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Dynamics of Inflation Responses to Monetary Policy in the EMU Area

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Abstract:

This paper analyses the effects of the European Central Bank's monetary policy on EMU member countries' inflation in a Bayesian structural vector autoregressive framework. The choice of variables capturing monetary conditions in the EMU area is guided by a simple closed economy New-Keynesian macro model in which the interest rate is the channel for monetary policy transmission. Drawn impulse responses suggest that inflation responses to common, expansionary monetary policy shock can be seen to be asymmetric in the EMU area.

JEL Classification: C11, E52

Keywords: European Central Bank; monetary policy; asymmetry; Bayesian structural vector autoregressive model; posterior model probabilities

1. Introduction

At the beginning of 1999 11 European countries¹ were shifted into a monetary system with a common monetary policy. Before the European Monetary Union (EMU) era central banks in member countries were able to conduct independent monetary policies. The policy operations could be suited solely on the basis of domestic economy conditions – an independent central bank could for instance stimulate the domestic economy if deemed necessary. The independency of domestic central banks can be seen to be lost when a country joined the EMU, and ever since monetary policy decisions have been made exclusively by the European Central Bank (ECB). An evident practical problem with this common monetary policy area is written in its history. Past economic conditions in EMU member countries imply that economic conditions have been and are heterogeneous *per se*, which means that common monetary policy actions will most likely cause asymmetric effects in member countries. Due to this, the ECB should find itself confronted with challenges in tuning and conducting monetary policy. It may well be the diversity of economic and institutional structures across the member countries which constitute the reason why common monetary policy shocks have impacts of different magnitudes in the economies in the EMU area, especially in inflation. The essence of this is manifested in the annual inflation figures of the various EMU member countries, where in only few cases inflation series have converged to the 2 per cent inflation target while the aggregate inflation has varied fairly closely around the target in the EMU era. We agree that the ECB's monetary policy can have a stabilizing role and might be optimal at aggregate level, but monetary policy effects in individual member countries can be crucially asymmetric.

The successful conduct of monetary policy in the EMU area requires a knowledge of how rapidly innovations in monetary policy are absorbed in member countries and how great monetary policy effects actually are. Then, for instance, it would be of interest to see how consumer price inflation in EMU member country responds to a common monetary policy shock in relation to EMU aggregate consumer price inflation. The monetary

¹ Greece joined the group two years later in year 2001.

response dynamics of consumer price inflation in EMU member countries is important in that the ECB declares that the (EMU area-wide) consumer price inflation plays a major role in the ECB's monetary policy strategy, and generally, relative price inflation among the EMU member countries must also be seen to be important for welfare reasons.

The literature² provides a plethora of studies concerned to depict monetary conditions and monetary policy effects in the EMU area. Unfortunately, however, there would seem to prevail no solid consensus as to a model specification which fits the EMU area. Recently the literature has focused on structural macroeconomic models (New Keynesian models) specified for both forecasting and policy analysis purposes; see Sungbae and Schorfheide (2007) for a survey. The drawback in these studies is that the models constructed are typically complex and carry an increased model uncertainty; see for instance Smets and Wouters (2003, 2005 and 2007). Hence, a descriptive statistical modeling approach would seem preferable to enable us to better understand the dynamics of conditions affecting the stance of monetary policy in the EMU.

In this paper we capture the monetary policy effects in price inflation dynamics with a statistical model which both allows analysis of the dynamic responses of model variables and is sufficiently flexible in setting ex-ante restrictions on the contemporaneous effects of variables specified in a model. We see the structural vector autoregressive (SVAR) models to be best suited for our purposes, since we agree with Peersman (2004) that to make valid cross-country comparisons we need to construct a model wherein all member countries are exposed to the same monetary policy shock. Moreover, SVAR models are

² For detailed surveys see for example Mojon and Peersman (2001) and Peersman (2004). Angeloni and Ehrmann (2004) use quarterly EMU panel data over the period 1998:1-2003:2 to track down the sources of the inflation differences among the EMU member countries. They employ an open economy version model letting the real exchange variable exist in the model. They find that the magnitude of inflation persistence is the driving force generating the inflation divergence. Batini (2006) and Batini et al. (2001) list and discuss three possible types of inflation persistence; 1) positive serial correlation in inflation, 2) lag between system monetary action and its effect of inflation and 3) lagged responses of inflation to shocks in monetary policy. With a SVAR model we can control for Type 1) and Type 3) inflation persistence. Clausen et al. (2006) provide a semi-structural VAR study on asymmetric effects of monetary policy in large EMU member countries and find that monetary transmission mechanisms in Germany, France and Italy are similar. Antipin and Luoto (2005) construct a SVAR model in which short-run interaction restrictions are derived from a simple, small-scale closed economy DSGE model. The authors report that price inflation responses to an unanticipated monetary policy shock could be seen to be asymmetric.

commonly applied in the monetary policy literature and the statistical properties of SVAR models are widely reported and known.

This paper provides updated empirical evidence on monetary policy transmission in the EMU area derived from the following contributions: first, we use updated EMU area data and a common reaction function across the EMU member countries and explicitly allow the size of the monetary policy shock to be the same across the member countries. Secondly, we derive posterior model probabilities to test the validity of ex-ante knowledge on the set of contemporaneous effects of the variables assumed to capture the monetary conditions for the EMU area. We rely on Bayesian inference since, for instance, posterior based error bands rather than classical confidence intervals allow us to use bands which characterize the shape of the likelihood more accurately, see Sims and Zha (1999).

The impulse response results obtained for an overidentified Bayesian SVAR model suggest that the EMU data lend support for existence of short-run asymmetric price inflation responses to an unexpected expansionary monetary policy shock in the EMU.

The rest of the paper is organized as follows: Section 2 presents econometric methods, Section 3 presents the data and results, and Section 4 comprises concluding remarks.

2. Econometric methods

European policy makers evince awareness of the existence of a delay between monetary policy action and its effect on inflation and on economies in general, since the ECB's declaration of medium-term price stability is widely accepted to constitute the first pillar of monetary policy in the EMU area and it is thus understood publicly that today's monetary policy actions are likely to have an impact on the future values of important macroeconomic variables such as inflation and output level. For the sake of dynamics, the monetary conditions in which the central bank needs to act should thus be seen as a

dynamic process involving multiple endogenous macroeconomic variables. Evidently, due to the aforesaid reasons we model monetary conditions for the EMU area applying a statistical model which captures both the evolution and the dynamics of an endogenous system of variables. The analysis in this paper is based on a SVAR model framework. The SVAR model takes the form

$$\Gamma_0 y_t = \delta + \sum_{i=1}^p \Gamma_i y_{t-i} + v_t, \quad (1)$$

where δ is a vector of constants, a nonsingular parameter matrix Γ_0 indicates how the variables listed in y_t simultaneously interact, matrices Γ_i contain parameters of lagged values of y_t , and unobservable structural shocks in v_t are assumed to be normally distributed with zero means and the diagonal covariance matrix denoted as Λ . The orthogonality property of structural shocks is typically assumed in the literature of SVAR models. The underlying idea of the SVAR approach is to impose theoretical restrictions³ on the data to identify structural shocks and then calculate the values of impulse response functions identified. In this study we identify the structural shocks of a SVAR model by specifying alternative short-run restriction schemes.

The literature lists a number of studies where the monetary policy transmission mechanism is examined using SVAR modeling methodology. To name but a few, Bernanke and Blinder (1992) analyze how unexpected changes in the Federal Funds Rate are transmitted to the U.S. economy, Sims (1992) explains the reasons for the price puzzle⁴ obtained in many VAR studies, Angeloni et al. (2003) compare euro area and U.S. monetary transmission mechanisms. Christiano et al. (1999) provide a survey of monetary policy SVAR models.

To capture the dynamics of the EMU area monetary conditions parsimoniously we collect in y_t the series of EMU area annual consumer price inflation (π_t), the output gap (x_t)

³ Short-run restrictions are set in the Γ_0 matrix and long-run restriction in matrices Γ_i .

⁴ A contradictory monetary policy action causes inflation to rise, whereas inflation is expected to drop.

which measures EMU area output deviations from steady state levels⁵, (r_t) to capture the status of monetary policy and ($\hat{\pi}_{t,j}$) to measure the annual consumer price inflation in a member country j . We thus specify $y_t = (\pi_t, x_t, r_t, \hat{\pi}_{t,j})'$ in a SVAR model for a member country j . The variables listed in y_t are in line with the models presented in an excellent survey of New Keynesian models by Clarida et al. (1999). Accordingly, we define price inflation (π_t) to capture the supply side and the output gap (x_t) to depict demand in the EMU area. As is common in the current monetary policy literature, the dynamics of monetary policy instrument (r_t) is modeled in the spirit of Taylor (1993); see also Hetzel (2000). Also in this paper, the Taylor rule -type reaction function interest rate responds to the output gap and inflation. The monetary transmission channel is the interest rate, since the central bank is assumed to be able to affect economic conditions by adjusting the real interest rate and thus affect aggregate consumption decisions; see Walsh (2003). The orthogonal property of structural shocks implies that for instance the cost-push shock of the inflation equation is independent of any monetary policy shock and vice versa. Evidently, the advantage of estimating individual member country-specific inflation and the EMU area aggregates simultaneously in the same model is that it enables efficient statistical investigation of possible asymmetries in the monetary policy transmission mechanism.

The above variables entering the SVAR model are assumed to adequately capture the monetary conditions in the EMU area. The member country-specific output variable is excluded from the variable list since the weight of a domestic output in relation to the EMU aggregate is minor and in general the variation in member-country output and inflation series can be seen to be driven by the interest rate. This is because of in the European Union both capital and labor force are free to move frictionlessly across the national borders. This paper comprises an analysis for twelve EMU member countries,

⁵ Galí and Gertler (1999) use the labor share of output as a proxy for marginal costs. However, Neiss and Nelson (2003), on the contrary, using data for the United States, the United Kingdom, and Australia report that labor costs do not suffice to explain inflation dynamics as well as the output gap. Hence, we describe marginal costs with a measure of the output gap.

i.e. Belgium, Germany, Greece, Spain, France, Ireland, Italy, Luxembourg, Netherlands, Austria, Portugal and Finland.

As in Sims and Zha (1999), the SVAR model in Eq. (1) is reparameterized such that

$$A_0 y_t = \delta + \sum_{i=1}^p A_i y_{t-i} + \eta_t, \quad (2)$$

where $A_0 = \Lambda^{-1/2} \Gamma_0$ and $\eta_t = \Lambda^{-1/2} v_t$. Hence $\eta_t \sim N(0, I) = I$ due to normalization. Thus $\text{Var}(A_0^{-1} \eta_t) = \Sigma = (A_0' A_0)^{-1}$.

The likelihood function of a SVAR model in Eq. (2) is

$$\begin{aligned} L(Y|X, \Sigma) &\propto |\Sigma|^{-0.5T} \exp\{-0.5 \text{tr}(\Sigma^{-1} E' E)\} \\ &= |\Sigma|^{-0.5T} \exp\{-0.5 \text{tr}(\Sigma^{-1} S) - 0.5 \text{tr}((B - \hat{B})' X' X (B - \hat{B}) \Sigma^{-1})\} \end{aligned} \quad (3)$$

where $E = (Y - XB)'(Y - XB)$, $S = (Y - X \hat{B})'(Y - X \hat{B})$, and the t th rows of Y , X , E are given by y_t' , $(I, y_{t-1}', \dots, y_{t-p}')$ and v_t , respectively. The matrix B is obtained by stacking the matrix product $(A_0^{-1} A_i)'$ and $\hat{B} = (X' X)^{-1} X' Y$ is a matrix of OLS parameter estimates.

As already noted, the matrix A_0 for the short-run effects of variables in y_t is the focal point of this study. We seek information on how variables in vector y_t simultaneously interact and thereby identify the impulse response functions of an estimated SVAR model. The traditional mode of SVAR model identification is to assume recursive restrictions i.e. Cholesky decomposition⁶. To allow for different simultaneous effects among the variables in y_t one needs to abandon the Cholesky restrictions and identify the SVAR model with different simultaneous effect restrictions. In specifying restrictions

⁶ The Cholesky decomposition supplies an exactly identified model. Setting underidentifying restrictions in matrix A_0 is of no interest, since in that case we cannot separate out the effects of a structural shock to model variables.

other than recursive restrictions we need ensure that the assumed simultaneous restrictions do not lead to an underidentified SVAR model. To avoid underidentification issues we verify that simultaneous restrictions fulfil the rank condition⁷ for identification. See Giannini et al. (1997) for a discussion of identification of SVAR models in econometrics.

We consider 7 different simultaneous effect schemes to identify the SVAR model in Eq. (2). The Cholesky restrictions (${}_{7}A_0$) and the six other suggested identification schemes for contemporaneous values of $y_t = (\pi_t, x_t, r_t, \hat{\pi}_{t,j})'$ in a SVAR model are as follows:

$${}_{1}A_0 = \begin{pmatrix} a_1 & a_{12} & 0 & 0 \\ 0 & a_2 & 0 & 0 \\ a_{31} & a_{32} & a_3 & 0 \\ 0 & 0 & a_{43} & a_4 \end{pmatrix}, \quad {}_{2}A_0 = \begin{pmatrix} a_1 & a_{12} & 0 & 0 \\ 0 & a_2 & 0 & 0 \\ a_{31} & a_{32} & a_3 & 0 \\ 0 & a_{42} & a_{43} & a_4 \end{pmatrix}, \quad {}_{3}A_0 = \begin{pmatrix} a_1 & 0 & 0 & 0 \\ a_{21} & a_2 & 0 & a_{24} \\ a_{31} & a_{32} & a_3 & 0 \\ 0 & 0 & a_{43} & a_4 \end{pmatrix}$$

$${}_{4}A_0 = \begin{pmatrix} a_1 & a_{12} & 0 & 0 \\ 0 & a_2 & 0 & 0 \\ a_{31} & a_{32} & a_3 & a_{34} \\ 0 & a_{42} & a_{43} & a_4 \end{pmatrix}, \quad {}_{5}A_0 = \begin{pmatrix} a_1 & 0 & a_{13} & 0 \\ 0 & a_2 & 0 & 0 \\ a_{31} & a_{32} & a_3 & a_{34} \\ 0 & a_{42} & a_{43} & a_4 \end{pmatrix}, \quad {}_{6}A_0 = \begin{pmatrix} a_1 & 0 & a_{13} & 0 \\ 0 & a_2 & 0 & 0 \\ a_{31} & a_{32} & a_3 & a_{34} \\ 0 & 0 & a_{43} & a_4 \end{pmatrix}$$

and the Cholesky identifying restrictions are

$${}_{7}A_0 = \begin{pmatrix} a_1 & 0 & 0 & 0 \\ a_{21} & a_2 & 0 & 0 \\ a_{31} & a_{32} & a_3 & 0 \\ a_{41} & a_{42} & a_{43} & a_4 \end{pmatrix}.$$

In the above matrices a_{kj} s denote the simultaneous effect of variable j on variable k . The lower-triangular matrix ${}_{7}A_0$ is a Cholesky factor of the covariance matrix Σ . The

⁷ See for instance Hamilton (1994). pp. 334, presenting a method to check the rank condition for identification.

restriction schemes ${}_4A_0$, ${}_5A_0$ and ${}_7A_0$ provide an exactly identified SVAR model, whereas other schemes constitute an over-identified SVAR model. All the matrices ${}_iA_0$ above fulfil the rank condition for SVAR model identification.

A closer inspection of the foregoing identification matrices reveals that besides the matrices ${}_4A_0$, ${}_5A_0$ and ${}_6A_0$, the monetary policy (instrumented by r_t) is allowed only simultaneously to be affected by EMU price inflation (π_t) and the output gap (x_t), which is in accordance with the declared ECB monetary policy targets. By specifying restrictions ${}_4A_0$, ${}_5A_0$ and ${}_6A_0$ we suggest that domestic consumer price inflation can have weight in the ECB's monetary policy decision-making by allowing member country inflation ($\hat{\pi}_{t,j}$) to have a simultaneous effect on a monetary policy instrument (r_t) (nonzero a_{34}) together with EMU aggregates (π_t) and (x_t). The difference between the restrictions in ${}_1A_0$ and ${}_2A_0$ is that the EMU area output gap is also allowed simultaneously to affect j th member country inflation dynamics (a_{42}). Matrix ${}_3A_0$ exhibits such restrictions that the EMU area inflation (π_t) cannot be seen to be contemporaneously affected by the EMU demand side (x_t) ($a_{13} = 0$), and monetary policy is assumed to have a simultaneous impact only on member country inflation⁸. Restrictions driven in ${}_2A_0$ and ${}_4A_0$ are almost the same except that member country price inflation ($\hat{\pi}_{t,j}$) is allowed simultaneously to affect the value of monetary policy instrument (r_t) in ${}_4A_0$. Restrictions in ${}_5A_0$ suggest that monetary policy simultaneously affects EMU area and member country inflation variables. Restrictions in ${}_6A_0$ exhibit restrictions similar to ${}_5A_0$ but member country inflation is not allowed simultaneously be affected by the EMU-wide output gap (x_t).

There are seven different competing identification schemes constituting 7 models among which we should choose. We apply posterior model probabilities to find the most likely restrictions matrix for the SVAR model. Given the data Y and seven rivaling

⁸ Since the EMU area inflation is a population- and GDP -weighted average of member country inflations and unanticipated movements in monetary policy instrument are diluted while averaging over member countries figures.

identification schemes, the posterior model probabilities in SVAR models identified with restrictions $iA_0, i = 1, \dots, 7$ can be given by

$$p(\text{Model}_k|Y) = \frac{p(Y|\text{Model}_k)p(\text{Model}_k)}{\sum_{i=1}^7 p(Y|\text{Model}_i)p(\text{Model}_i)}, \quad (4)$$

where the marginal likelihood of model k is defined as

$$p(Y|\text{Model}_k) = \int p(Y|\theta_k, \text{Model}_k)p(\theta_k|\text{Model}_k)d\theta_k.$$

Model_k and parameter vector θ_k refer to a SVAR model in Eq. (2) identified with iA_0 restrictions. $p(\theta_k|\text{Model}_k)$ is prior density function of θ_k under model k ; $p(Y|\theta_k, \text{Model}_k)$ is the likelihood function. We assume that the prior model probability of model k , $p(\text{Model}_k)$, is the same (one over seven, i.e. $1/7$) for all seven SVAR models.

We follow Garratt et al. (2007) in conducting the model selection and approximate the value of the marginal likelihood of model k . In line with Garratt we base the model probability analysis on Schwarz (1978), who presents an asymptotic approximation to the marginal likelihood function of the form

$$\log p(Y|\text{Model}_i) \approx l - K \times \log(T)/2, \quad (5)$$

where l is the log of the likelihood function evaluated at maximum likelihood estimates, K is the number of parameters and T is the number of available observations.

To measure posterior model probabilities in Eq. (4) we specify a likelihood function of a SVAR model for given restrictions iA_0 ($i = 1, \dots, 7$). For a Cholesky-restricted SVAR model (γA_0) the concentrated likelihood function evaluated at maximum likelihood estimates \hat{B} and $\hat{\Sigma}$ takes the form

$$\begin{aligned}
L(Y|X, {}_7A_0) &= (2\pi)^{-0.5Tm} \left| \frac{S^*}{T} \right|^{-0.5T} \exp \left\{ -0.5 \text{trace} \left(\left(\frac{S^*}{T} \right)^{-1} S^* \right) \right\} \\
&= (2\pi)^{-0.5Tm} \left| \frac{S^*}{T} \right|^{-0.5T} \exp \{-0.5Tm\},
\end{aligned} \tag{6}$$

where $S^* = ({}_7A_0' A_0)^{-1} = (Y-X\hat{B})'(Y-X\hat{B})$ under Cholesky restrictions and the maximum likelihood estimate of Σ is hence $S = (Y-X\hat{B})'(Y-X\hat{B})/T$, where $\hat{B} = (X'X)^{-1}X'Y$. The trace of an identity matrix $I_{m \times m}$ is m , the number of diagonal elements and m is the number of variables.

The concentrated likelihood function evaluated at maximum likelihood estimates for SVAR models identified with other than Cholesky restrictions, i.e. ${}_iA_0, i \neq 7$, is

$$\begin{aligned}
L(Y|X, {}_iA_0) &= (2\pi)^{-0.5Tm} \left| ({}_iA_0' A_0)^{-1} \right|^{-0.5T} \exp \left\{ -0.5 \text{trace} ({}_iA_0 S {}_iA_0') \right\} \\
&= (2\pi)^{-0.5Tm} \left| {}_iA_0 \right|^T \exp \left\{ -0.5 \text{trace} ({}_iA_0 S {}_iA_0') \right\}.
\end{aligned} \tag{7}$$

To obtain a value for Eq. (4) we maximize Eq. (6) for a Cholesky-restricted SVAR model and Eq. (7) for SVAR models identified with restrictions ${}_iA_0, i = 1, \dots, 6$.

With a SVAR model with suitable simultaneous restrictions in matrix A_0 we will track down whether there exist member country-specific asymmetric price inflation responses to an unanticipated expansionary common monetary policy shock. Following Sims and Zha (1999) we update non-informative prior knowledge of the reduced form parameter values of a VAR model with the information summarized by the likelihood function. Sims and Zha (1999) presume flat prior distributions for A_0 and B of a SVAR model identified with non-recursive restrictions (${}_iA_0, i \neq 7$). The full conditional and marginal posterior densities for SVAR model specified with non-recursive restrictions in A_0 are

$$\beta|X, Y, {}_iA_0 \sim N\left(\hat{\beta}, \left({}_iA_0' {}_iA_0\right)^{-1} \otimes (X'X)^{-1}\right) \quad (8)$$

and

$$q\left({}_iA_0|X, Y\right) \propto \left|{}_iA_0\right|^{(T-k)} \exp\{-0.5\text{trace}({}_iA_0 S {}_iA_0)\}, \quad (9)$$

where $k = mp + 1$. The full conditional posterior distribution in Eq. (8) is the multivariate normal and the marginal posterior distribution in Eq. (9) is not in a form of standard distribution, so that we need to use numerical integration methods to draw samples from it.

Having the Jeffrey's non-informative as the joint prior p.d.f. for B and Σ in a Cholesky identified SVAR model gives the marginal posterior distribution of Σ the following form

$$\Sigma|X, Y \propto |\Sigma|^{\frac{-(T-(mp+1)+m+1)}{2}} \exp\{-0.5\text{trace}(\Sigma^{-1}S)\}. \quad (10)$$

Eq. (10) is the kernel of the inverse Wishart distribution for Σ , i.e. $\Sigma \sim iW_m(S, T-(pm+1))$. The parameters β in a Cholesky identified SVAR model follow the multivariate normal distribution of Eq. (8).

Vectors β and $\hat{\beta}$ in Eqs (8) – (10) are formed by stacking the columns of B and \hat{B} , respectively. The motivation for using the Jeffrey's prior in a Cholesky identified SVAR model is that the posterior distributions for B and Σ are known and drawing samples from these is trivial. The information content of a Jeffrey's prior is in practice the same as in a flat prior for B and ${}_iA_0$, $i = 1, \dots, 6$ which Sims and Zha (1999) suggest to be used in non-recursive identification schemes. For a good reference for the Bayesian statistics one might consider Zellner (1971).

To analyze the possibility of asymmetric price inflation responses to a common monetary policy shock we draw impulse responses for SVAR models identified with simultaneous restrictions which are supported by the data. For the impulse responses the size of a shock

in monetary policy instrument (r_t) is normalized to one standard deviation in a SVAR model in Eq. (2). The properties of the standard impulse response function for linear models are well known and documented in the literature; see for example Hamilton (1994) and Sims and Zha (1999). For a general case we define the standard impulse response function by letting c_{lk} be the response of variable $y_{l,t+s}$ to shock $\eta_{k,t}$, i.e.,

$$c_{lk,s} = \frac{\partial y_{l,t+s}}{\partial \eta_{k,t}}. \quad (11)$$

The values of the response function depend only on the parameters of the structural model of Eq. (2), and the values can be obtained using basic matrix operations.

3. The Data and Results

We operate with monthly EMU data spanning the period from 1999:1 to 2007:10 (106 observations). The data are collected from the sources of the online data bank services of the EuroStat. The series for the HICP (harmonized index of consumer prices) of twelve EMU countries and the EMU area aggregate are neither work-day nor seasonally adjusted. Seasonally adjusted series of the index of industrial production⁹ (IIP) (excluding construction) in the EMU area are used in the formation of the output gap (x_t). The IIP series look back to the year 1980. Monthly values for Eonia-12 are used as a proxy for the ECB's monetary policy instrument (r_t). The series for Eonia-12 interest rate is calculated using the day-to-day interest rates without seasonal adjustment. The reference year for all HICP series is 2005, and year 2000 is the reference year for the EMU area IIP series.

⁹ One could use aggregate GDP series instead of IIP series, but the problem is that there are no monthly data available for the GDP in the EMU area. Furthermore, we could consider the IIP series to depict the manufacturing sector more accurately. Aksoy, De Grauwe and Dewachter (2002) use monthly industrial production series in their study tracking down the impact of economic and institutional asymmetries on the effectiveness of monetary policy in the euro zone with an explicit policy target rule.

The output gap (x_t) is measured as the logarithmic difference between the actual and the potential output level. The logarithm of the potential output is proxied by a one-sided Hodrick-Prescott (HP) trend estimate of the unobserved trend component τ_t in a model

$$g_t = \tau_t + \zeta_{1t}, \quad (12)$$

$$(1 - L)^2 \tau_t = \zeta_{2t}, \quad (13)$$

where g_t is the logarithm of a measure of actual output, L is the lag operator and ζ_{1t} and ζ_{2t} are mutually uncorrelated white noise sequences with a relative variance of $q = \text{var}(\zeta_{1t}) / \text{var}(\zeta_{2t})$. The value of $q = 0.67 \times 10^{-3}$ is taken from Stock and Watson (1999).

The price inflation series, π_t and $\hat{\pi}_{t,j}$, are constructed¹⁰ on an annual basis for both the EMU area and member countries, respectively.

Figures 1-5 in Appendix section A plot the series of annual HICP inflation ($\hat{\pi}_{t,j}$) for EMU member countries together with the EMU area HICP inflation (π_t) and the Eonia interest rate (r_t). Figure 1 shows the annual HICP inflation in the EMU area to be more or less an average of inflation figures for Germany, France and Italy. Additionally, the inflation series plotted in Figure 1 tend all to converge to an overall 2 percent inflation target. In Figure 2 the annual HICP inflation series for the EMU area, Belgium, Greece and Spain are plotted against time. The series for Spain and Greece vary similarly at higher levels than those for Belgium and the EMU area. Convergence to the overall inflation target is not evident for these member countries.

Figure 3 implies that annual inflation series for Finland have been at lower levels than in any other EMU country. The HICP inflation in Ireland and Portugal has been historically higher than on average in the EMU area. Figure 4 shows that since the beginning of 2003 HICP inflation in Netherlands and Austria have followed the EMU inflation. Meanwhile, the annual HICP inflation in Luxembourg has been fluctuating relatively strongly,

¹⁰ We use 1998 HICP values in calculating 1999 inflation figures.

indicating no convergence to the overall 2 per cent annual target. Thus, a striking observation is that the aggregate EMU area annual HICP inflation has varied closely around the declared inflation target, while the inflation series for member countries have been fluctuating at different levels.

Finally, in Figure 5 the Eonia interest rate (r_t), output gap (x_t) and EMU area inflation (π_t) are plotted. The output gap, as a proxy variable for marginal costs, has not followed a constant pattern – it has been fluctuating mainly on the negative side. Observations on consumer price inflation and the output gap suggest that in the EMU inflation stabilization is allocated greater weight while the ECB decides the optimal value of monetary policy instrument (r_t). From Figures 1-4 we hypothesize that asymmetric price inflation responses to a monetary policy shock are to be expected due to the somewhat heterogeneous HICP inflation dynamics among the EMU member countries.

The posterior model probabilities in Eq. (4) are calculated for seven SVAR models for each member country. Specifically, a SVAR model in Eq. (2) with restrictions iA_0 ($i = 1, \dots, 7$) and Eq. (6) and (7) are maximized respectively conditional on the member country data. The data we feed into Eq. (6) and (7) are $y_t = (\pi_t, x_t, r_t, \hat{\pi}_{t,j})'$, where $\hat{\pi}_{t,j}$ is the annual HICP inflation in the j th member country. This means that for each member country we get 7 posterior model probabilities, one for each identification scheme. Posterior model probabilities are reported in Table 1 below. A lag length of five (5) was chosen, since it turned out to be the shortest lag length providing homoscedastic and autocorrelation-free SVAR model errors.

SVAR(p), p=5	Posterior model probabilities						
Member country	Identification scheme						
	${}_1A_0$	${}_2A_0$	${}_3A_0$	${}_4A_0$	${}_5A_0$	${}_6A_0$	${}_7A_0$
Belgium	0.0000	0.0000	0.3223	0.2201	0.0237	0.1968	0.2371
Germany	0.0000	0.0000	0.2322	0.1795	0.0180	0.3908	0.1795
Ireland	0.1484	0.0000	0.2136	0.1419	0.0221	0.2530	0.2209
Greece	0.0004	0.0001	0.0004	0.0597	0.0087	0.8432	0.0875
Spain	0.0000	0.0000	0.1079	0.1459	0.0208	0.5170	0.2084
France	0.0000	0.0000	0.6105	0.0888	0.0218	0.0608	0.2181
Italy	0.0006	0.0005	0.0490	0.2747	0.0509	0.1156	0.5088
Luxembourg	0.0000	0.0000	0.1294	0.1799	0.0180	0.4929	0.1799
Netherlands	0.4613	0.0562	0.0068	0.0571	0.0072	0.3390	0.0723
Austria	0.0000	0.0000	0.0000	0.1679	0.0168	0.6475	0.1679
Portugal	0.8359	0.0981	0.0153	0.0072	0.0013	0.0291	0.0131
Finland	0.0000	0.0000	0.6870	0.0716	0.0072	0.1627	0.0716

Table 1. Posterior model probabilities. Bolded figures indicate the most probable identification scheme for a member country

The highest posterior model probability of a member country is highlighted in bolded font in Table 1. The last column indicates that Cholesky restrictions (${}_7A_0$) are relatively weakly supported in the data. Only for Italy do Cholesky restrictions seem to produce the best model fit. For the rest of the member countries the model fit of Cholesky restrictions is more or less moderate. Generally, the data support restrictions ${}_7A_0$, ${}_3A_0$ and ${}_6A_0$ and restrictions according to ${}_2A_0$ are in fact faintly supported. A slightly striking finding is the posterior model probability of a SVAR model under ${}_5A_0$ restrictions is low despite that the restriction scheme being very similar to ${}_6A_0$ restrictions. The difference between ${}_5A_0$ and ${}_6A_0$ restrictions is that in ${}_5A_0$ the output gap (x_t) is allowed simultaneously to affect member country inflation ($\hat{\pi}_{t,j}$). It emerges from Table 1 that ${}_6A_0$ restrictions are best supported in the data i.e. the restrictions which allow the EMU area (π_t) and member country consumer price inflation to be both simultaneously affected by monetary policy shock. Furthermore, from the posterior model probabilities for identification schemes ${}_4A_0$, ${}_5A_0$ and ${}_6A_0$ we see that the data also lend support to the conception that the ECB takes into account the inflation of an individual member country.

To attain identified impulse responses we restrict the analysis to a SVAR model using δA_0 and γA_0 restrictions. The Cholesky restrictions (γA_0) are also taken into the impulse response analysis, since typically SVAR models are identified with a recursive identification scheme. For a Cholesky identified (γA_0) SVAR model it is assumed that the contemporaneous effect of monetary policy (r_t) on EMU area inflation (π_t), i.e. contemporaneous interest rate elasticity, is zero by definition. This is not the case with δA_0 restrictions, since the monetary policy can have an immediate effect on both the EMU area inflation (π_t) and member country inflation ($\hat{\pi}_{t,j}$). In total we will be estimating 24 SVAR models, two for each member country – one SVAR model identified with Cholesky restrictions (γA_0) and one with δA_0 restrictions.

A SVAR model with δA_0 restrictions is such that the posterior p.d.f. in Equation (9) is not in the form of standard p.d.f. To generate a Monte Carlo sample from the posterior of δA_0 we use a version of the random walk Metropolis algorithm for Markov Chain Monte Carlo (MCMC). The algorithm uses multivariate normal distribution for the jump distribution on changes in parameters in δA_0 . We first simulate 15,000 draws using a diagonal covariance with diagonal entries 0.00001 in the jump distribution. These draws are then used to estimate the posterior covariance matrix of parameters in δA_0 and scale it by the factor $2.4^2/9$ to obtain an optimal covariance matrix for the jump distribution; see Gelman et al. (2004). In estimating the SVAR models identified with δA_0 restrictions, we use 100,000 draws, discarding the first 10,000 as a burn-in period. As a convergence check three chains with different starting values are simulated. For each chain we pick every 100th draw to achieve a nearly independent sample. The potential scale reduction factor of Gelman and Rubin (1992) is between 1 and 1.08 for each parameter in δA_0 . The multivariate version of Gelman and Rubin's diagnostic, proposed by Brooks and Gelman (1997), is between 1.00 and 1.05. Finally, the frequencies of accepted jumps are roughly 0.24. Eventually our results for δA_0 restricted SVAR models are based on 2,700 draws for each member country. For a Cholesky identified (γA_0) SVAR model we generate 3000 draws from p.d.fs given in Eq. (8) and Eq. (10). The conditional posterior p.d.f is multivariate normal and the marginal posterior p.d.f. of Σ is inverse Wishart distribution, as already noted.

When computing the posterior of impulse responses we follow Sims and Zha (1999) and calculate Bayesian 68-percent error bands. In Figures 6-17 in Appendix section B, for each member country in turn, the impulse responses drawn are

$$D_{s,j} = \frac{\partial \pi_{t+s}}{\partial \eta_{r,t}} - \frac{\partial \hat{\pi}_{j,t+s}}{\partial \eta_{r,t}} \text{ for } s = 0, \dots, 12 \text{ and } j = 1, \dots, 12. \quad (14)$$

The first term in Eq. (14) is the annual EMU area inflation response to an unanticipated, one standard deviation expansionary monetary policy shock. The latter term in Eq. (14) is the member country's inflation response. Black lines in Figures 6-17 are for a SVAR model with ${}_6A_0$ restrictions and gray, dotted lines a SVAR model identified with Cholesky restrictions, ${}_7A_0$. In both identification schemes the middle line is the median impulse response value. For both ${}_6A_0$ - and ${}_7A_0$ -identified SVAR models we calculate impulse responses up to 13 periods (the length of a period is 1 month), including the shock period denoted as time 0 in the figures. If 68-percent error bands contain the value $D_{s,j} = 0$, then the inflation responses are statistically the same in the EMU area and in member country j at 68-percent posterior probability.

Next we will first discuss the impulse responses drawn for Cholesky identified SVAR models and thereafter comment on impulse responses obtained from an overidentified SVAR model with restrictions in ${}_6A_0$. For Cholesky identified SVAR models, the monetary response of EMU inflation (π_t) is identically zero, i.e. $\frac{\partial \pi_{t+s}}{\partial \eta_{r,t}} = 0$ for $s = 0$. This

implies that the immediate ($s = 0$) impulse response value is dictated solely by the second term $\frac{\partial \hat{\pi}_{j,t+s}}{\partial \eta_{r,t}}$ in Eq. (14). Thus, if the immediate response of j th member country inflation ($\hat{\pi}_{j,t}$) to a shock in monetary policy instrument ($\eta_{r,t}$) is positive, it will be shown in Figures 6-17 such that the $D_{0,j}$ assumes negative value.

Impulse responses drawn for a Cholesky identified SVAR model with $\hat{\pi}_{t,j}$ series for Belgium, Germany, Greece, France and Finland imply that the immediate inflation responses are asymmetric with 68-percent posterior probability for these countries. However, with the exception of Greece, inflation responses for later periods are statistically the same as the EMU inflation response. This means that the posterior intervals contain the zero level of $D_{s,j}$ for $s > 0$. Figure 9 for Greece shows that during the last 5 periods (i.e. $s = 8, \dots, 12$) drawn responses exhibit persistent asymmetric inflation responses.

Drawn differences between the impulse response of the EMU and Luxembourg, Dutch, Portuguese and Italian inflation convey asymmetric inflation responses. Specifically, the Luxembourg inflation response is statistically more moderate than the EMU inflation response between 3 and 7 months after the initial monetary policy shock. Portuguese asymmetric inflation responses begin 5 months after the shock and have ever since remained different from those of EMU area inflation. The Dutch responses are more aggressive/moderate than the EMU inflation one/eleven months after the shock. Two months after the shock the inflation response in Italy is more moderate than the EMU response for 1 month. We see that the results from the Cholesky identified SVAR models suggest heterogeneous type 3 inflation persistence among the EMU member countries (see footnote on page 3). Inflation responses in Ireland, Spain and Austria are statistically the same as the EMU inflation responses. Symmetric responses imply that an unanticipated monetary policy shock does not produce statistically significant terms of trade divergence.

Posterior distributions show that immediate inflation responses ($D_{0,j}$) are mixed for $6A_0$ -identified SVAR models. In the model for Germany, France, Italy and Finland the immediate inflation response is statistically stronger than the EMU area inflation response. For Belgium, Ireland, Greece, Spain, Luxembourg and Austria the adjustment to a shock in monetary policy instrument is not as rapid as it is on average in the EMU area. The immediate inflation response in the Netherlands and Portugal is statistically the same as in EMU on average, as shown in Figures 14 and 16.

The δA_0 restricted SVAR model for the Belgian, Italian, Austrian and Finnish inflation responses shows that it takes 1 month for Belgium to adapt, Italy waits for 3 months, whereas Austria and Finland need 2 months to adjust. The difference between the EMU area and Greek inflation responses exhibit lagged response behavior (type 3 inflation persistence;) during the 3 and 6 months after the shock the difference in responses is statistically positive, indicating a more sluggish inflation adjustment process in Greece than in the EMU area. Inspecting the impulse responses drawn for Germany, France and Italy we see that inflation responses are similar, suggesting a similarity in price transmission mechanisms for monetary policy shock¹¹. This is in line with Clausen et al. (2006). Furthermore, for δA_0 restrictions no persistent asymmetric inflation responses can be obtained for any member country.

The important findings derived from the impulse response analysis are; firstly, there occur statistically significant asymmetric immediate ($s = 0$) inflation responses for SVAR models specified with both δA_0 and γA_0 restrictions for Belgium, Germany and Greece. Another finding is that the different adjustment speeds (compared to EMU inflation) in response to a monetary policy shock indicate that in the EMU area inflation persistence is heterogeneous among the member countries.

4. Conclusions

In this paper we provide empirical evidence of transmission of the ECB's monetary policy actions in annual consumer price inflation with updated monthly data. Evidence is obtained using an actual monetary policy instrument and error bands for impulse responses which characterize the true shape of the likelihood.

We calculate posterior model probabilities for SVAR models identified with a set of plausible identification schemes. We find that the data weakly support Cholesky factorization, while the strongest support goes to an identification scheme possessing

¹¹ Note that the total weight of Germany, Italy and France in constructing the EMU aggregate series is high.

overidentifying restrictions which also let the EMU member country inflation simultaneously interact with the monetary policy instrument and allow the monetary policy shock to have an immediate effect on the EMU area inflation and EMU member country inflation.

Given the impulse response function calculations based on the posterior distributions we may state that the data lend support to short-run asymmetric consumer price inflation responses to a monetary policy shock across the member countries in the EMU area. This verifies the hypothesis that the ECB's monetary policy conduct needs to be seen as a complicated task, because if the ECB conducts its monetary policy conditional on union-wide aggregates (where the target of aggregate EMU-wide inflation has a substantial role), unforeseen shocks in the monetary policy must be seen to have an asymmetric impact on consumer price inflation across the member countries.

One possible way to understand the asymmetric inflation responses addressed is for instance to allow nominal rigidity in firms' price-setting, i.e. assuming that firms in individual member countries follow Calvo (1983) pricing with different price adjustment probabilities. Under Calvo pricing firms may adjust their prices with some constant probability, and since the adjustment probabilities vary across the member countries, deviations from the optimal price level will occur when adjustments are needed. This will evidently show in asymmetric inflation responses to a common monetary policy shock.

Finally, as a consequence of asymmetric inflation responses and a fixed exchange rate across the member countries, unanticipated monetary policy actions will influence relative prices in member countries, causing disturbances in mutual price competition and thereby indirectly altering consumption schemes, this leads to changes in EMU member country welfare levels in the short-run.

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APPENDIX

A. Data Figures

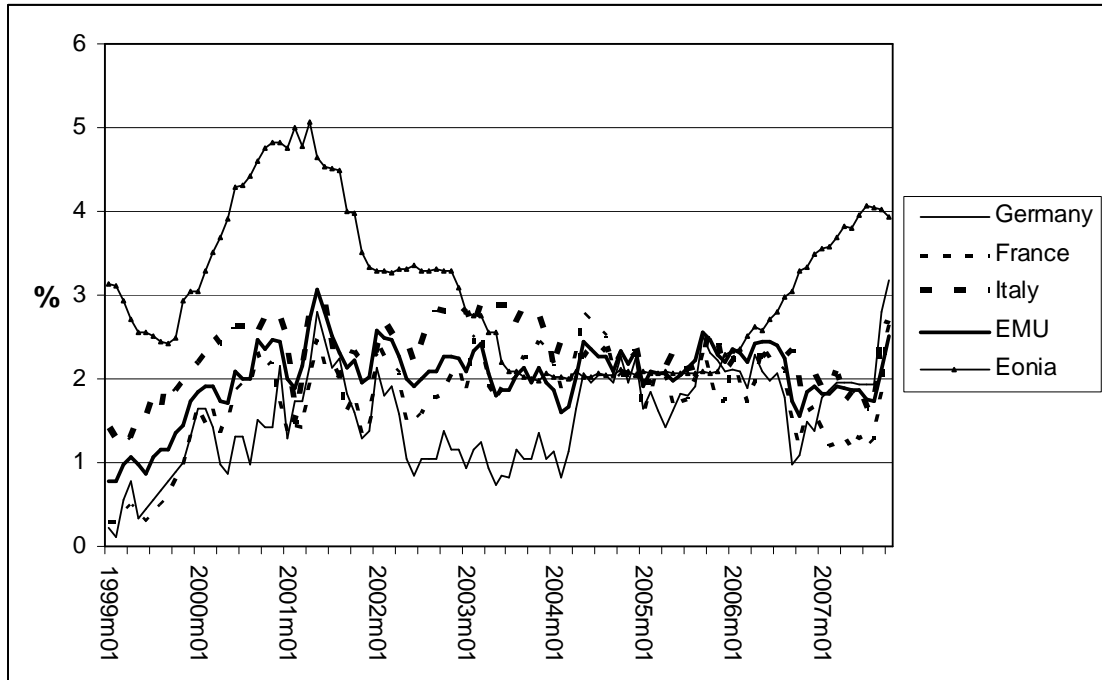


Figure 1. Annual HICP inflation rates for Germany, France, Italy and EMU area together with Eonia interest rate

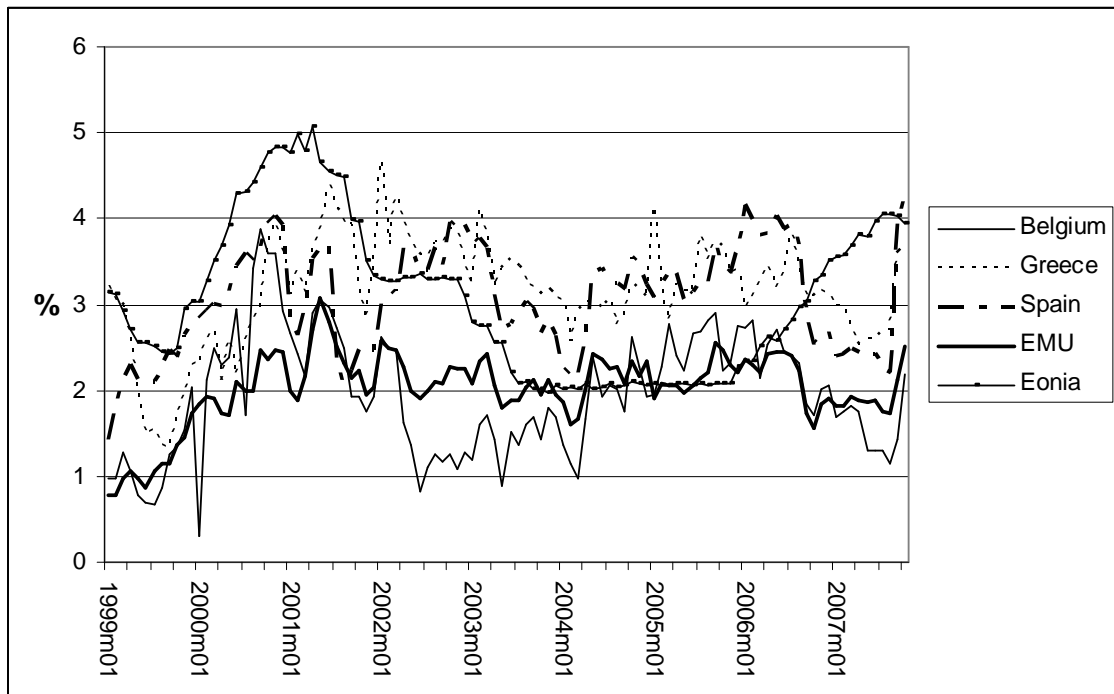


Figure 2. Annual HICP inflation rates for Belgium, Greece, Spain and EMU area together with Eonia interest rate

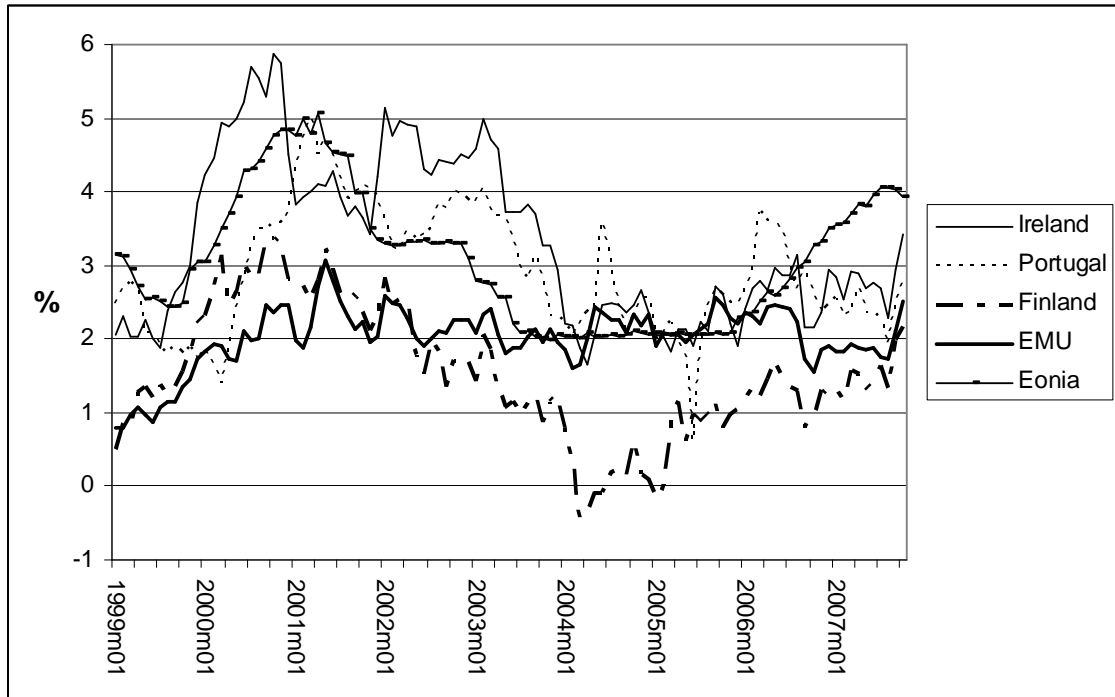


Figure 3. Annual HICP inflation rates for Ireland, Portugal, Finland and EMU area together with Eonia interest rate

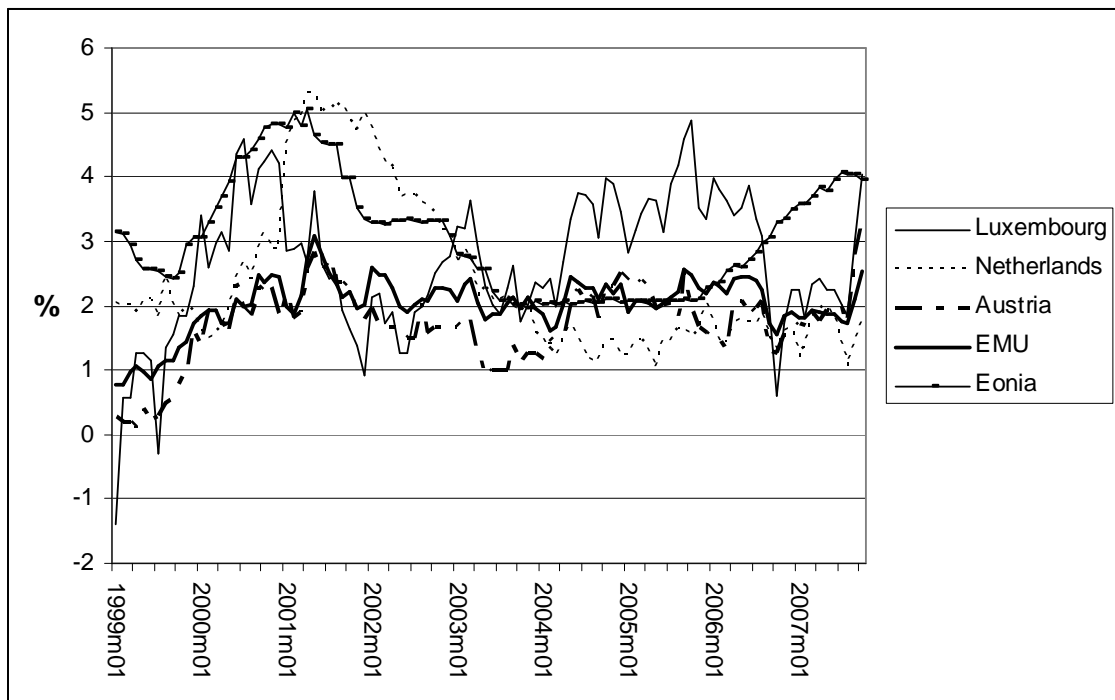


Figure 4. Annual HICP inflation rates for Luxembourg, Netherlands, Austria and EMU area together with Eonia interest rate

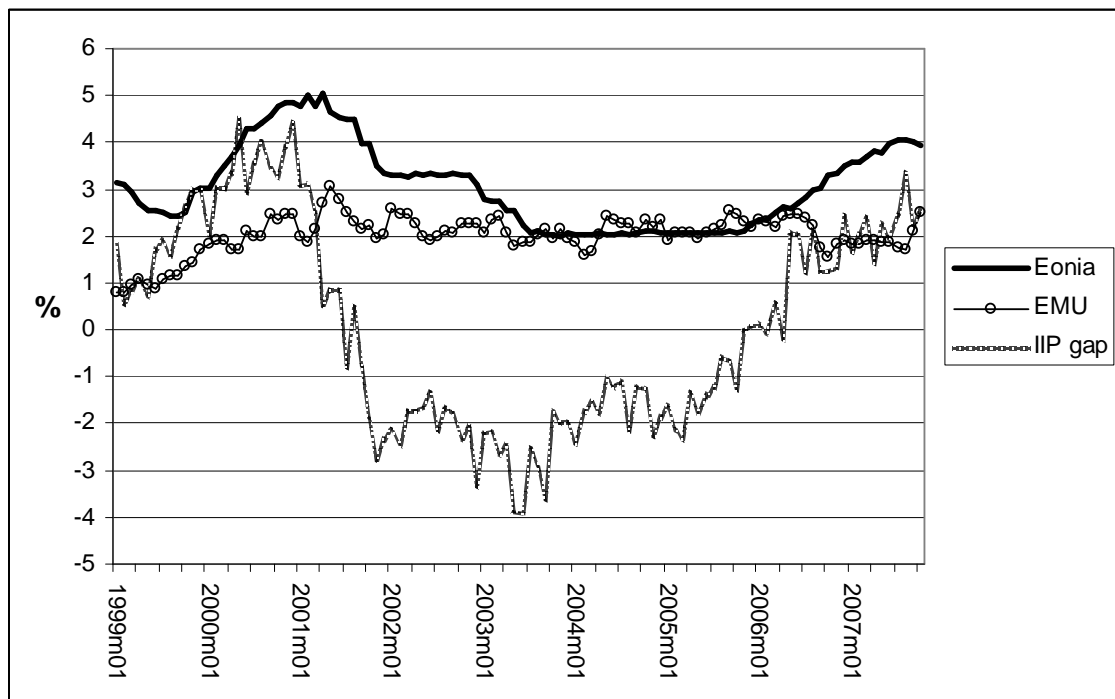
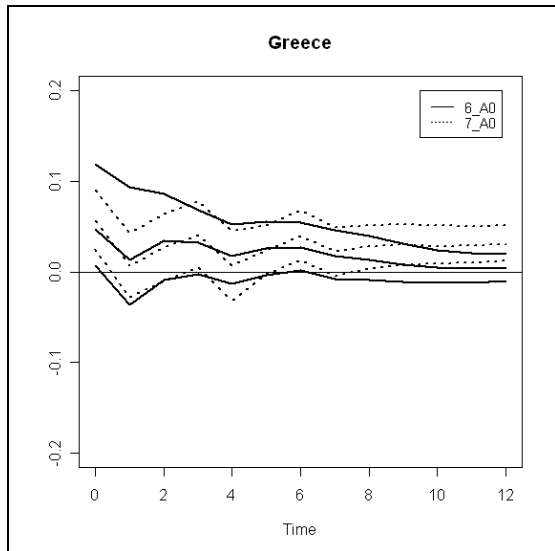
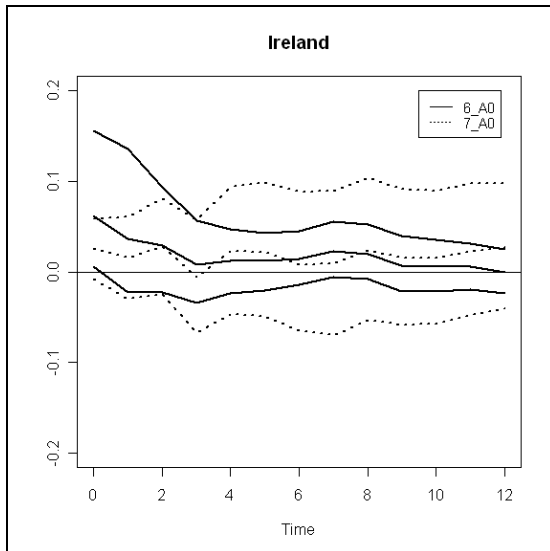
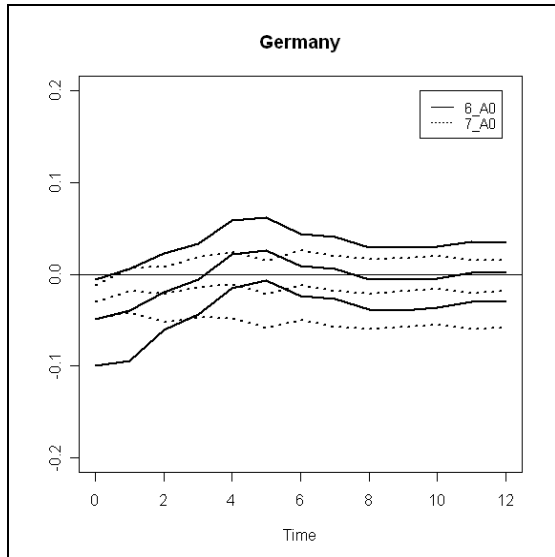
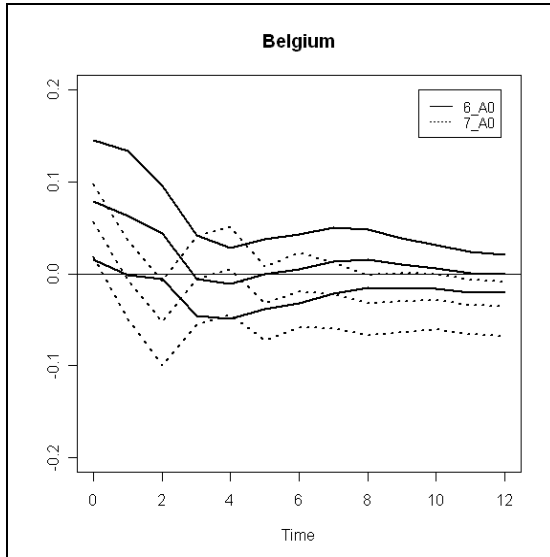


Figure 5. EMU area output gap, EMU area HICP inflation rate and Eonia interest rate

B. Impulse Responses for SVAR Model with ${}_6A_0$ and ${}_7A_0$ restrictions



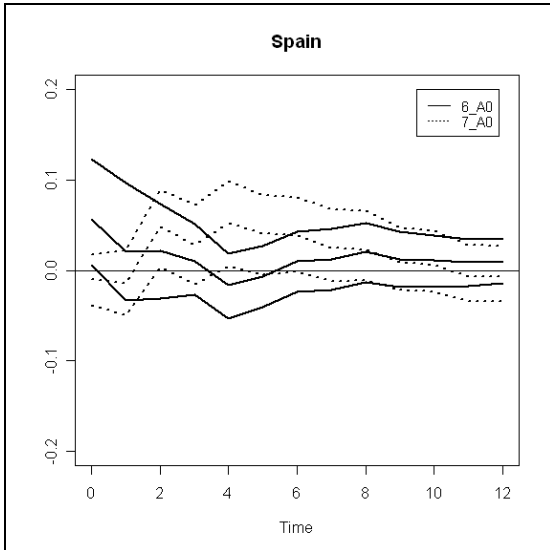


Figure 10. Spain

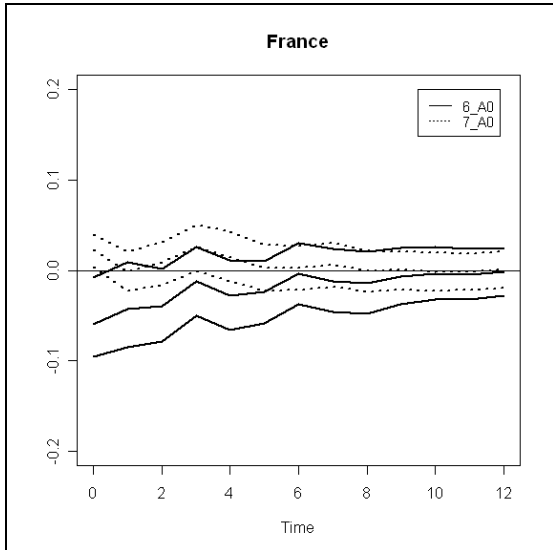


Figure 11. France

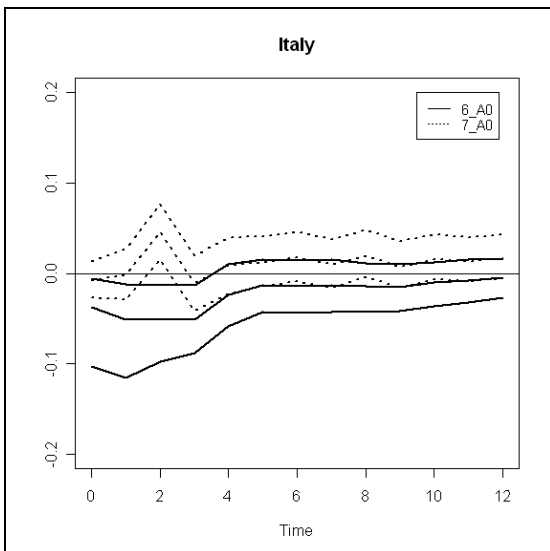


Figure 12. Italy

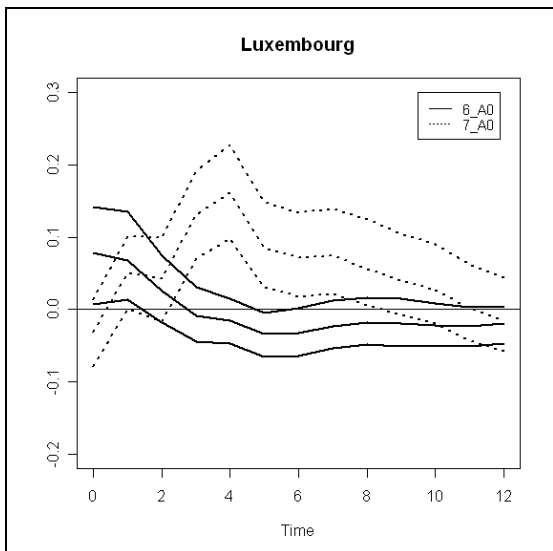


Figure 13. Luxembourg

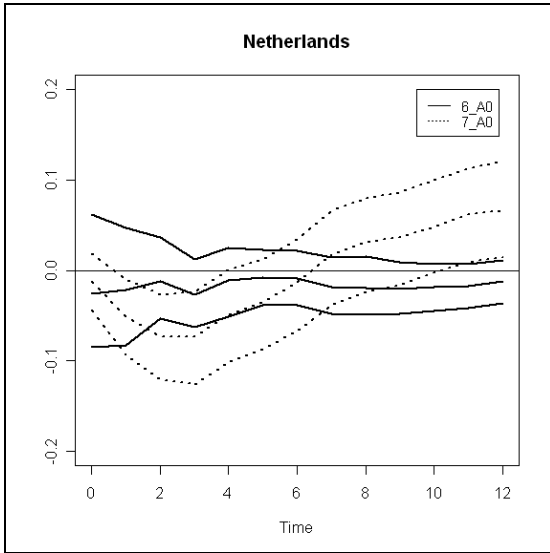


Figure 14. Netherlands

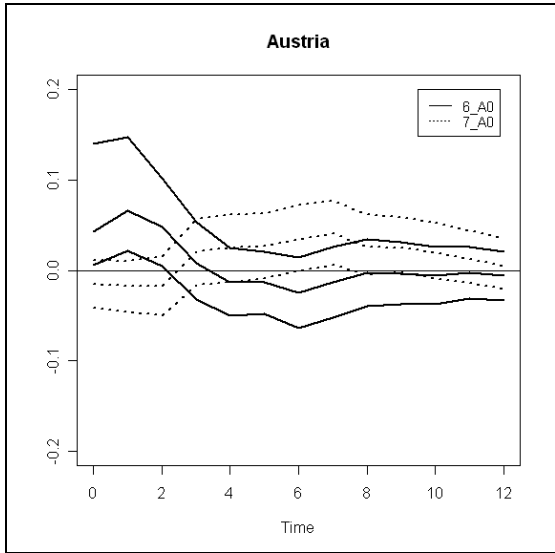


Figure 15. Austria

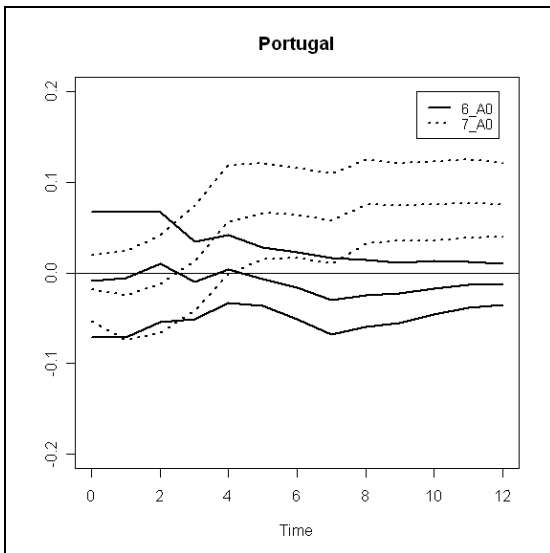


Figure 16. Portugal

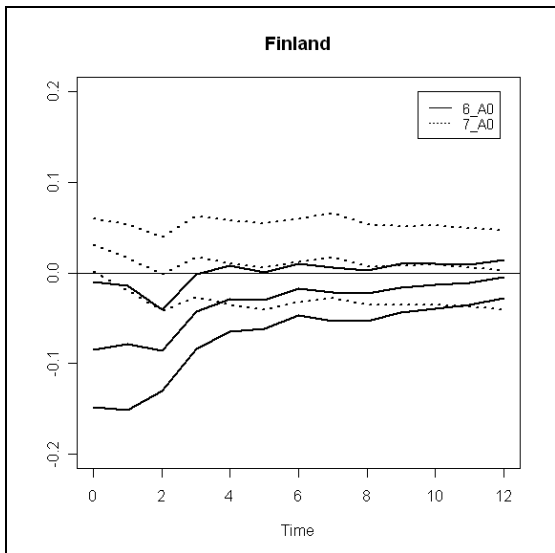


Figure 17. Finland