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NOVEMBER 2003

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Monetary Policy Shocks and Business Cycle Fluctuations in a Small Open Economy: Sweden 1986-2002

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Sveriges Riksbank Working Paper Series
No. 153
November 2003

Abstract

This paper contains an empirical analysis of the dynamic effects of monetary policy on Swedish data within a framework consistent with the theoretical New-Keynesian type of small open economy models. Because of what appears to be time-varying seasonal patterns in the data, I argue that it is of crucial importance to use the annual inflation rate rather than the quarterly inflation rate in the empirical analysis. After a monetary policy shock, the impulse response functions for output and inflation display a "hump-shaped" pattern with peak effects after $1.5 - 2$ years. There also seems to be considerable inertia in the real exchange rate. Sensitivity analysis suggests that the shape of the obtained impulse response functions is fairly robust with respect to the number of lags in the VAR, sample size, and the formulation of the policy rule. Also, we find evidence that foreign shocks are very important for understanding Swedish business cycles. In particular, they account for a large fraction of the lower frequency movements in output and inflation, whereas domestic shocks generate most of the high frequency movements in the data.

Keywords: Monetary policy shocks; Impulse response functions; VAR models; New-Keynesian models; Real exchange rates; Business cycles.

JEL Classification Numbers: E52, C52, C22.

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I am indebted to Anders Vredin for very helpful discussions and suggestions. Discussions and comments by Malin Adolfson, Tor Jacobson, Marianne Nessén and Ulf Söderström have also been very helpful. Special thanks to Daniel Hammarsten and Josef Svensson for collecting the data. The paper has also benefitted from comments by seminar participants at Sveriges Riksbank. The views expressed in this paper are those of the author and should not be interpreted as reflecting the views of the Executive board of Sveriges Riksbank.

1 Introduction

The effects of monetary policy shocks have been studied extensively on data for the U.S. economy, see e.g. Leeper, Sims and Zha (1996) and Christiano, Eichenbaum and Evans (1999). The consensus in this literature seems to be that a contractionary shock to monetary policy (i.e. an unanticipated increase in the fed funds rate) leads to "hump-shaped" decreases in output and inflation with peak responses after 1-2 years' time. In empirical studies of open economies, there seems to be less consensus about the effects of monetary policy shocks, see e.g. Jacobson et al. (2002) and the references therein.

One natural question to ask is why it is interesting to study the effects of monetary policy shocks empirically. One obvious answer is that given the institutional setup in many countries where central banks have been assigned the task of stabilizing inflation using some kind of nominal interest rate as instrument, we want to understand if and how monetary policy affects the economy. In addition to establishing a nominal anchor (i.e an inflation target level), we want to know if movements in nominal interest rates matter for the short-run movements in aggregate quantities and prices. Unfortunately, theory cannot provide a clear cut answer to this question because it is possible to write down plausible theoretical models with very different implications regarding the real effects of monetary policy. This is the reason why I want to examine what the data suggest. Another answer to the question is that the data seem to reveal a lot of information that can be used to distinguish between competing theoretical models. Christiano, Eichenbaum and Evans (2001) argue that in order for a dynamic general equilibrium model to fit the estimated impulse response functions to a monetary policy shock, it needs to include adjustment costs to capital, habit persistence in consumer preferences, variable capacity utilization and sticky wages (Calvo style wage setting). Interestingly, the same set of parameters that makes this model able to replicate the dynamic effects of a monetary policy shock on US data, is also shown by Altig et al. (2002) to be able to replicate the dynamic effects of a permanent technology shock too. This finding suggests that impulse response functions to a monetary policy shock are useful to unravel the underlying structure of the economy.[1] Another finding by Altig et al. (2002) is that monetary policy shocks account for a small fraction of the variation in the data, the identified technology shocks account for about 50 percent of the fluctuations in output, whereas

[1] Of course, it is possible that more than one theoretical model can produce impulse response functions to a monetary policy shock that are in line with those in the unrestricted VAR. So one cannot, of course, draw the conclusion that the model at hand is the true model, but at least conclude that it is not inconsistent with the data. Although this latter conclusion may not sound that strong, it has turned out to be difficult to write down models with microeconomic foundations that can replicate the effects of monetary policy shocks in the data.

the policy shocks account for less than 5 percent, suggesting that what is interesting to study once equipped with a "good" model is the systematic response of monetary policy to the state of the economy (i.e. the systematic part of the "policy rule").

Jacobson et al. (2002) estimate a VAR model on Swedish data and use standard open economy assumptions to identify the effects of a shock to monetary policy. They obtain "price puzzles" (see Sims, 1992) when using standard techniques to compute impulse response functions to a policy shock, whereas when they use a "narrative" approach to identify policy shocks (i.e. estimating the effects of the large devaluations in 1981, 1982, and 1992) they obtain results that are more in line with the results on U.S. data. In this paper, I argue that VAR models on Swedish data that include a price index in levels (as in the case in Jacobson et al. 2002) or the first difference of the price level - i.e. the quarterly inflation rate - may produce misleading results about the dynamic effects of monetary policy. The reason for this is that these series seem to contain a lot of variation - although deseasonalized with the X12-method - which cannot simply be explained either by including more lags or by introducing more variables in the VAR. Although the series have been deseasonalized, simple regression analysis reveal clear signs of remaining time-varying seasonal variation in the quarterly inflation rate in Swedish data.[2]

One possible way to mitigate the effects of changing seasonal patterns is to use the fourth difference of the price level, i.e. the annual instead of the quarterly inflation rate, in the empirical model. I demonstrate that for a standard VAR model estimated on U.S. data, the use of the annual inflation rate rather than the annualized quarterly inflation rate leads to basically identical impulse response functions for the other variables (e.g. output) after a shock to monetary policy, whereas the peak response of inflation gets smoothed out with a couple of quarters' delayed peak response. By generating artificial data from a simple theoretical model, I also show that these findings are completely in line with what one would have expected in an economic model as long as inflation is positively autocorrelated with an autocorrelation coefficient that is in line with what we observe in the data. In contrast, the shape of the impulse response functions for output and inflation on Swedish data are very different when the annual inflation rate is used in the VAR instead of the quarterly inflation rate. However, the results on Swedish data are well in line with those obtained for the U.S. when the annual inflation rate is used: the responses of output and inflation to a monetary policy shock are both hump-shaped with peak

[2] An additional important difference to the study by Jacobson et al. (2002) is that the benchmark sample period is restricted to $1986Q1 - 2002Q4$, due to successively deregulated financial markets and recurrent devaluations during the 1970's and beginning of the 1980's. See Berg et al. (1993) for a chronological exposition and discussion of the deregulations' in the Swedish financial markets during the 1970's and 1980's.

responses after $1.5 - 2$ years, whereas the effect on the interest rate are much less persistent. Based on the evidence reported in the paper, my conjecture is that remaining time-varying seasonal variation in the inflation rate accounts for these large qualitative differences in the impulse responses to a policy shock, and that we should have confidence in the ones generated using the annual inflation rate in the VAR.

Another interesting finding in the empirical analysis is that it is of crucial importance to condition on foreign variables in the VAR model, but at the same time we find that the foreign variables we include (the trade weighted foreign output, inflation and short-term nominal interest rate) are not affected by the Swedish economy. Consequently, the small economy assumption that domestic shocks do not affect the foreign economy which is typically imposed in theoretical models of small open economies is supported by the data. This result enables us to investigate the relative importance of foreign and domestic shocks for Swedish business cycles. I find that foreign shocks are an important source of business cycles. In particular, they seem to account for most of the lower frequency movements in output and inflation, whereas domestic shocks are more important for high frequency movements. As in Altig et al. (2002), I provide evidence that monetary policy shocks have not contributed much to macroeconomic fluctuations. The exception being the year 1995, when the Riksbank pursued a tight monetary policy. This policy led to a significant contraction in economic activity and a drop in inflation during the subsequent years $1996 - 1997$. The conclusion that monetary policy was geared toward a strong tightening during this period is in line with previous results reported by Berg, Jansson and Vredin (2002) who use a completely different empirical approach.[3] Rudebusch (1998) criticizes VAR models estimated on U.S. data for their inability to produce identified policy shocks that are in line with common beliefs about the stance of monetary policy during certain episodes, but interestingly that critique does not appear to be valid here.

The structure of the paper is as follows. In the next section, I present the VAR model and data that are used in the paper. I draw particular attention to data problems when measuring inflation on quarterly data for the Swedish economy. In Section 3, I discuss the dynamic effects of monetary policy shocks. Section 4 reports the sources behind Swedish business cycles $1986 - 2002$. Finally, some concluding remarks are provided in Section 5.

[3] More specifically, Berg et al. (2002) compute the policy shocks as the residuals in a regression of the short-term interest rate (that the Riksbank controls) using the Riksbank's own real-time forecasts of inflation and output as regressors. During the year 1995, they record a series of fairly large positive residuals which they, along with minute statements, interpret as a tightening of monetary policy. There is no single economic story that have been put forward as an explanation behind this tightening, but one possible candidate is that the Riksbank wanted to establish credibility for the inflation target that was in effect from January 1, 1995.

2 The VAR model

2.1 Basic specification

The VAR(p) model with p lags is specified as

$$X_t = C + \delta_1 D_{701854} + \delta_2 D_{923} + \delta_3 D_{931024} + \tau T_t + \sum_{i=0}^{p} \Upsilon_i Z_{t-i} + \sum_{i=1}^{p} \Gamma_i X_{t-i} + \varepsilon_t \tag{1}$$

where D_{701854} is a dummy variable equal to 1 $1970Q1 - 1985Q4$ and 0 otherwise, D_{923} is a dummy variable equal to 1 $1992Q3$ and 0 otherwise, D_{931024} is a dummy variable equal to 1 $1993Q1$ and thereafter, T_t is a linear timetrend, and Z_t is a vector with exogenous variables. During the 1970's and first half of 1980's, the Swedish financial markets were highly regulated and the dummy variable D_{701854} is intended to capture potential regulation effects, whereas the dummy variable for the third quarter in 1992 is included to capture the exceptionally high interest rate increase (up to 500 percent) implemented by the Riksbank in order to defend the fixed Swedish exchange rate. Despite the efforts to defend the Swedish krona, Sweden entered into a floating exchange rate regime in late November 1992, and the dummy variable D_{931024} is included in order to capture possible effects of the new exchange rate regime. Choice of lag length is discussed in Section 2.3.

The variables in X_t and Z_t are

$$X_t = \begin{bmatrix} y_t & \pi_t & R_t & Q_t & \pi_t^{imp} \end{bmatrix}' \tag{2}$$

and

$$Z_t = \begin{bmatrix} y_t^* & \pi_t^* & R_t^* \end{bmatrix}'.$$

where y_t^* denotes the foreign trade-weighted (TCW) GDP to market prices (seasonally adjusted), π_t^* is a measure of foreign trade-weighted (CPI) inflation, R_t^* is the foreign trade-weighted 3-month nominal interest rate, y_t is domestic GDP in 1995 years' prices (seasonally adjusted in logs additively, allowing for breaks in the seasonal pattern in $1979Q1$, $1985Q1$, $1990Q1$ and $1997Q1$), π_t is the inflation rate on domestic goods (measured with the GDP deflator, see below), R_t is the nominal REPO interest rate (or its equivalence prior to June 1, 1994), Q_t is the real trade-weighted exchange rate, and π_t^{imp} is the inflation rate on imported goods. All variables are measured in logs except inflation and interest rates. The theoretical foundation behind the choice of these variables is the New-Keynesian type of small open economy models outlined in e.g. Clarida, Galí and Gertler (2002), and Svensson (2000), and the variables in X_t are similar to those included by Smets and Wouters (2002) who estimate a VAR model for the Euro area. The hope is that these variables suffice to approximate the reduced-form equilibrium in the theoretical model. Jacobson et al. (2002), who estimate a VAR on Swedish data for the period 1970-2000 use a different theoretical model - and thus a different set of variables - to impose restrictions on the VAR.

Giordani (2001) argues that instead of including output in levels as measures of y_t and y_t^* in (2), we should either (i) include the deviation of output from trend (i.e. the "output gap") for these variables, or (ii) include output in levels along with the trend level of output in the VAR. If one incorrectly includes only output in levels in the VAR, Giordani shows that one might obtain misleading results if the data generating process is a New-Keynesian type of model with a time-varying trend output. In particular, Giordani argues that a so-called "price puzzle" (i.e. the price level goes up instead of down for a long time following an unexpected tightening of monetary policy) might arise. However, a practical problem that Giordani does not address is how to compute trend output (and thus also the "output gap") in the data. By including a linear trend in the VAR model (1) along with output in levels I try to address this problem by modelling the trend implicitly in the VAR.[4]

The specification of the VAR-model in (1) imposes the strict small open economy (SOE) assumption that domestic shocks in the small open economy (i.e. Sweden) have no effects on the world economy. This SOE assumption is strongly supported by the data. A block-exogeneity test, which can be thought of as a multivariate Granger-causality test, clearly shows that the domestic variables are highly dependent on the foreign variables.[5] The p-value for the null hypothesis of block exogeneity of X_t w.r.t. Z_t is 0.01 for the sample period $1986Q1 - 2002Q4$. At the same time, I find that the foreign variables are not dependent on the domestic variables, the p-value for a block exogeneity test of Z_t w.r.t. X_t is 0.99. Consequently, there is strong support for the conventional SOE assumption for this sample period. It should be pointed out that these p-values are for the preferred lag length in the VAR (see below), but the support for the SOE assumption from the block-exogeneity test results are robust to perturbations of the

[4] I have experimented by estimating (1) with Hodrick-Prescott filtered y and y^* (setting the smoothing coefficient to 1600) included in the system, but the results were found to be qualitatively similar.

[5] The block-exogeneity test of $X_{1,t}$ w.r.t. $X_{2,t}$ is carried out as follows: (i) Estimate the VAR(p) model $X_{1,t} = c + B(L)X_{1,t-1} + e_{1,t}$ where p is the length of the polynomial $B(L)$ and c is a vector with constant terms and compute the determinant of the estimated covariance matrix for $e_{1,t}$, $\mathrm{d}_1 \equiv \det\left(\hat{\Sigma}_1\right)$. ($ii$) Estimate the VAR($p$) model $X_{1,t} = c + B(L)X_{1,t-1} + CL)X_{2,t-1} + e_{2,t}$ and compute the determinant of the covariance matrix for $e_{2,t}$, $\mathrm{d}_2 \equiv \det\left(\hat{\Sigma}_2\right)$. ($iii$) Then the test statistic $T \cdot \log(\mathrm{d}_1/\mathrm{d}_2)$ follows the χ^2-distribution with $n_{x_1} \cdot n_{x_2} \cdot p$ degrees of freedom. See Hamilton (1994) pages 311-313 for further details.

lag length in the VAR between 1 and 4 lags.[6] These results imply that an analysis with the foreign variables included in X_t, thus allowing for potential spillover from domestic shocks to abroad, i.e.

$$X_t = \left[\begin{array}{cccccccc} y_t^* & \pi_t^* & R_t^* & y_t & \pi_t & R_t & Q_t & \pi_t^{imp} \end{array} \right]', \tag{3}$$

yields very similar results as the ones that are presented in this paper.

There is no explicit modeling of cointegrating relationships in the estimated VAR model given by (1) and (2), so the presumption is that the VAR conditional on the trend and foreign variables is stationary. For the benchmark specification, this presumption is strongly supported by the data, the largest root of the characteristic polynomial being 0.89 in modulus which is well below unity.[7]

As already mentioned in the introduction, due to changes in the economic environment during the 1970's and beginning of the 1980's, e.g. recurrent devaluations and successively deregulated financial markets, the benchmark estimation period for the VAR-model in (1) is restricted to the period $1986Q1 - 2002Q4$. We will nevertheless do some robustness checks of the results by expanding the sample backwards in time.

2.2 Measuring the inflation rate

Before we present the results from the VAR estimations, we have to decide how we should measure the inflation rates π_t^* and, in particular, π_t on available data. Some studies include the price level rather than the inflation rate in the VAR, and then back out the effects on the inflation rate. This is the approach chosen by e.g. Jacobson et al. (2002). In other studies in the literature that use inflation rates instead, typically the annualized first difference, i.e.

$$\tilde{\pi}_t = 4 \ln \left(\tilde{P}_t / \tilde{P}_{t-1} \right), \tag{4}$$

is used where $\tilde{P}_t$ is the seasonally adjusted price level, see e.g. Christiano, Eichenbaum and Evans (2001). An alternative to measure inflation is to use the fourth difference of the seasonally

[6] In particular, this is always true for block-exogeneity of Z_t w.r.t. X_t. When 4 lags is used, the p-value for the block-exogeneity of X_t w.r.t. Z_t goes up to 0.50 from 0.12 using 3 lags. But in this case, there are very few degrees of freedom because there are 66 observations in the sample and 8 variables in the VAR, and hence large uncertainty about the power of the test. Moreover, although block-exogeneity of the domestic variables to the foreign shocks cannot be statistically rejected in this case, the foreign variables are still a very important source of fluctuations in the domestic variables.

[7] More specifically, let $\Pi = \left[\begin{array}{cc} \Gamma_1 & \Gamma_2 \\ \mathrm{I}_n & 0_n \end{array} \right]$ where n is the number of variables in the VAR, and let λ_j, $j = 1, ..., np$, denote the eigenvalues of the Π-matrix. Then the modulus for eigenvalue j, denoted $|\lambda_j|$, is computed as $\sqrt{\left(\lambda_j^r\right)^2 + \left(\mathrm{real}\, \lambda_j^i\right)^2}$ where λ^r and λ^i denote the real and imaginary part of eigenvalue j respectively.

unadjusted price level, i.e.

$$\bar{\pi}_t = \ln\left(P_t / P_{t-4}\right), \tag{5}$$

which is approximately equivalent to a 4-quarter moving average of $\tilde{\pi}_t$. Note that $\bar{\pi}_t$ can also be seen as a particular seasonal adjustment of the price level P_t where the adjustment coefficients change every quarter since $\left(1 - L^4\right) = (1 - L)\left(1 + L + L^2 + L^3\right)$.

In this section, I will show that the choice between using the first difference, i.e. (4), or the fourth difference (5) does not cause any problems for identifying monetary policy shocks for the U.S., but that it is of crucial importance when pinning down the effects of a shock to monetary policy on Swedish data. In the end, I will argue that one should measure inflation on Swedish data with the annual inflation rate rather than the annualized quarterly inflation rate.[8]

Figure 1 shows a plot of $\bar{\pi}_t$ and $\tilde{\pi}_t$ in the U.S. and in Sweden, in both countries measured with the CPI.[9] We immediately see that the similarity between the two measures is much higher in the U.S. than in Sweden. The standard deviation for the difference of the series, i.e. $\tilde{\pi}_t - \bar{\pi}_t$, is also much higher in Sweden (2.22 percent) than in the U.S. (0.88 percent). If inflation is measured with the GDP deflator, the difference is even larger (2.77 and 0.57 percent in Sweden and the U.S., respectively).

In Figure 2, we show impulse response functions of output and inflation (solid line $\pm$ 2 standard deviations indicated by the dashed lines) for a shock to monetary policy in the U.S. using both the annualized first difference of GDP deflator and the fourth difference of the GDP deflator as measure of inflation. We identify the policy shock using the recursiveness assumption outlined in Christiano, Eichenbaum and Evans (2001).[10] From the figure, we see that the effects on output are almost identical, whereas the effects on inflation in the latter case is smoothed out somewhat as expected. Consequently, it does not matter for the identification of policy shocks in the U.S. whether you use the first of the fourth difference of inflation in the VAR. To verify this latter statement, I have computed impulse response functions on simulated data from a small structural macroeconomic model similar to the one used by Lindé (2001).[11] Figure 3 reports

[8] When measuring $\tilde{\pi}_t$ on Swedish data, I use the X12-method (multiplicative) to seasonally adjust the price level, and when computing $\bar{\pi}_t$ I use the fourth difference of the seasonally unadjusted price level.

[9] The U.S. dataset I am using in this paper is an updated version of the dataset used by Christiano, Eichenbaum and Evans (2001).

[10] The VAR-model for the U.S. contains the variables capacity utilization, inflation, consumption, output, investment, real wage rate, hours worked, fed funds rate and money growth. All variables are in logs except inflation, the Fed Funds rate and money growth. I use 2 lags and include a constant in the estimated VAR. The sample period is $1967Q3 - 1999Q4$. The interest rate shock size period 1 in Figure 2 is roughly .70.

[11] The model used to generate the data in Figure 3 includes the following equations; (i) a New-Keynesian Phillips curve $\pi_t = \omega_f \mathrm{E}_{t-1}\pi_{t+1} + (1-\omega_f)\pi_{t-1} + \gamma y_t + \varepsilon_{\pi,t}$, (ii) a New-Keynesian type of output equation $y_t = \beta_f \mathrm{E}_{t-1} y_{t+1} + \left(1 - \beta_f\right) y_{t-1} - \beta_r \mathrm{E}_{t-1}\left(R_t - \pi_{t+1}\right) + \varepsilon_{y,t}$, and (iii) a simple Taylor-type interest rate rule

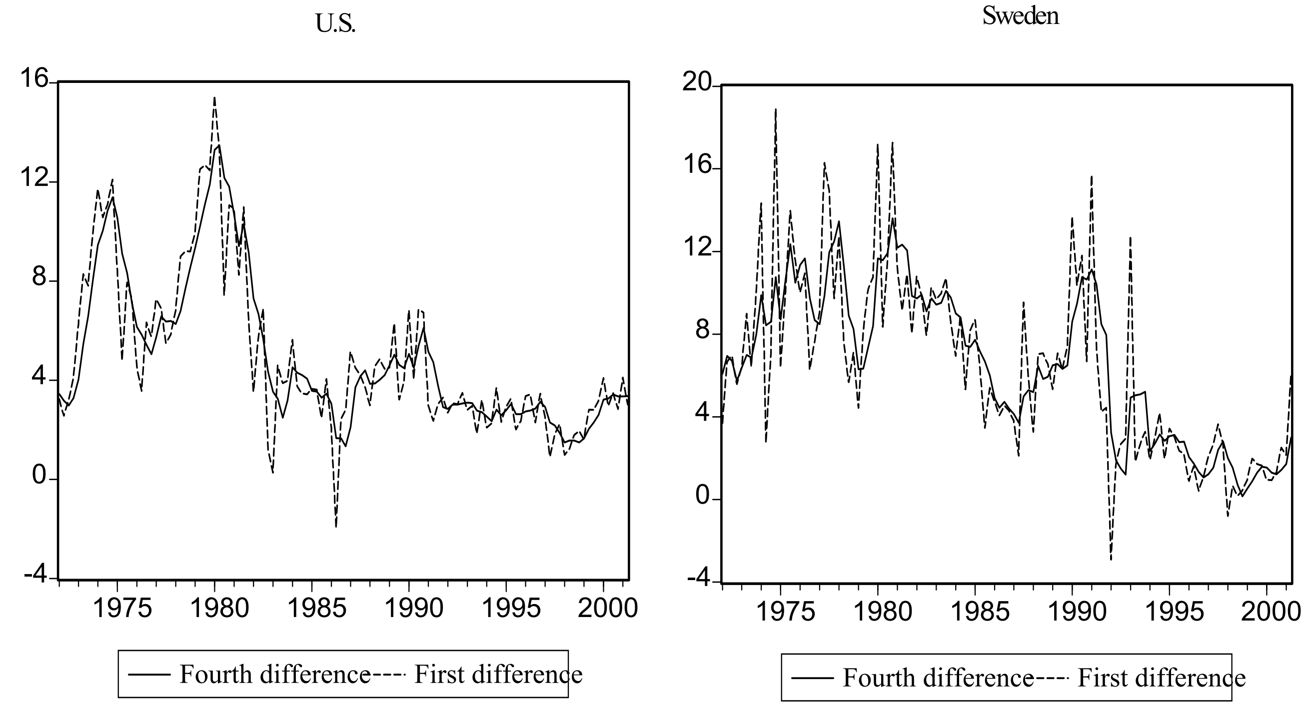

Figure 1: Annual (fourth difference) and annualized quarterly (first difference) CPI inflation rates in the U.S. and Sweden $1972Q1 - 2001Q2$.

the results of this exercise. The solid line denoted "True" denotes the true impulse response function to a policy shock in the theoretical model in which the "recursiveness assumption" is fulfilled since I use predetermined expectations in the output and inflation equations in the data generating process. The dashed lines "VAR" and "Moving average" show the impulse response functions in a VAR model with 2 lags (which is sufficient to describe the dynamics in the theoretical model) estimated on simulated data (sample size $20,000$ observations) where inflation is measured in first and fourth differences, respectively. As expected, we see that the "True" and "VAR" impulse response functions coincide whereas the impulse response function when inflation is measured with fourth differences (moving average) is smoothened out with the peak response delayed in time. The impulse response functions for output and the nominal interest rate are, however, very similar also for the moving average case. Since the features we

$R_t = (1-\rho)\left(\gamma_\pi \pi_t + \gamma_y y_t\right) + \rho R_{t-1} + \varepsilon_{R,t}$ where the shocks are assumed to follow AR(1)-processes. I used the following model parameters (for a motivation, see Lindé 2001) $\omega_f = .30$, $\gamma = .13$, $\beta_f = .50$, $\beta_r = .15$, $\rho = .50$, $\gamma_\pi = 1.5$ and $\gamma_y = .50$. The parameters in the AR(1)-processes for $\varepsilon_{\pi,t}$, $\varepsilon_{y,t}$ and $\varepsilon_{R,t}$ were set so that the model was able to (roughly) replicate the standard deviations and first-order autocorrelations of π, y and R on U.S. data $1970Q1 - 2001Q2$.

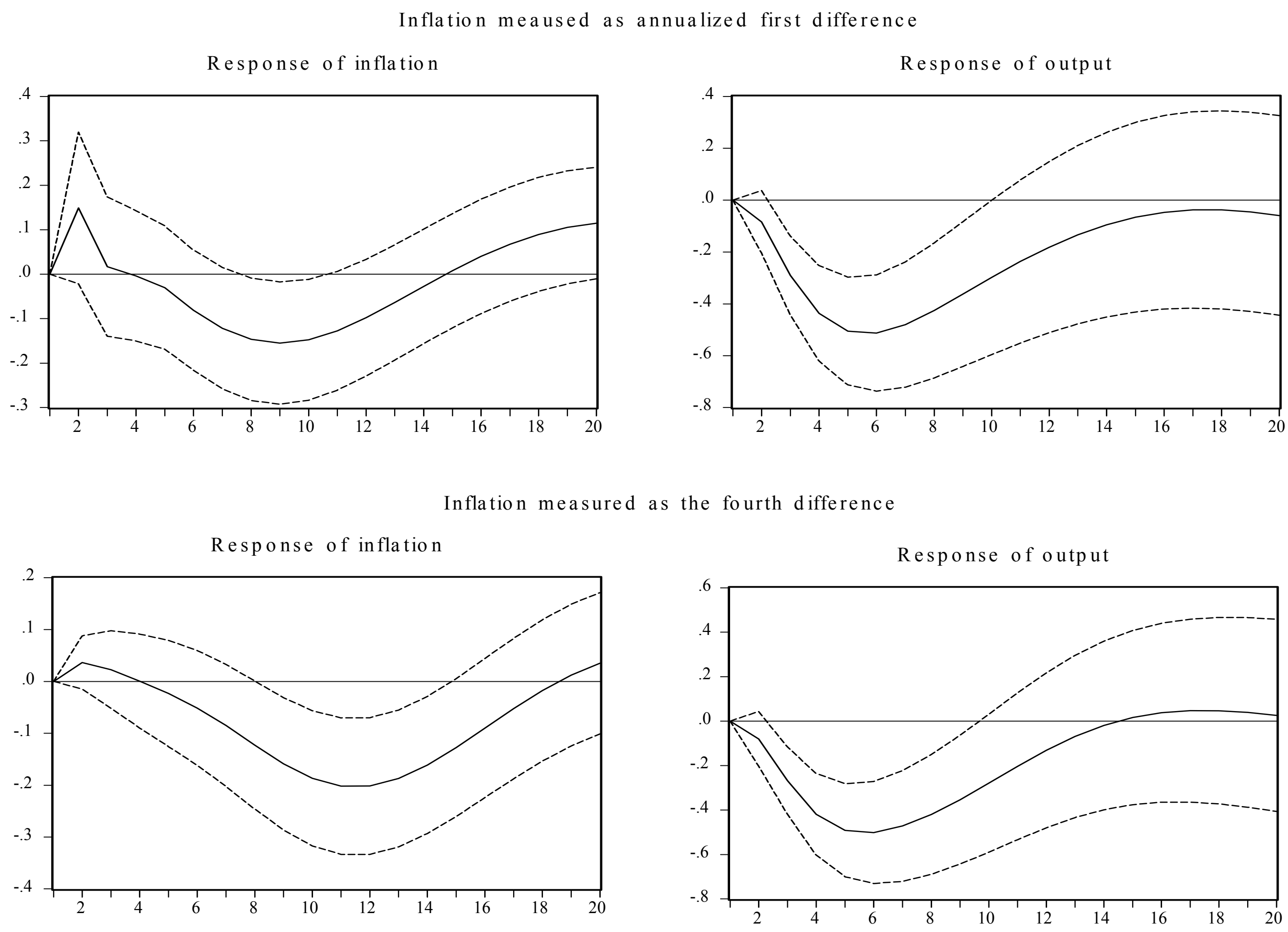

Figure 2: Impulse response functions on U.S. data to a monetary policy shock using different measures of inflation in the VAR.

observe on simulated data are essentially the same as we observe on U.S. data, we conclude that as long as inflation is positively autocorrelated with a fairly high autocorrelation coefficient, the use of $\bar{\pi}_t$ and $\tilde{\pi}_t$ should not make a big difference for the identification of monetary policy shocks, a finding verified by the results in Figure 2.

In Figure 4 we plot the impulse response functions for output and inflation on Swedish data for the preferred specification chosen in Section 4 using the GDP deflator as price index. As can be seen from the figure, choosing between the two measures of inflation in this case has very different consequences for the judgment of the effects of monetary policy on inflation and output, in contrast to the results reported on U.S. data in Figure 2. I would like to emphasize that the corresponding results when inflation is measured with the CPI are very similar to those

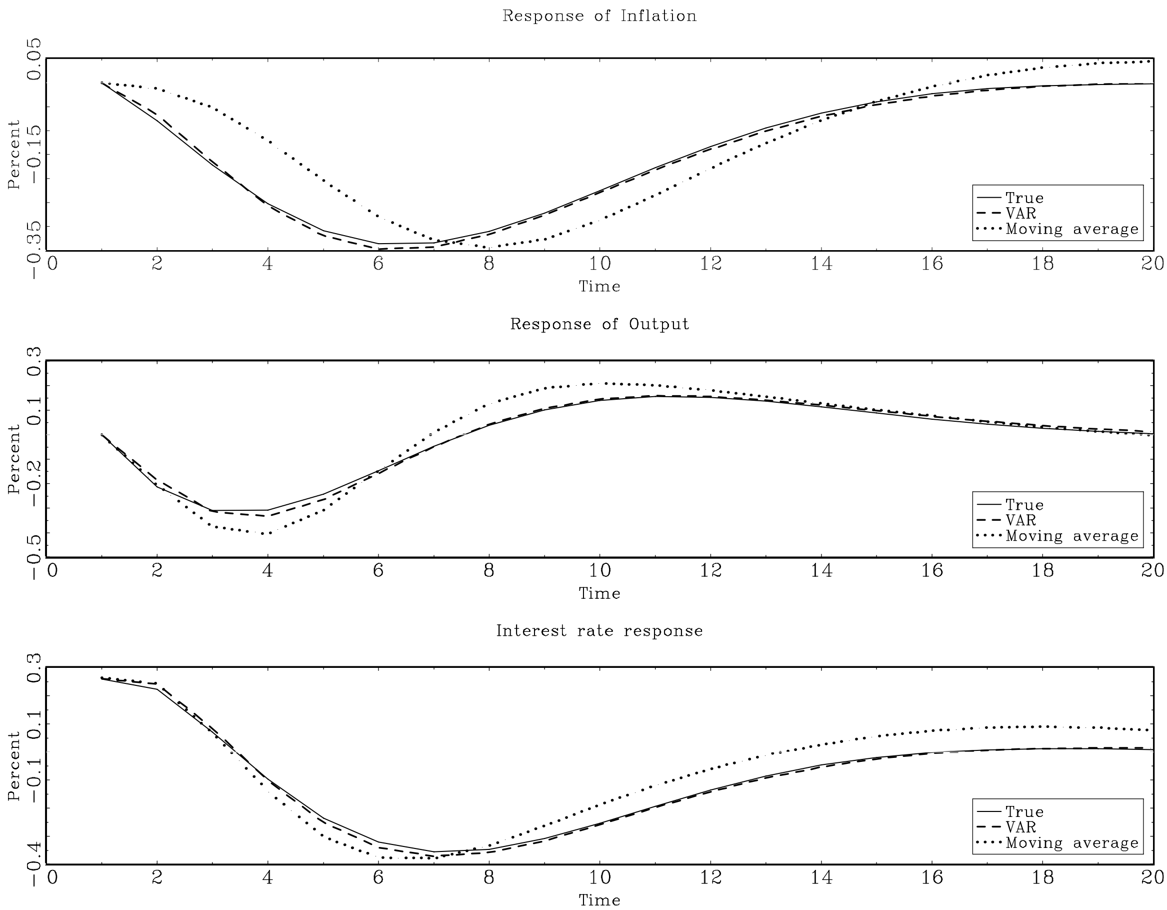

Figure 3: Impulse response functions based on simulations with a small structural macroeconomic model.

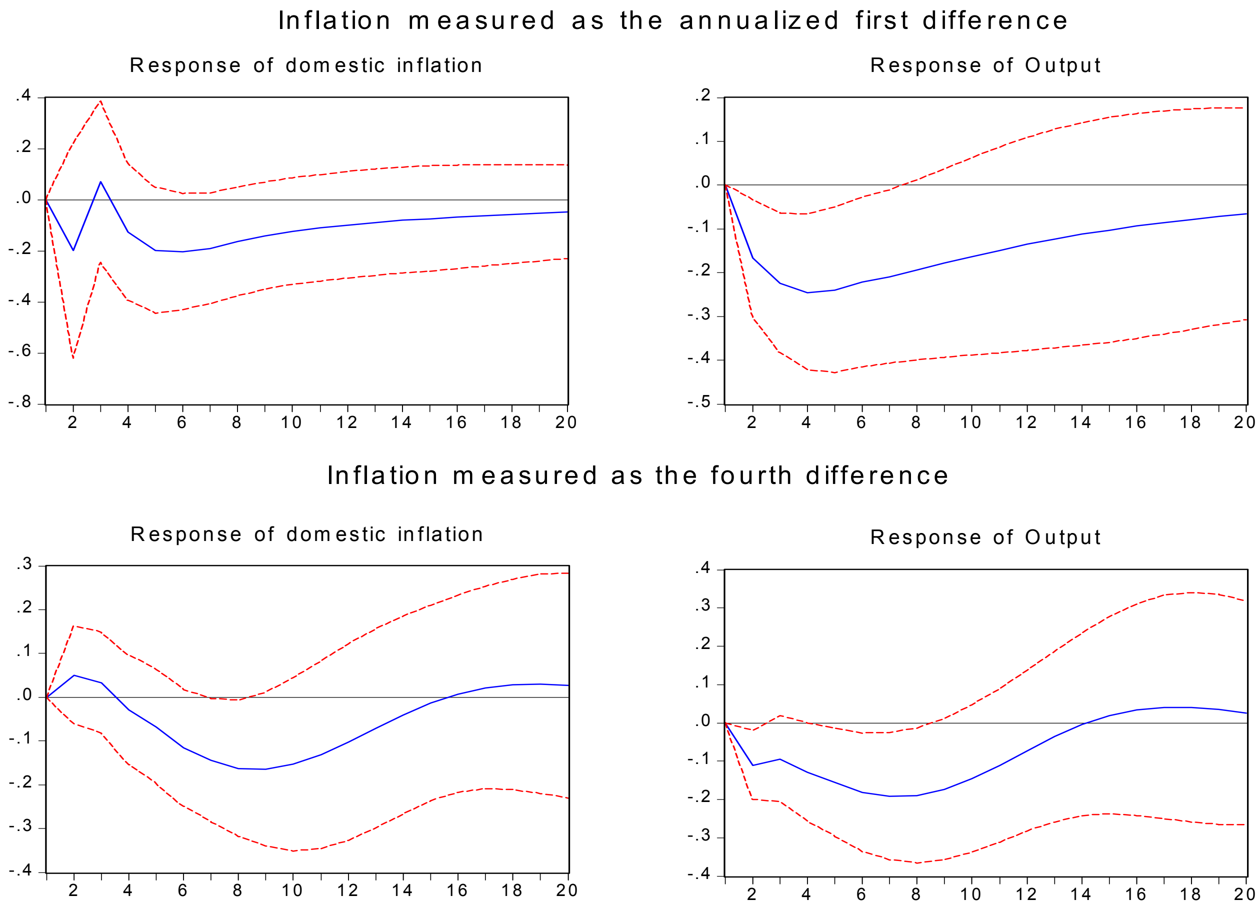

Figure 4: Impulse response functions on Swedish data to a monetary policy shock using different measures of inflation in the VAR.

shown in Figure 4.

The question then arises: why do these big differences in the impulse response functions occur in the VAR estimated on Swedish data? One potential explanation is omitted variables in the VAR estimated on Swedish data. To investigate if this is likely to be the case, I have performed a number of regressions of the difference between the annualized quarterly inflation and the annual inflation rate, using different sets of explanatory variables. The set of explanatory variables include the variables in the VAR, dummy variables (to investigate the possibility of some remaining seasonality effects in the data), changes in indirect taxes, and money growth (measured with M3). The regression results are reported in Table 1. In the table, we report regression results both for the CPI and the GDP deflator.

Table 1: Regression results for the difference between annualized quarterly and annual inflation 1976Q1-1999Q4.

Explanatory variables	Dependent variable $\tilde{\pi}_t - \bar{\pi}_t$ measured with GDP deflator		CPI	
y_t^*	0.26 (0.37)	−0.06 (0.09)	0.35 (0.25)	−0.07 (0.05)
$\tilde{\pi}_t^*$	0.12 (0.23)	0.11 (0.28)	0.19 (0.16)	0.38 (0.18)
R_t^*	−0.01 (0.25)	0.01 (0.25)	−0.03 (0.18)	−0.16 (0.17)
y_t	−0.05 (0.13)	0.03 (0.12)	0.01 (0.09)	0.12 (0.07)
$\bar{\pi}_t$	−0.02 (0.20)	−0.18 (0.21)	−0.08 (0.13)	−0.21 (0.13)
R_t	0.01 (0.13)	0.01 (0.13)	0.07 (0.09)	0.15 (0.08)
Q_t	0.07 (0.07)	0.03 (0.06)	0.08 (0.05)	0.05 (0.04)
$\tilde{\pi}_t^{imp}$	−0.03 (0.02)	−0.03 (0.02)	0.03 (0.02)	0.02 (0.02)
$D2$		−1.02 (1.25)		−0.51 (0.81)
$D3$		0.38 (1.13)		1.14 (0.73)
$D4$		−2.53 (0.95)		−0.79 (0.61)
$\Delta_1 T^{ind}$		1.22 (0.42)		1.24 (0.27)
$\Delta_1 \ln m_t$		−0.01 (0.04)		0.05 (0.03)
R^2	0.03	0.17	0.11	0.29
$\sqrt{\frac{SS}{T}}$	3.07	2.89	2.09	1.88

Notes: Standard errors in parentheses. A constant and linear time trend are included in the regressions but are not reported in the table. $\tilde{\pi}_t$ is defined as $\ln\left(\tilde{P}_t/\tilde{P}_{t-1}\right)$ where $\tilde{P}_t$ is the seasonally adjusted price index P_t (the X12-method, multiplicatively). $\bar{\pi}_t$ is defined as $\ln\left(P_t/P_{t-4}\right)$ where P_t is a seasonally unadjusted price index. $\Delta_1 T_t^{ind}$ is the quarterly change in indirect taxes (in percent) which have been subject to seasonal adjustment with the X12-method (additively). $\Delta_1 \ln m_t$ is the quarterly growth rate in the money stock (measured with M3) where m_t has been subject to seasonal adjustment with the X12-method (multiplicatively). $\sqrt{\frac{SS}{T}}$ is the standard error of the regression residuals.

As can be seen from the regression results in Table 1, the variables in the VAR model cannot explain much of the difference between the two series of inflation either when inflation is measured with the CPI or when it is measured with the GDP deflator (compare $\sqrt{\frac{SS}{T}}$ with the numbers discussed previously in text) . Therefore, when including $\tilde{\pi}_t$ instead of $\bar{\pi}_t$ in the VAR model, we introduce a lot of unexplained variation that presumably have large implications for the shape of the obtained estimated impulse response functions. So given these regression results, the results in Figure 4 are not so puzzling. When we extend the number of variables in the regressions in Table 1 to include changes in indirect taxes, money growth and seasonal dummies, we see that we can explain more of the variation in the dependent variable. In particular this is true when inflation is measured with the CPI. Among the new variables, changes in indirect taxes seem most important with a coefficient around unity which seems plausible.

Perhaps surprisingly, the regression results in Table 1 reveal tendencies of remaining seasonal variability in $\tilde{\pi}_t$, although the price indices have been seasonally adjusted with the X12-method. To shed more light on the potential problem with remaining seasonality in the $\tilde{\pi}_t$ series, I run the regression

$$\pi_t = C + D2 * T_{2,t} + D3 * T_{3,t} + D4 * T_{4,t} + u_t \tag{6}$$

for both the quarterly and annual inflation rates where C is a constant, and $T_{i,t}$ is a dummy variable with value 1 in quarter i and 0 otherwise. I estimate (6) recursively expanding the sample with 1 quarter from the smallest sample period $1971Q1 - 1975Q1$ until $2002Q4$. Before turning to the results of these regressions, let us discuss what we should expect as the outcome. First, since we do not detrend the inflation series prior to estimating (6), we expect that C declines over time (since inflation is falling over time in the sample, see Figure 1). Second, if the seasonal adjustment has been done correctly, we also expect that the estimates of $D2$, $D3$ and $D4$ are roughly zero at each point in time. A trend or persistent changes in the estimates would suggest that there are remaining seasonal patterns in the data. Looking at the results of this exercise in Figure 5, we see that the regression results for the fourth difference of inflation show little or no signs of remaining seasonality in the data; the constant term C changes over time in the expected direction and the coefficients $D2$, $D3$ and $D4$ are very close to 0 for each extended subsample. Turning to the results for the first difference instead, we see that the estimates of $D2$ and, although to a lesser extent, $D4$ are trending upwards when the sample period is extended, indicating that there are remaining seasonal components that are changing over time. The time-varying seasonal patterns for the first difference are visualized in Figure 6,

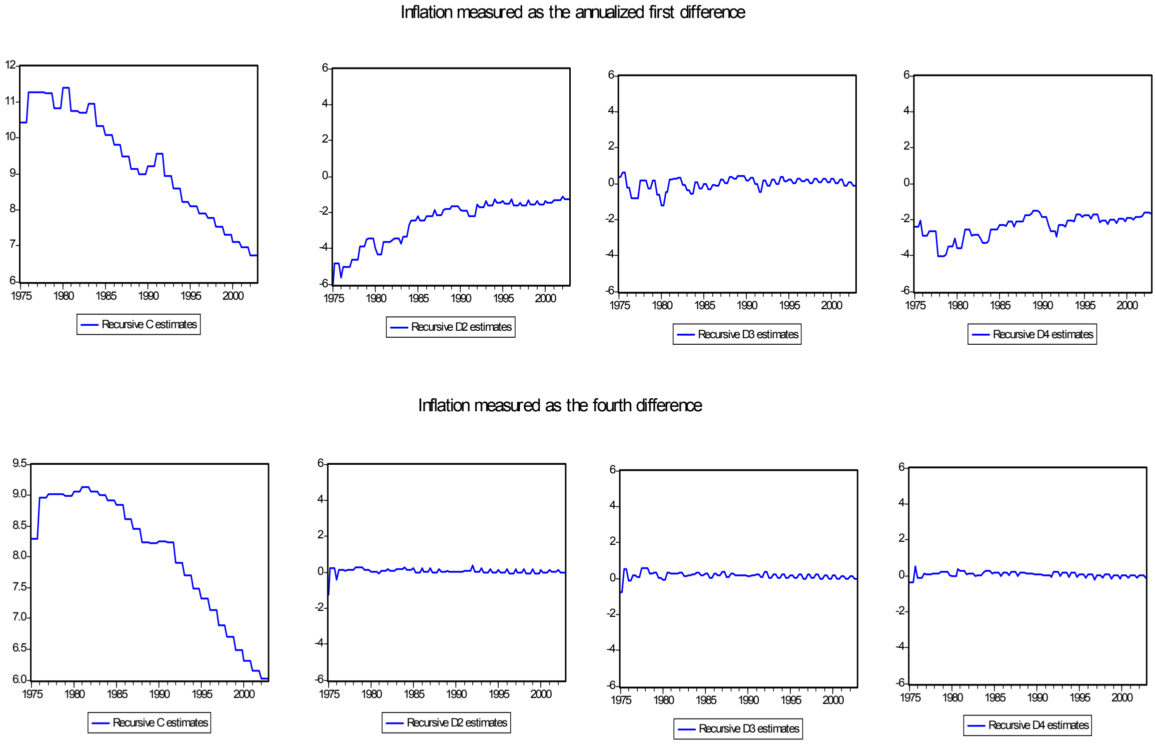

Figure 5: Recursive OLS regression results for quarterly (annualized) and annual inflation on a constant and seasonal dummy variables $1972Q1 - 2002Q4$.

which shows rolling window estimates of (6) starting $1972Q1$ with 8 quarters in each subsample. For $\bar{\pi}_t$, the rolling estimates for $D2$, $D3$ and $D4$ are nonsignificant noise around zero, whereas the estimates for $\tilde{\pi}_t$ are sometimes significantly different from zero and large in magnitude.

This makes the use of the first difference of the price level in the VAR undesirable. The results in Figures 5 and 6 are based on the GDP deflator, but the results are qualitatively similar - although somewhat less pronounced - when the CPI is used instead. The notion of time-varying seasonality effects is also consistent with VAR estimation results for the quarterly inflation rate, because R^2 (the share of explained variation in $\tilde{\pi}_t$) does not increase much when increasing the number of lags from 2 to 4 (only from .61 to .68), indicating that the different VAR results for $\tilde{\pi}_t$ and $\bar{\pi}_t$ are not due to including too few lags in the model for $\tilde{\pi}_t$.

The implication of the analysis in this subsection is that since it is not possible to explain the variation in the first difference in neither the CPI nor the GDP deflator with a broad set of variables that are included in our theoretical models of the economy, the excess variability in

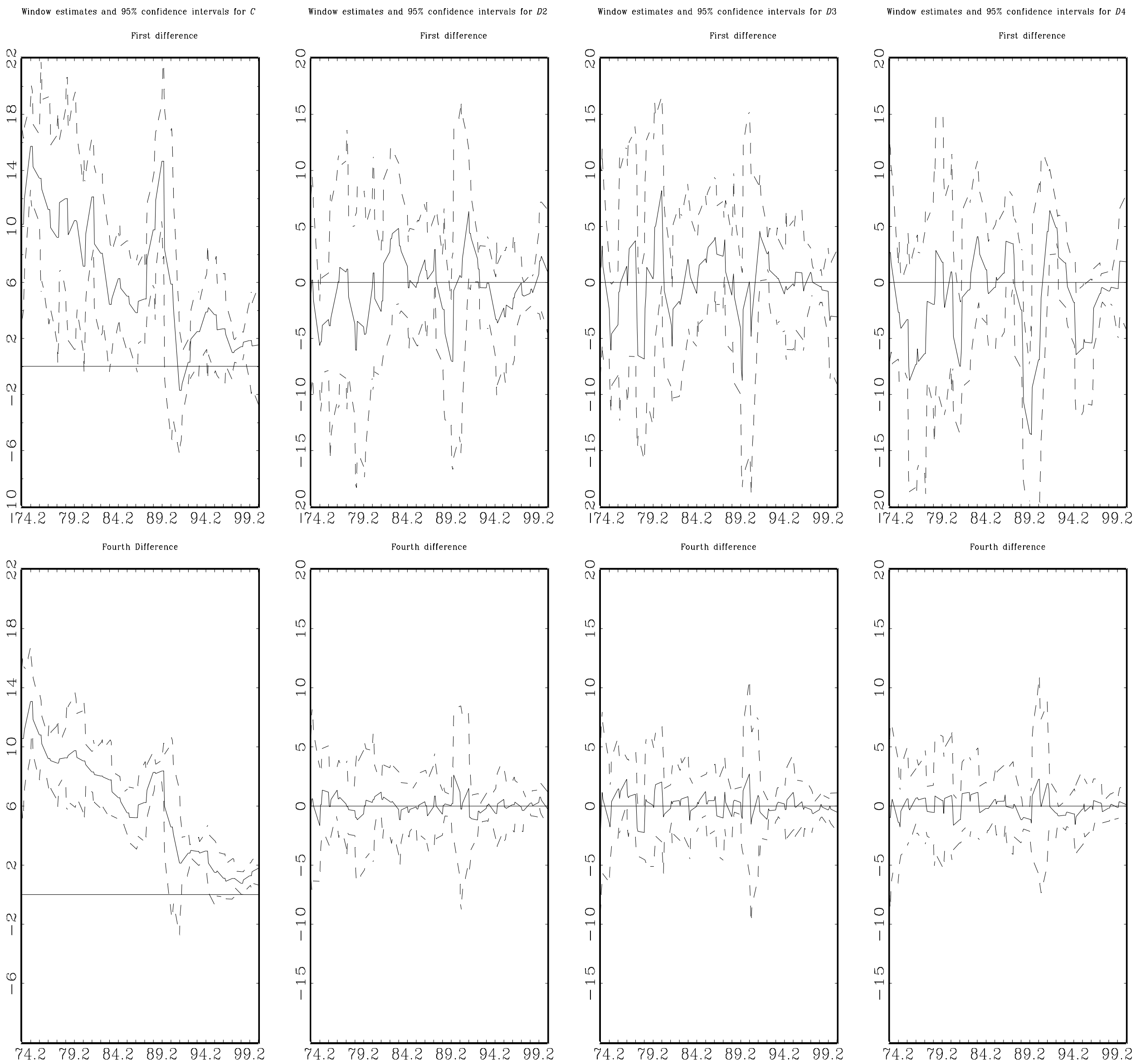

Figure 6: Rolling window estimates with 8 quarters in each subsample for quarterly and annual inflation rates $1972Q1 - 2002Q4$.

the annualized quarterly inflation rate compared to the annual inflation rate is likely to stem from non-economic factors, such as time-varying seasonal patterns. Therefore, and since I have demonstrated that the effects of falsely using the annual inflation rate rather than the quarterly inflation rate in the VAR are not large as long as inflation is positively autocorrelated (which is the case in the data), I will use the fourth difference of the price indices in the VAR analysis below rather than the first difference.

2.3 Lag selection

The specification test results for the VAR model are provided in Table 2. The results in the table refer to the VAR model in (1) when the foreign variables are included in X_t, see (3). The reason for this is that we will later use a separate VAR model for the foreign variables (under the supported assumption of block-exogeneity) in order to compute historical decompositions, and we therefore want to get an idea of how many lags we need to use in both VARs.

Table 2: Specification test results for the VAR model in (1)

lag length p	Test statistic & Information criteria					
	LR-test	Autocorrelation	Normality	Akaike	Hannan-Quinn	Schwartz
1	36.92 (0.99)	514.08 (0.05)	34.24 (0.04)	19.91	21.17	23.09
2	34.48 (0.99)	496.30 (0.06)	32.19 (0.04)	21.05	23.14	26.35
3	38.44 (0.99)	487.28 (0.39)	13.84 (0.46)	22.08	25.01	29.51
4	43.83 (0.97)	481.72 (0.99)	15.65 (0.44)	22.74	26.51	32.29

Notes: The results in the table refer to the VAR model where the foreign variables are endogenous, see (3). Approximate $p-$values for the Likelihood Ratio (LR-) test in parentheses (for testing the null hypothesis of lag length p vs. lag length $p+1$). The LR-test statistic is defined as $(T-k)\left[\ln\left(\det\left(\hat{\Sigma}_p\right)\right)-\ln\left(\det\left(\hat{\Sigma}_{p+1}\right)\right)\right]$ where T is the number of included observations (here, $T=66$), $k=4+np_1$ (where n is the number of variables in the system, $p_1=p+1$, and 4 is the number of exogenous variables in the VAR; a constant, linear trend and the two dummies). $\det\left(\hat{\Sigma}_p\right)$ is the determinant for the estimated covariance matrix in the VAR with p lags. The LR-statistic follows the χ^2 distribution with n^2 degrees of freedom (here, $n=8$ variables are included in the VAR). k is the small sample modification suggested by Sims (1980). Autokorrelation is the Portmanteau test for no autocorrelation up to 8 lags (bootstapped $p-$values in parentheses). Normality is the multivariate test for normality in the residuals (bootstapped $p-$values in parentheses). The sample period is $1986Q1-2002Q4$.

As can been seen from Table 2, the LR-test statistic is not very informative about the lag length selection in the VAR when allowing for a small sample modification. If anything, it seems as if 2 lags are preferred. The tests for autocorrelation and normality tests, on the other hand, suggest that at least 3 lags should be used. Turning to the information criteria, they all suggest that 1 lag is sufficient. Given the short sample (66 observations for $1986Q3-2002Q4$), this is perhaps not too surprising. By thinking of the VAR as an approximation of the solution to a theoretical open economy model, we note that a theoretical model with some inertia in the behavioral equations and possibly autocorrelated shocks implies that at least 2 lags should be used (see e.g. Lindé, Nessén and Söderström, 2003). As a compromise between theory and degrees of freedom in the VAR, I will therefore work with 2 lags as the benchmark specification, but do some sensitivity analysis w.r.t. the number of lags in the model.[12] It is interesting to note that it is possible to estimate what appears to be a stable VAR model, despite the change from fixed exchange rate regime to an inflation targeting regime during the sample period.

3 The dynamic effects of monetary policy shocks in the estimated VAR

The structure of this section is as follows. First, I will discuss the identifying assumption that is used to identify the dynamic effects of a shock to monetary policy in the estimated VAR. Second, I report the benchmark impulse response functions for a positive monetary policy shock (i.e. an exogenous increase in the nominal interest rate that is not part of the systematic response of policy to the state of the economy). Third, I conduct some sensitivity analysis along four dimensions; (i) by extending the sample backwards in time, (ii) by changing the number of lags in the VAR, (iii) by changing the order of the nominal interest rate R_t and recalculating the impulse response functions, and (iv) by including money in the VAR and examining if shocks to money seem important for the variation of the variables in the VAR.[13]

3.1 Identification of policy shocks

To identify the monetary policy shock, we will make use of the so-called "recursiveness assumption" that has become standard in the closed economy literature (see e.g. Christiano, Eichenbaum and Evans, 1999, 2001, and Angeloni, Kashyap, Mojon and Terlizzese, 2002), and assume that a shock to monetary policy has no contemporaneous effects on aggregate quantities and prices except for the nominal exchange rate (and thus the real exchange rate). By ordering Q_t after R_t in X_t, we assume that the Riksbank does not contemporaneously respond to shocks

[12] It should be noted that the specification test results for the model are qualitatively very similar if the foreign variables are included as exogenous variables instead of endogenous. The reason for this is that there appears to be very small spillover effects from the domestic variables to the foreign variables.

[13] Notice that by changing the ordering of the nominal interest rate in X_t we assume that the interest rate rule has the form $R_t = b_x X_{t-1} + \varepsilon_t$ instead of $R_t = b_{x,1} X_{1,t} + b_{x,2} X_{2,t-1} + \varepsilon_t$ where $X_{1,t}$ are the variables in X_t ordered before R_t and $X_{2,t}$ the variables ordered after.

to the nominal exchange rate, i.e. R_t is assumed to be contemporaneously unaffected by risk premium shocks.[14] Thus, we implicitly have in mind a policy rule of the form

$$R_t = \beta_1 (L) X_{1,t} + \beta_2 (L) X_{2,t-1} + \beta_3 (L) R_{t-1} + \varepsilon_t \tag{7}$$

where $X_{1,t}$ consists of y_t^*, π_t^*, R_t^*, y_t and π_t, and $X_{2,t}$ consists of Q_t and π_t^{imp}, and ε_t is the monetary policy shock which is orthogonal to the information set in (7). A shock to monetary policy is also assumed to affect import prices contemporaneously via the nominal exchange rate. This implies that domestic inflation (π_t) should be measured with the GDP deflator rather than the CPI. However, Smets and Wouters (2002) - who use the same identifying restrictions - use CPI as measure of π_t nevertheless, by assuming that the contemporaneous effect of a monetary policy shock through import prices on CPI-inflation is negligible due to imperfect pass-through. In this paper, π_t is measured with the GDP deflator, but the obtained results are very similar if the CPI is used instead as in Smets and Wouters (2002). Inflation on imported goods is measured at the producer price level, so I expect a greater coherence of inflation on imported goods with respect to the real exchange rate than if inflation on imported goods had been measured at the consumer level.

3.2 Benchmark impulse response functions

In this subsection we present the impulse response functions for the benchmark specification of the VAR model (2 lags, foreign variables exogenous, identification scheme given by 7, and sample period $1986Q1 - 2002Q4$).

The impulse responses in Figure 7 for output and inflation are similar to those estimated on U.S. data. Output and inflation display a "hump-shaped" pattern with peak effects occurring after $1.5 - 2$ years' time. The real exchange rate appreciates with a peak effect after 4 quarters. Although the interest rate response is not persistent, the effect on the real exchange rate are persistent and gradual. This is an intriguing result, which seems at odds with the idea of an uncovered interest parity condition commonly used in theoretical open economy models. The UIP condition suggests that the real exchange rate should appreciate contemporaneously as the policy shock occurs, and then gradually return to zero from below. Here, we instead obtain the maximum effect on the real exchange rate 4 quarters *after* the policy shock has occurred. The shape of the effects on inflation of imported goods are very similar to those for

[14] With this assumption, we avoid the identification problem of a shock to monetary policy emphasized by Giordani (2002).

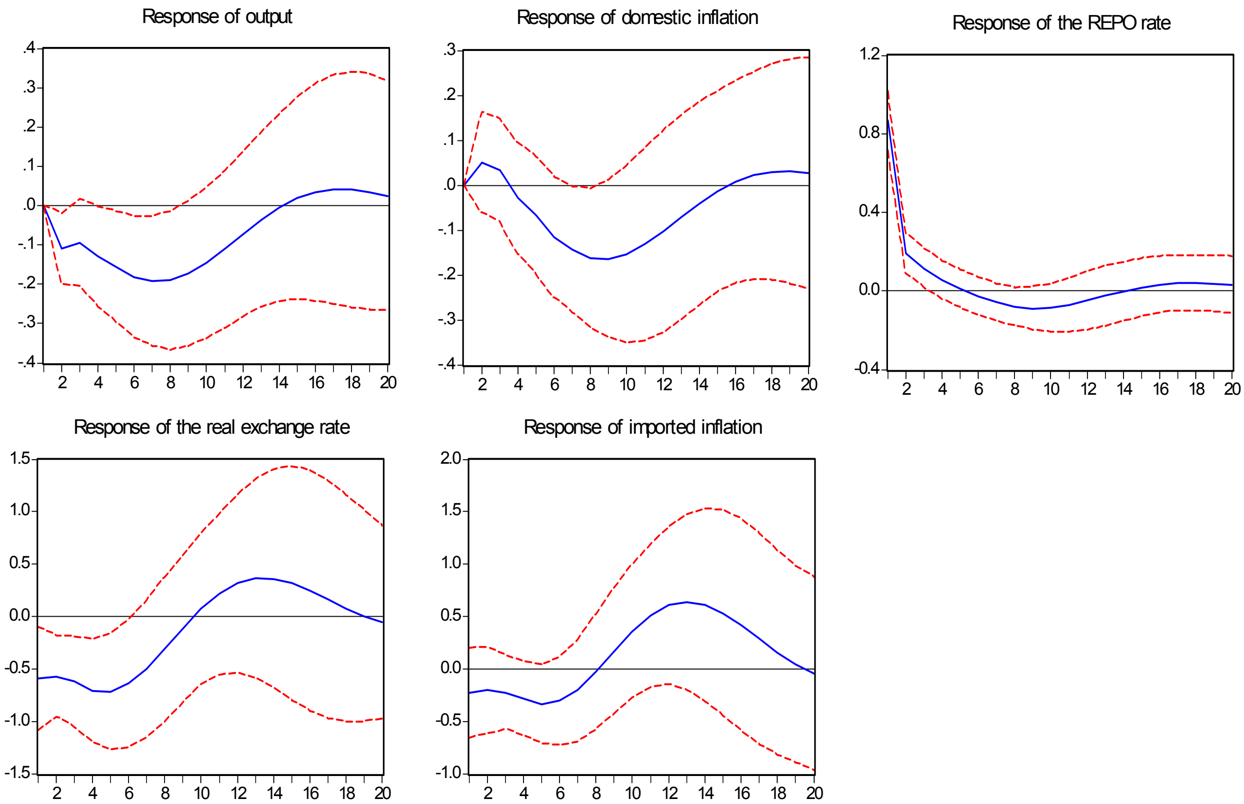

Figure 7: Impulse response functions for a shock to monetary policy in the baseline VAR.

the real exchange rate. At the same time, the point estimates suggest that pass-through in the first period is incomplete because the real exchange rate appreciates by more than the fall in imported inflation.[15]

As can clearly be seen in Figure 7, the impulse response functions are not significantly different from zero for many periods, although the asymptotic 95-percent confidence intervals are not symmetric around the zero line. Throughout the paper, I will use the terminology that

[15] I define pass-through as $\frac{d\pi_t^{imp}}{ds_t}$ where s_t is the nominal exchange rate. If pass-though is complete, then this ratio equals unity, i.e. prices are completely adjusted to account for the change in the nominal exchange rate. But here, due to the adopted identifying assumption that the domestic price level is not affected in the first period by a monetary policy shock along with the exogeneity assumption of the foreign price level, we have that $\frac{dq_1}{ds_1} = \frac{ds_1}{ds_1} = 1$, and it follows that $\frac{d\pi_1^{imp}}{ds_1} = \frac{d\pi_1^{imp}}{dq_1}$. And since this ratio is less than unity according to the estimated impulse response functions, we conclude that pass-through is incomplete in the first period. However, we cannot draw any conclusions about the following periods, because then the changes in the real and the nominal exchange rates do not coincide because the domestic price level changes. Finally, replacing the measure of inflation on imported goods at producer prices with consumer prices in the VAR, the initial drop is even lower and the shape of the impulse response function is very different than for the real exchange rate, suggesting that pass-through is far from complete at the consumer level.

the impulse response functions for a variable is small if the point estimates are close to zero and the associated confidence interval are roughly symmetrically distributed around 0.

Furthermore, we note in Figure 7 that the nominal interest rate relatively quickly goes to zero after the shock: after 4 quarters the point estimate of the impulse response function is only about 0.1 percent. For the other variables, the effects are much more persistent. This implies that in order for a theoretical model to be able to account for these impulse response functions, the model must embody internal propagation. The model by Christiano, Eichenbaum and Evans (2001) is an example of one such model for a closed economy. Finally, I would like to emphasize that the results in Figure 7 are essentially unaffected if the foreign variables instead are treated as endogenous in the VAR (see equation 3).

3.3 Sensitivity analysis

In this subsection, we will examine the robustness of the obtained impulse response functions in the previous subsection in a variety of dimensions.

3.3.1 Sample period

I have recalculated the impulse response functions in Figure 7 for the extended sample period $1976Q3 - 2002Q4$. The results are reported in Figure 8.[16]

Comparing Figures 7 and 8, we see that extending the sample period back to $1976Q3$ has quite large quantitative effects on the estimated impulse response functions, although some of the qualitative aspects of the results are similar; output falls gradually, we see a "price puzzle" in the first 4 quarters following the shock and thereafter a small decrease in inflation with a maximum drop after about 10 quarters. The real exchange rate appreciates, and the shape of the estimated impulse response function for imported inflation is similar to the one for the real exchange rate. But clearly, as might could have been anticipated due to the reasons discussed at the end of Section 2.1, the quantitative results for the impulse response functions in Figure 7 are not robust when extending the sample size, although some of the qualitative aspects are.

[16] With the VAR specification used in the paper, it is not possible to go further back in time (to $1970Q1$, say) because data on import prices (measured at the producer level) are only available from $1975Q1$ and onwards.

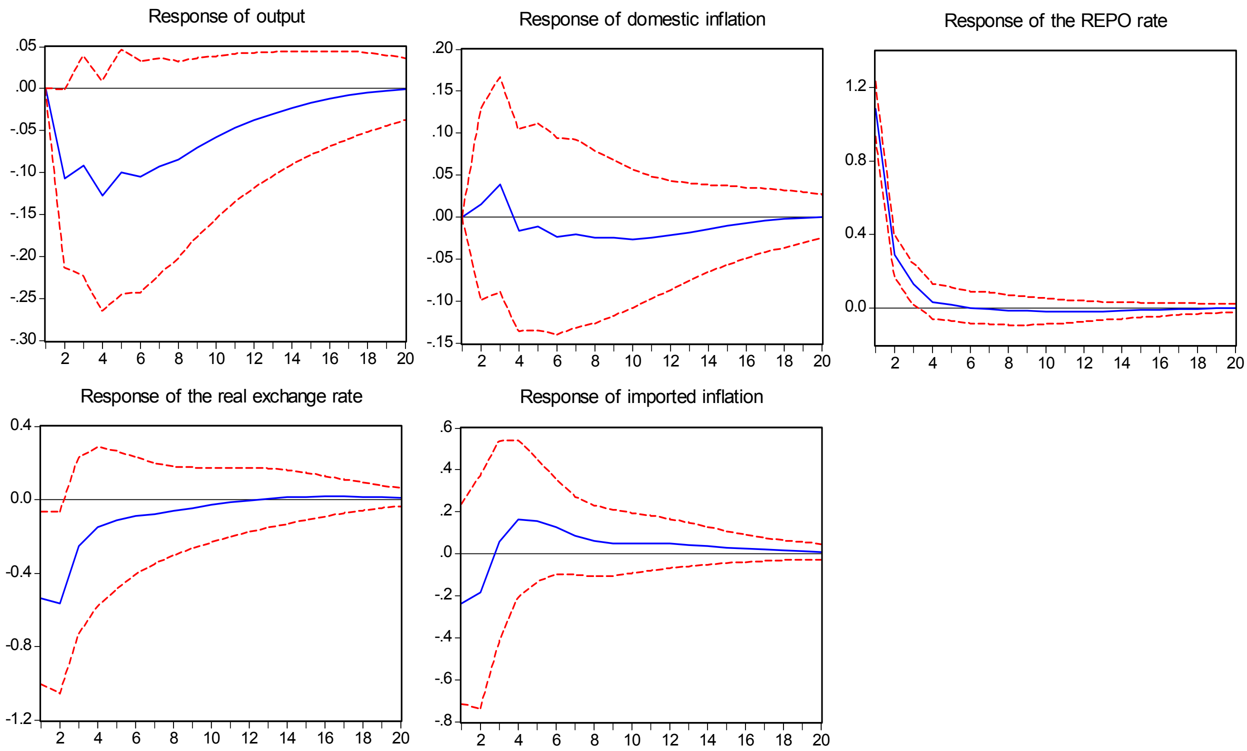

Figure 8: Impulse response functions with foreign variables exogenous in the VAR for the period $1975Q3 - 2002Q4$.

3.3.2 Number of lags in the VAR

Figure 9 shows the impulse response functions of output and inflation in the VAR model (using the same number of lags for the exogenous foreign variables as the endogenous variables, see equation 1) for 4 different lag lengths p; $p = 1$, $p = 2$, $p = 3$ and $p = 4$. It should be noted that for $p = 4$, the VAR is dynamically unstable and the largest root of the characteristic polynomial is above 1 in absolute modulus (1.01). As can clearly be seen from the figure, the choice of lag length does not have a great impact on the benchmark results in Figure 7. The impulse response functions still have roughly the same shape and peak effects around the same dates, with the possible exception for $p = 1$ when the peak effects appear to arise a couple of quarters earlier.

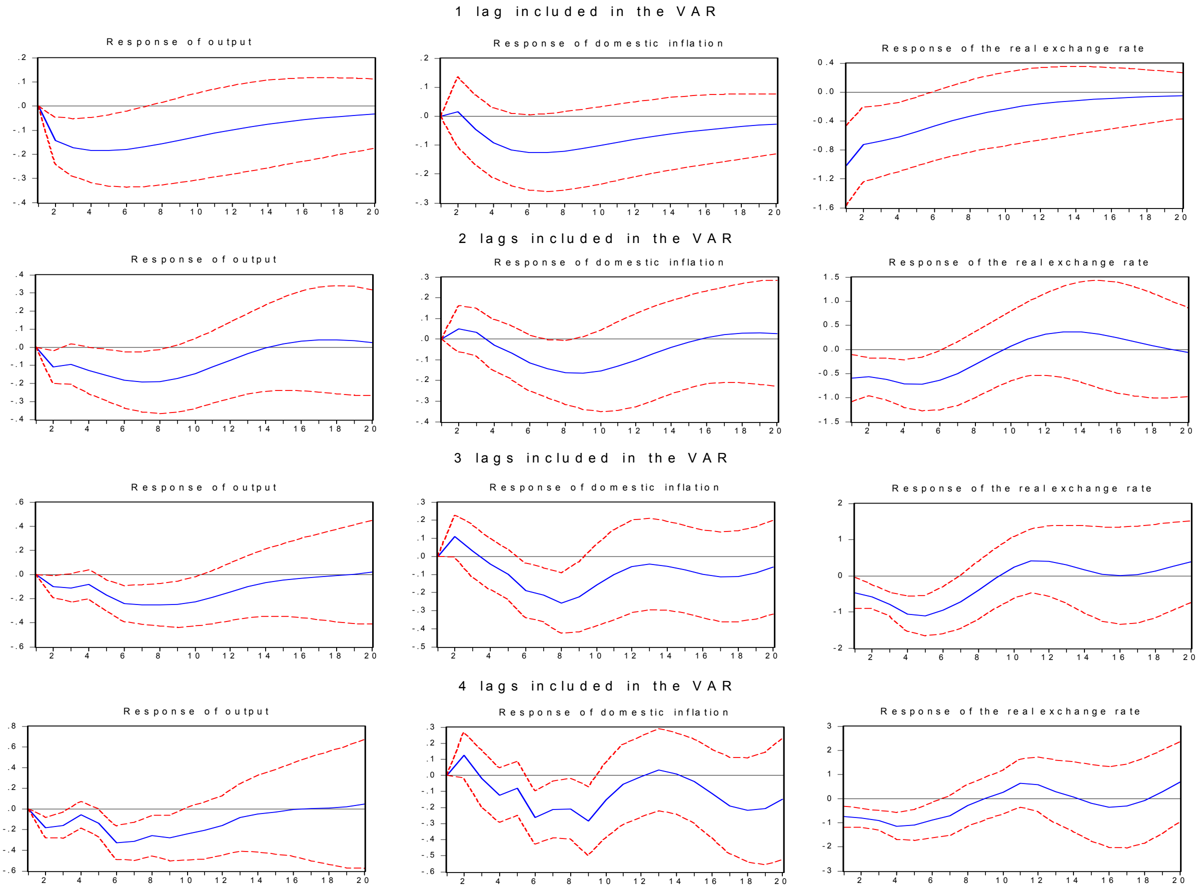

Figure 9: Impulse response functions to a policy shock using different number of lags in the VAR model.

3.3.3 Ordering of the interest rate

I have also changed the ordering of the interest rate in X_t (see 2) to

$$X_t = \left[\begin{array}{ccccc} R_t & y_t & \pi_t & Q_t & \pi_t^{imp} \end{array} \right]'$$

which means that the monetary authorities are assumed not to observe domestic output and inflation the same quarter they set the interest rate. This assumption seems somewhat more common on monthly data, see e.g. Sims (1992). For completeness, I also present results when the interest rate is ordered last in X_t, so that

$$X_t = \left[\begin{array}{ccccc} y_t & \pi_t & Q_t & \pi_t^{imp} & R_t \end{array} \right]'$$

thus assuming that the monetary authorities respond to all variables in period t when setting the nominal interest rate, but that policy has no effects on the other variables in X_t contemporaneously. The resulting impulse response functions for the other variables in X_t are plotted in Figure 10.

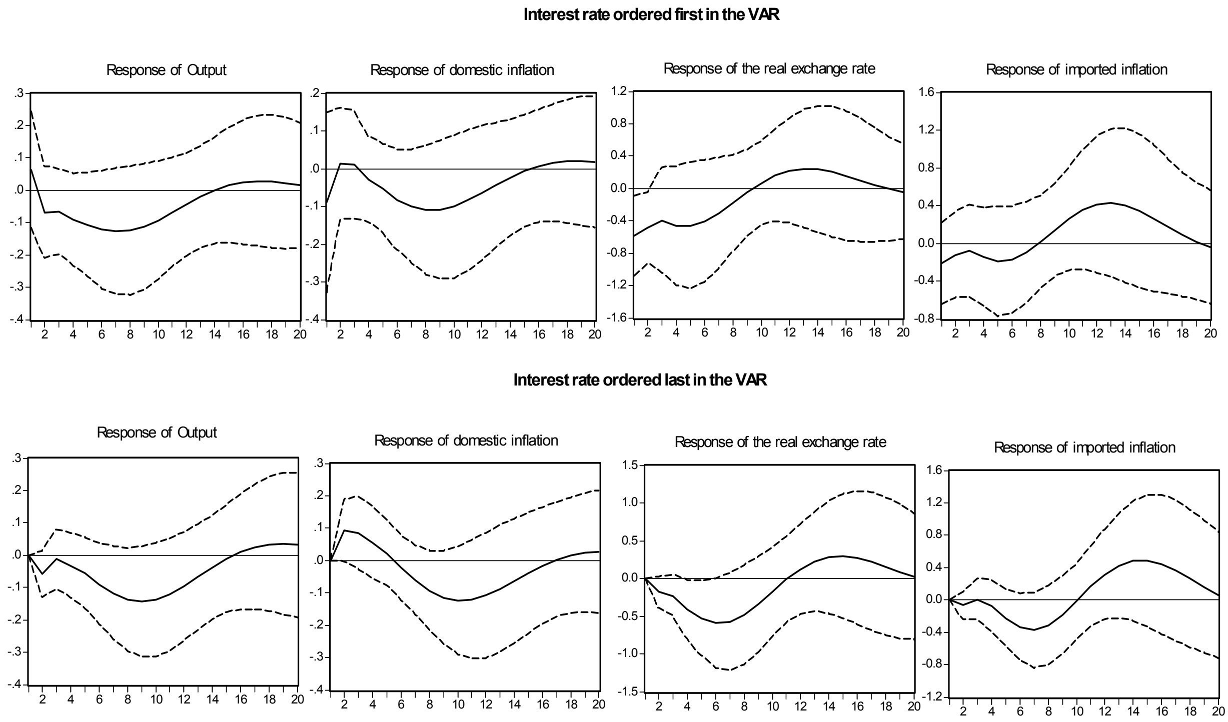

Figure 10: Impulse response functions to a policy shock when changing the ordering of variables in the VAR.

Comparing Figure 7 and 10, we see that the shape of the impulse response functions are very similar for all variables. The biggest difference is that the confidence intervals are somewhat wider than in the benchmark specification, so I can no longer exclude the possibility that the real effects of monetary policy on all variables except possibly the real exchange rate are zero or very small using these alternative identification schemes. So the identification scheme adopted in this paper is robust to the shape of the economic effects of monetary policy, but not to the statistical significance of the real effects of monetary policy. More technically, the result that the impulse response functions to a policy shock are robust w.r.t. the ordering of R_t implies that the off-diagonal elements in the A_0 matrix for the VAR model (1) written on structural form (dropping all exogenous variables),

$$A_0 X_t = A(L) X_{t-1} + \varepsilon_t, \tag{8}$$

are close to zero.[17] For the estimated VAR, I cannot reject the null hypothesis that all off-diagonal elements except the contemporaneous response of imported inflation to the real exchange rate are zero (i.e. allowing the A_0^{54}-element to be non-zero), the $\chi^2(9)$-statistic for the Likelihood Ratio test for over-identification being 10.26 with a p-value of 0.33.[18]

3.3.4 Including money in the VAR

There is no money variable included in the VAR model (1). The reasons for this are twofold. First, the Riksbank has during the estimation period pursued an interest rate policy, which should imply that money demand shocks have no effects on real quantities, interest rates, exchange rates and inflation, assuming that the long-term target growth of money supply has been determined once and for all by the Riksbank (i.e. an inflation target is the goal). Second, the theoretical framework underlying the specification of the VAR model assigns little or no role to monetary aggregates. If both these arguments are approximately correct, including money as the last variable in (2) and reestimating the VAR and computing the effects of a shock to the money variable should result in insignificant and small impulse response functions for the other variables in the VAR. In Figure 11, I plot the results of this exercise on Swedish data for the VAR specification where the foreign variables are exogenous and money is measured with $M0$.[19]

The results in Figure 11 suggest that the VAR model without a money variable included seems acceptable: the point estimates for the impulse response functions are small and the confidence intervals are in most cases symmetric around the zero line. This suggest that the Riksbank has almost completely insulated money demand shocks during this period. Shocks to the money variable do not account for much of the fluctuations in the other variables, with the possible exception of inflation on imported goods. A variance decomposition confirms that the monetary variable accounts for between 3 (output) and 5 percent (inflation on imported goods) of the variation in the other variables. These findings are in sharp contrast to the findings by Favara and Giordani (2002) for the U.S. economy. In their study, money shocks are quite important for aggregate quantities and prices, and the nominal interest rate.

[17] In (8), $\varepsilon_t \sim N(0, D)$ where D is a diagonal matrix.

[18] A test of A_0 being a diagonal matrix gives the $\chi^2(10)$-statistic 23.27 with a p-value of 0.01. Thus, I reject the null-hypothesis that the A_0^{54}-element is zero.

[19] I have also experimented with other monetary aggregates such as $M3$, but the results were essentially identical to those reported in Figure 11.

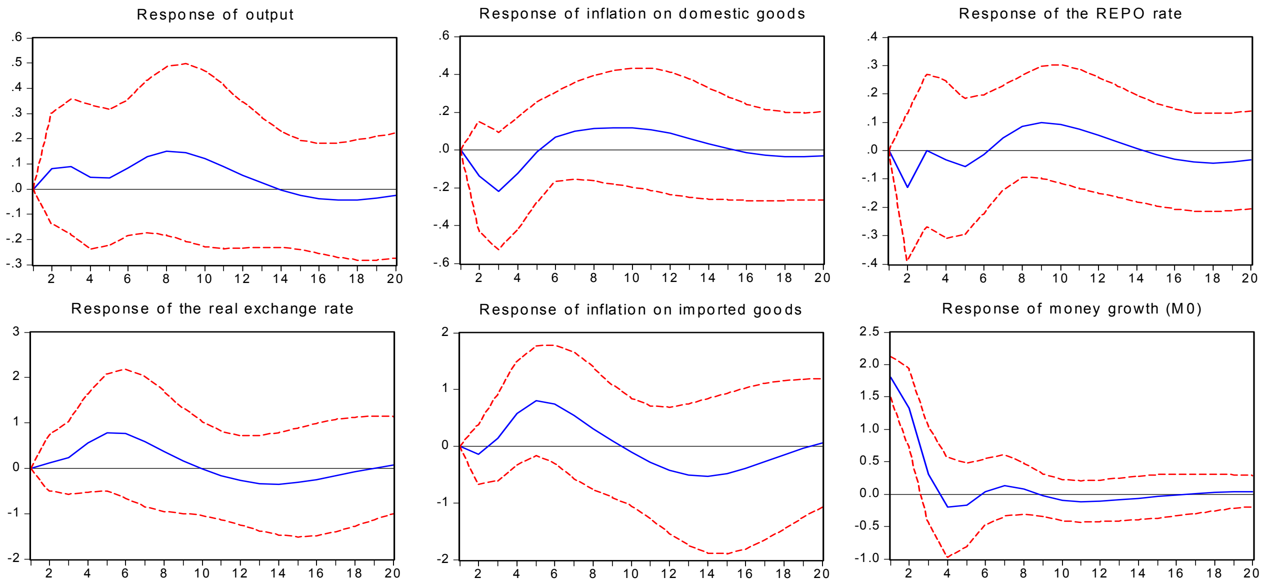

Figure 11: Impulse response functions to a shock in the growth rate of $M0$ when included as the last variable in the VAR model (assuming foreign variables are exogenous).

4 Sources behind macroeconomic fluctuations

In this section, I will use the estimated VAR to calculate the fraction of fluctuations around trend that are due to foreign and domestic shocks during the baseline sample period $1986Q1-2002Q4$. The reason why this is of great interest is that it will enable us to say something interesting about, for instance, to what extent the deep recession in the beginning of the 1990's in Sweden was homemade or caused by international factors.

Since block exogeneity for the foreign variables w.r.t. to the domestic variables holds in the benchmark VAR that we are working with, it is straightforward to investigate the role of foreign and domestic shocks behind macroeconomic fluctuations as long as we are only interested in the importance of *all* the foreign/domestic shocks, since we then do not need to identify each foreign and domestic shock separately. Among the domestic shocks we will, however, specifically look at how the identified monetary policy shocks have contributed to business cycles.

With macroeconomic fluctuations, we mean deviations around trend in the estimated VAR. The trend path for the variables in X, denoted $\bar{X}_t$, can be computed with a dynamic simulation of

$$\bar{X}_t = C_X + \delta_1 D_{923} + \delta_2 D_{931013} + \tau_X T_t + \sum_{i=0}^{2} \Upsilon_i \bar{Z}_{t-i} + \sum_{i=1}^{2} \Gamma_i \bar{X}_{t-i} \tag{9}$$

where $\bar{Z}_t$ is the trend path for the foreign variables which in turn is given by a dynamic simulation of the following estimated VAR(2) model

$$\bar{Z}_t = C_Z + \tau_Z T_t + \sum_{i=1}^{2} \Psi_i \bar{Z}_{t-i}.$$

As starting values when generating the trend, we use the actual values $1986Q1$ and $1986Q2$ for X_t and Z_t. Deviations around trend, denoted $\hat{X}_t$, are defined as $\hat{X}_t = X_t - \bar{X}_t$. Fluctuations due to foreign shocks, denoted $\hat{X}_t^f$, are then given by a dynamic simulation of (9) with $\bar{Z}_t$ replaced by Z_t (the actuals) minus $\bar{X}_t$, whereas fluctuations due to domestic shocks, denoted $\hat{X}_t^d$, are given by a dynamic simulation of (9) with the estimated one-step ahead forecast errors (i.e. the shocks) included minus $\bar{X}_t$. Note that $\hat{X}_t = \hat{X}_t^d + \hat{X}_t^f$. Among the domestic shocks, we can use our identifying assumption for the monetary policy shock to decompose $\hat{X}_t^d$ into two parts, fluctuations that are due to policy shocks ($\hat{X}_t^{pol}$), and fluctuations that are due to not identified domestic shocks ($\hat{X}_t^{nid}$).

Figure 12 shows the deviations around trend, $\hat{X}_t$, in the VAR along with the fluctuations due to foreign shocks, $\hat{X}_t^f$. There are a couple of interesting observations worth discussing. First, we notice that the VAR captures the idea of a boom in the late 1980's in Sweden, followed by the deep recession in the beginning of the 1990's with a trough in the last/first quarters 1992/1993. The huge and rapid depreciation of real exchange rate due to the introduction of the floating nominal exchange rate restored the competitiveness of the Swedish exporting firms, and got the economy "jump-started" at the end/beginning of 1993/1994. Interestingly, there is a clear upward trend in the "trend" real exchange rate (not shown in Figure 12) after the introduction of the floating exchange rate in the VAR, implying a continued deprecation of the real exchange rate after the adoption of the inflation targeting regime in Sweden. As could have been expected, movements in imported inflation at the producer price level is highly positively correlated with movements in the real exchange rate during this period. The deep recession was also associated with a dramatic decrease in domestic inflation. Finally, we see that the current downturn in economic activity in Sweden is mainly driven by foreign shocks, i.e. an international recession spilling over to the Swedish economy.

From Figure 12, it is evident that foreign shocks have contributed substantially to Swedish business cycles during this period. In particular, they seem to account for a substantial part of

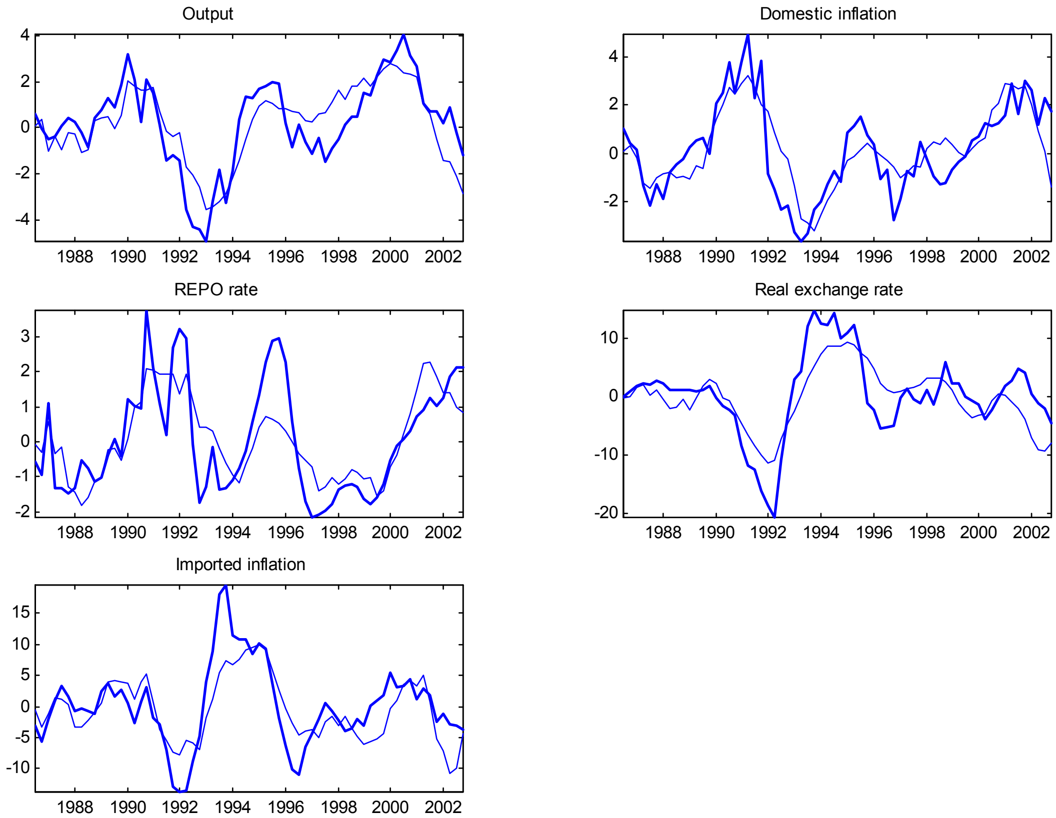

Figure 12: Deviations around trend (thick line) and fluctuations due to foreign shocks (thin line) in the estimated VAR.

the variation of output and domestic inflation. Quite naturally, they also seem more important for the fluctuations in the lower frequencies than the higher frequencies. Figure 13 shows deviations around trend along with fluctuations due to all domestic shocks. From the Figure, we notice that the domestic shocks account for more of the higher frequency movements. A variance decomposition shows that foreign shocks account for 54 and 52 percent in the fluctuations around trend for output and domestic inflation, whereas they account for 49 and 45 percent of the fluctuations the Hodrick-Prescott filtered $\hat{X}_t$, demonstrating the relative importance of domestic shocks for high frequency movements. Comparing Figures 12 and 13, we see that foreign shocks were contributing substantially to the deep recession in the beginning of the 1990's, so the VAR clearly suggest that this was not an entirely domestic story. Moreover, according to the VAR model, the boom in the late 1990's was to a large extent driven by an international boom as well.

An interesting feature according to Figure 13 is that the persistent and sharp increase in the nominal interest rate - the REPO rate - during 1995 and the beginning of 1996 was due to domestic shocks. Berg, Jansson and Vredin (2002) argue, by estimating various policy rules based on a unique Riksbank real-time dataset consisting of forecasts on inflation and the output-gap, that the high REPO rates recorded during this period are best explained by a sequence of positive policy shocks. That is, the Riksbank pursued a tightening of monetary policy in order to establish credibility for the inflation target that was introduced in 1995. By applying the identification scheme discussed in Section 3.1, we can compute the fluctuations in Figure 13 that are caused by the identified policy shocks. In Figure 14, I report the fluctuations due to policy shocks.

Interestingly, we see from Figure 14 that the estimated VAR supports the arguments by Berg, Jansson and Vredin (2002) using a completely different empirical approach and different dataset. According to the estimated VAR, the tightening of monetary policy during this period were mainly driven by policy shocks, and the persistent increase in the REPO rate during this period contributed to a clear drop in economic activity and domestic inflation and an appreciation of the real exchange rate in the subsequent period 1996-1998. For the other parts of the sample period, we see that policy shocks do not contribute much at all to macroeconomic fluctuations, a finding that are in line with results for the U.S. economy (see e.g. Altig et al., 2002).

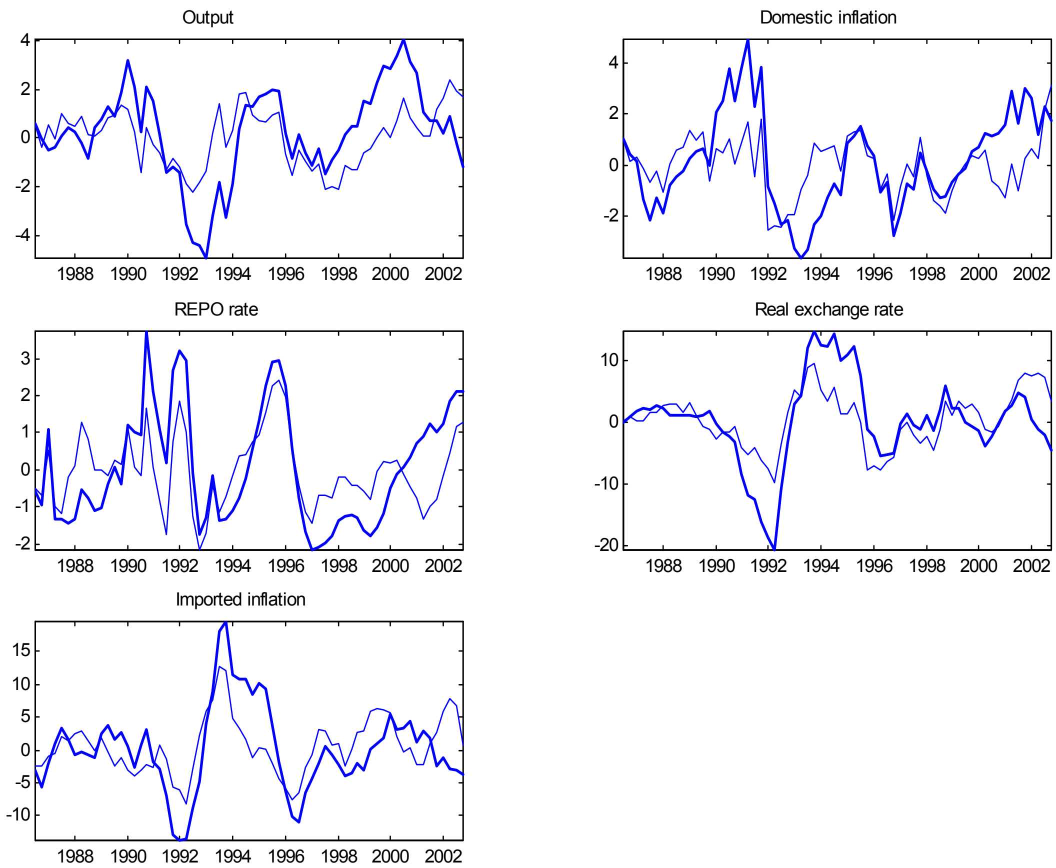

Figure 13: Deviations around trend (thick line) and fluctuations due to domestic shocks (thin line) in the estimated VAR.

5 Concluding remarks

The findings of the paper suggest that time-varying seasonal patterns and problems in fitting a stable VAR to a period spanning a transition from a highly regulated to a deregulated credit market create difficulties to obtain "reasonable" effects of monetary policy shocks on Swedish data using standard techniques. When limiting the sample to the period after most of the important financial deregulations had been implemented, and using annual inflation rates as a way to deseasonalize inflation, we obtain dynamic effects of monetary policy shocks that are very similar to those reported elsewhere in the literature, e.g. by Christiano, Eichenbaum and Evans (2001). In addition to obtaining hump-shaped and gradual real effects on output and domestic inflation to a policy shock, I also document that the real exchange rate response is very inertial.

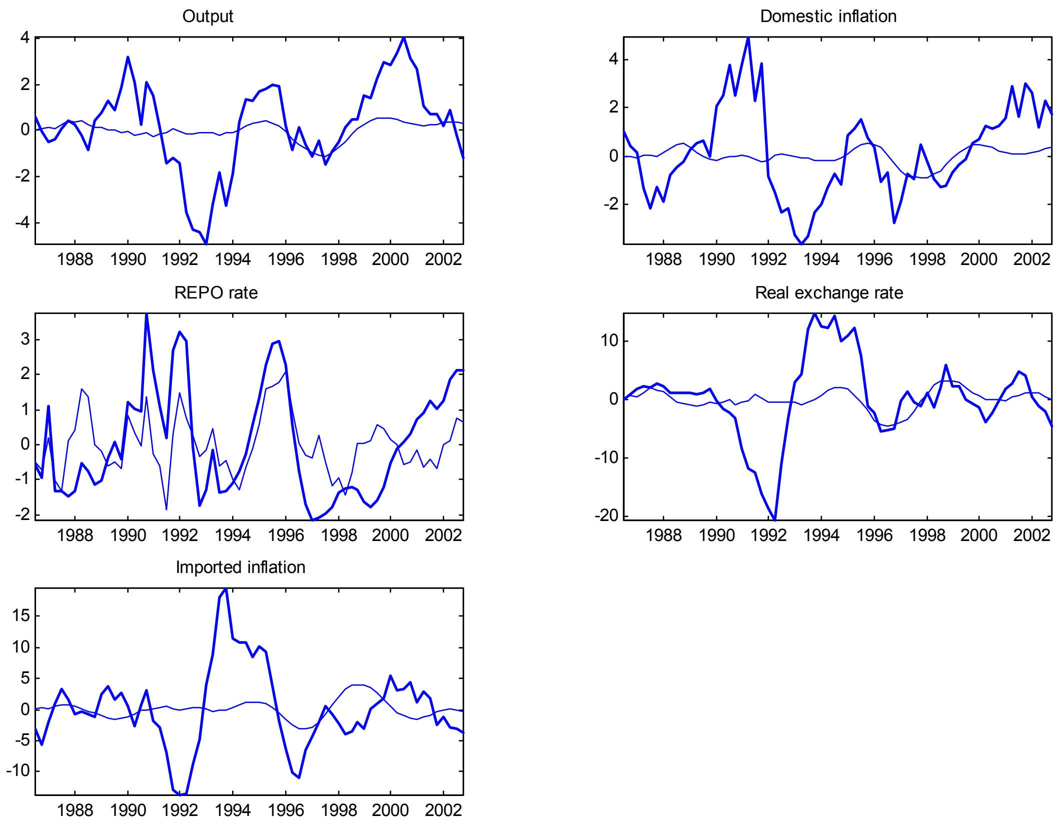

Figure 14: Deviations around trend (thick line) and fluctuations due to policy shocks (thin line) in the estimated VAR.

Any theoretical model of an open economy with flexible prices and a pure forward-looking uncovered interest parity condition included will have a hard time replicating this response. So in a sense, this finding complements the huge empirical literature which suggest that the uncovered interest parity condition does not hold in the data, see e.g. Lewis (1995). In essence, these findings suggest that in order for a theoretical open economy model to be able to account for the impulse responses of a shock to monetary policy, it must embody internal propagation, see Lindé, Nessén, and Söderström (2003).

I also performed an extensive sensitivity analysis of the results, and that the results are robust to various perturbations in ways that can be expected. For instance, the results are not sensitive w.r.t. the number of lags, the adopted identification scheme, or the inclusion of money growth in the VAR model, but as expected they are sensitive to the choice of sample period.

Since I found strong support in the VAR for the conventional small open economy assumption that the domestic economy is highly dependent of the international development but not the converse (i.e. the foreign variables are not influenced by the domestic variables), I was able to examine the relative importance of domestic and foreign shocks for Swedish business cycles during the period $1986 - 2002$. I found that foreign shocks account for slightly more than 50 percent of the macroeconomic fluctuations in Sweden, and that they were an important determinant of the deep recession in the beginning of the 1990's. Interestingly, the VAR model suggest that the international recession is the major driving source behind the current drop in economic activity in Sweden. I also documented that significant and persistent movements in the short-term interest rate have important real effects on the economy, suggesting that the conduct of monetary policy is important for the macroeconomic outcome. The results suggest a larger role for foreign shocks than reported by Lindé (1998) for the period $1950 - 95$. Most likely, the reason why they differ is that the results in Lindé (1998) are influenced by the huge government expansion during the period $1950 - 80$ in Sweden, which implied a larger role for domestic shocks during this period.

One issue that remains to be investigated further is to what extent measuring the inflation rate with the fourth difference of the seasonally unadjusted price level eliminates the seasonality effects in the data. Another question is whether the findings of time-varying seasonal patterns carry over to other small open economies, or is just a feature of Swedish data. For instance, we know that inflation for all EU countries are more volatile than inflation for the EU area as a whole, suggesting the existence of idiosyncratic shocks hitting each country that wash out in the aggregate. One interpretation is that part of these idiosyncratic shocks are time-varying seasonal patterns, but this a subject for further research. A possibility here would be to apply the Canova and Hansen (1995) test of changing seasonal patterns for inflation measured with the first difference in the EU countries separately and the EU aggregate.

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ISSN 1402-9103 Riksbank Printing Office Stockholm 2003