r_{\Delta\pi,t} \Delta \hat{\pi}_t^c + r_{\Delta y,t} \Delta \hat{y}_t + \epsilon_R^t \tag{A.13}$$

The government consumes domestic good, collects taxes from households and transfers back to them any surplus or fiscal deficit. The government together with the foreign economy are estimated ahead of the DSGE model with identified VAR models and then exogenously put in the the DSGE model.

In equilibrium the following constraints should hold for clearance in the final goods market, the foreign bond market, and the loan market for working capital:

$$C_t^d + I_t^d + G_t + C_t^x + I_t^x \leq z_t^{1-\alpha} \epsilon_t K_{i,t}^\alpha H_{i,t}^{1-\alpha} - z_t \phi \tag{A.14}$$

$$S_t B_t^* = S_t P_t^x (C_t^x + I_t^x) - S_t P_t^* (C_t^m + I_t^m) + R_{t-1}^* \Phi(a_{t-1}, \tilde{\phi}_{t-1}) S_t B_t^* \tag{A.15}$$

$$\nu W_t H_t = \mu_t M_t - Q_t \tag{A.16}$$

Since the economy in this model (both versions) is subject to a unit root technology shock, z_t , the model is made stationary by scaling the real quantities with it. The model is log-linearized around a constant steady state. Both model M1 and M2 are estimated with Bayesian techniques using Swedish data for the period 1980Q1-2004Q4 using a vector of 15 observables.

B Empirical analysis

Table B.1: Selected papers on the effect of monetary policy on the exchange rate with VARs

Study	Countries*	Period	Identification	Results
Eichenbaum and Evans (1995)	Germany, UK, Japan	1974-1990 (monthly)	recursive	DOP, up to 3 years
Scholl and Uhlig (2008)	Germany, UK, Japan	1975-2002 (monthly)	sign restric- tions DOP, 1 to 2 years	
Peersman and Smets (2003)	aggregate Euro area	1980-1998 (quarterly)	recursive, non-recursive, long-run restriction	DOP, up to 1 year
Cushman and Zha (1996)	Canada	1974-1993 (monthly)	non-recursive	No DOP
Kim and Roubini (2000)	G7	1974-1992 (Canada, 1974-1992) (monthly)	recursive, non- recursive	Exchange rate puzzle
Faust and Rogers (2003)	Germany, UK	1974-1997 (monthly)	sign restric- tions	Inconclusive
Bjornland (2009)	Australia, Canada, New Zealand, Sweden, UK	1983-2004 (quarterly)	recursive, long-run restriction	DOP, No DOP

Notes: *Bilateral rates with US. DOP - Delayed Overshooting Puzzle.

B.1 Data description for Sweden

The main source of the data is the OECD database and when not, it is stated.

GDP_t - Gross Domestic Product, chained volume estimates, s.a. with X11 ARIMA.

CPI_t - Consumer Price Index, 2005=100.

R_t - Three month interbank rate, per cent per annum.

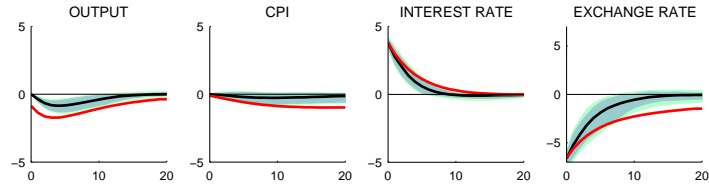
S_t - Nominal effective exchange rate (TCW), 1992 = 100. Source: Riksbank.

GDP_t^* - Foreign GDP (TCW). Source: OECD and authors calculations.

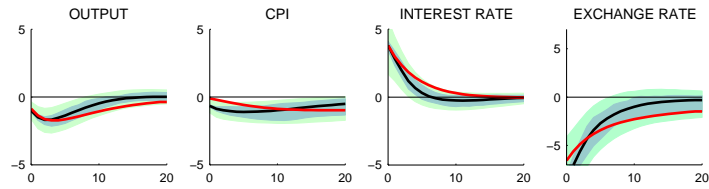
R^* - Foreign interest rate (TCW). Source: OECD and authors calculations.

CPI^* - Foreign Consumer Price Index (TCW). Source: OECD and authors calculations.

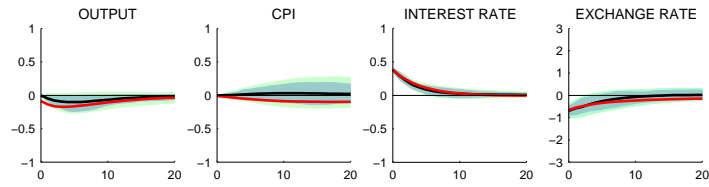
TCW - Total Competitiveness Weights index, a set of trade weights for 20's largest trade partners of Sweden. Source: IMF.



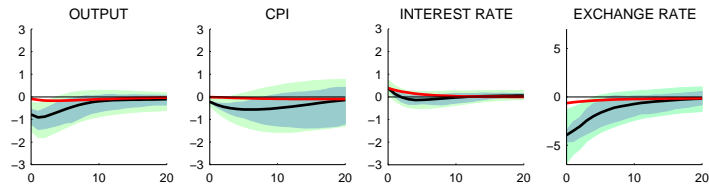
(a) Large MP shock: Cholesky



(b) Large MP shock: Sign restrictions



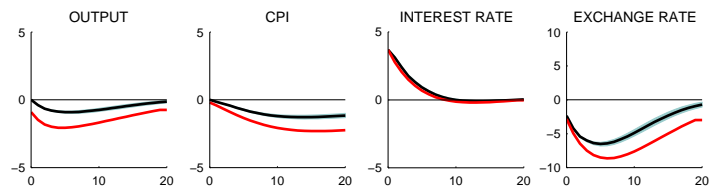
(c) Muted DM shock: Cholesky



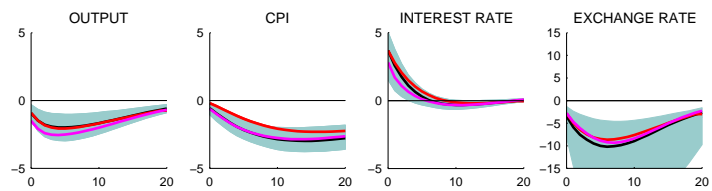
(d) Muted DM shock: Sign restrictions

Figure B.1: MC exercise, Counterfactual DSGE, M1 model

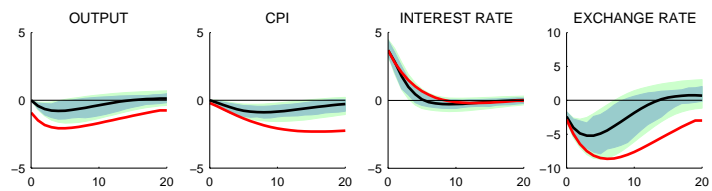
Notes: The solid line in black denotes the pointwise median of impulse responses from all samples. The shaded area in light blue indicates the 16th and 84th percentile region. The deep blue area corresponds to the 68 percent distribution of the pointwise sample medians. The VAR for each sample is estimated with $T=150$ and lag selected with BIC and the chosen lag is 3. The line in red denotes the true impulse response from the DSGE model. A decrease in exchange rate implies appreciation. All sign restrictions are set only for one period (the shock impact period). Horizontal axis is lag horizon in quarters.



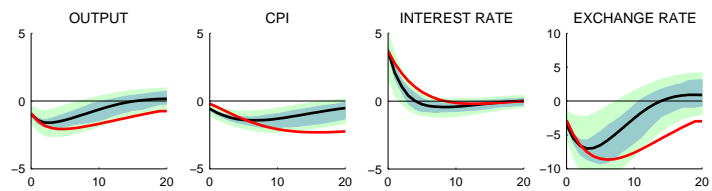
(a) Long sample: Cholesky



(b) Long sample: Sign restrictions



(c) Short sample: Cholesky



(d) Short sample: Sign restrictions

Figure B.2: Monetary policy shock 10 fold higher in model M2

Notes: The line in red denotes the true impulse response from the DSGE model. A decrease in exchange rate implies appreciation. All sign restrictions are set only for one period (the shock impact period). Horizontal axis is lag horizon in quarters.

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