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Responses to Monetary Policy Shocks in the East and the West of Europe: A Comparison

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Editorial

On the occasion of the 65th birthday of Governor Klaus Liebscher and in recognition of his commitment to Austria's participation in European monetary union and to the cause of European integration, the Oesterreichische Nationalbank (OeNB) established a "Klaus Liebscher Award". It will be offered annually as of 2005 for up to two excellent scientific papers on European monetary union and European integration issues. The authors must be less than 35 years old and be citizens from EU member or EU candidate countries. The "Klaus Liebscher Award" is worth EUR 10,000 each. The winners of the second Award 2006 were Petra Geraats and Marek Jarociński. Marek Jarociński's winning paper is presented in this Working Paper, while Petra Geraats' contribution is contained in Working Paper 123.

In this paper Marek Jarociński compares responses to monetary shocks in the EMU countries (in the pre-EMU sample) and in the New Member States (NMS) from Central Europe. The small-sample problem, especially acute for the NMS, is mitigated by using a Bayesian estimation procedure which combines information across countries. A novel identification scheme for small open economies is used. The estimated responses are quite similar across regions, but there is some evidence of more lagged, but ultimately stronger price responses in the NMS economies. This contradicts the common belief that monetary policy is less effective in post-transition economies, because of their lower financial development. NMS also have a probably lower sacrifice ratio, which is consistent with the predictions of both the imperfect information model of Lucas (1973) and the New-Keynesian model of Ball et al. (1988).

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Responses to Monetary Policy Shocks in the East and the West of Europe: A Comparison

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Abstract

This paper compares responses to monetary shocks in the EMU countries (in the pre-EMU sample) and in the New Member States (NMS) from Central Europe. The small-sample problem, especially acute for the NMS, is mitigated by using a Bayesian estimation procedure which combines information across countries. A novel identification scheme for small open economies is used. The estimated responses are quite similar across regions, but there is some evidence of more lagged, but ultimately stronger price responses in the NMS economies. This contradicts the common belief that monetary policy is less effective in post-transition economies, because of their lower financial development. NMS also have a probably lower sacrifice ratio, which is consistent with the predictions of both the imperfect information model of Lucas (1973) and the New-Keynesian model of Ball et al. (1988).

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Contents

1 Introduction

Prior to the creation of the Economic and Monetary Union (EMU), much research was devoted to the question of possible heterogeneity of responses to monetary shocks in the prospective member countries. The question was motivated by significant differences in some structural characteristics of the EMU economies and their possible implications for monetary transmission. If responses to monetary shocks are significantly heterogeneous, and the reasons of this heterogeneity do not disappear in the monetary union, conducting common monetary policy will be politically difficult (Dornbusch et al., 1998): The burden of disinflation will fall disproportionately on some countries, while other will have to accept higher than average inflation.[1]

Examples of papers discussing the impact of structural characteristics of the European economies on their monetary transmission are

[1]This problem is independent from the long debated question of whether current and potential future EMU member countries constitute an optimal currency area.

Dornbusch et al. (1998), Guiso et al. (1999), Mihov (2001), Ehrmann et al. (2003). These papers first look for indicators of interest sensitivity of output, size, health and structure of the banking sector, stock market capitalization and other, and relate them, by theoretical reasoning, to the strength of monetary transmission. The results of this type of analysis are often ambiguous, as different characteristics sometimes have conflicting implications, and their relative quantitative importance is unclear. The ultimate judgment has to come from macroeconomic data, usually analyzed with a Structural VAR technique. Papers in this line or research naturally fall into two categories: those that find significant and interpretable differences among the examined countries, and those that don't. Examples of the first group are Mihov (2001) and Ramaswamy and Slok (1998). Kieler and Saarenheimo (1998) and Ehrmann et al. (2003)/Mojon and Peersman (2001), among others, find that whatever asymmetries in monetary transmission might exist among EU countries, they are not strong enough to be robustly detected in the available data.

Now research along similar lines is being extended to the New Member States (NMS) from the Central and Eastern Europe, which joined the European Union in 2004, and which are legally obliged to adopt the euro some time afterwards. Examples are Anzuini and Levy (2004), Elbourne and de Haan (2005) and Creel and Levasseur (2005). Ganev et al. (2002) and Coricelli et al. (2005) contain surveys of this literature. Most authors agree that, mainly because of the small size of the financial markets, monetary policy in transition countries should have little effect, although its effectiveness is likely to be increasing with time, as the market economies in the region become more mature.

What has been missing so far, is an explicit comparison of responses to monetary shocks across the two regions: Central-Eastern and the Western Europe. Such comparison can be expected to be more meaningful and interesting than intra-regional comparisons performed so far, as the structural differences between these regions dwarf those within them. This paper fills this gap, by using a novel econometric technique, which allows to robustly estimate responses for both regions despite short data series and in a unified framework.

The comparison yields interesting results: First, monetary shocks in the NMS are associated with larger movements of the interest rate. Second, in spite of the structural differences between the regions, we find no support for the relative ineffectiveness of monetary policy in the NMS. Responses of output and prices are broadly similar, once controlling for the size of the interest rate shock. If anything, price responses are more lagged, but later more vigorous in the NMS and in the medium term these countries may be facing a more favorable sacrifice ratio.

We conclude that, on the one hand, the structural weakness of monetary transmission in the NMS is quantitatively less important than widely believed, possibly because it is compensated by a relative strength of the exchange rate channel, by more limited access to foreign financial markets and other discussed factors. On the other hand, prices appear to be more responsive, once aggregate demand is affected. This finding is consistent with greater volatility of aggregate demand and higher average inflation in the NMS, through well known mechanisms proposed in Lucas (1973) and Ball et al. (1988).

As regards the implications of these results for the adoption of the euro by the NMS, the common limitation of this and previous studies is that they are subject to the Lucas critique: responses to monetary shocks are likely to change after a further EMU expansion. Nevertheless, as argued in the above quoted papers, the empirical results based on past data provide stylized facts, which are a reasonable departure point for further speculations. Taken at face value, results of this paper downplay the structural weakness of monetary transmission as an argument against further EMU expansion.

The principal obstacle in the study of the former communist countries are the short available data series. To mitigate this problem, we perform a Bayesian estimation with the prior (called 'exchangeable prior'), which conveys the intuition that parameters of VAR models for individual countries are similar across the region, since all economies in the region are special cases of the same underlying economic model. This prior results in estimates which are shrunk towards a common mean. The Bayesian setup used here allows to formulate the problem in such a way, that both the degree of shrinking, and the weights of countries in the common regional mean are endogenous and optimal for the sample at hand. This guarantees the most efficient use of the scarce available data.

Applications of the estimation with the exchangeable prior in economics include Zellner and Hong (1989) and Canova and Marcet (1995). The former find that the exchangeable prior improves the out of sample forecasting ability in time series models, which has been also exploited in the forecasting time-varying VARs of Canova and Ciccarelli (2004). This finding suggests, that it should also increase reliability of a structural analysis. However, it has not been used for structural VARs, except for Canova (2005), which uses a different technique of working with a similar prior. The present paper adapts to VARs the formulation of Gelman et al. (1995), called Hierarchical Linear Model, which allows to avoid the specification of a subjective prior about the degree of similarity between units, but instead determines it solely from the data.

The second methodological novelty of this paper is the proposed identification of monetary shocks in a VAR. They are identified by assuming that they influence output and prices with at least one month lag, and that they involve a negative comovement of interest rate and exchange rate innovations. This is a combination of standard zero restrictions with more recently proposed sign restrictions.[2]

The structure of the paper is following: Section 2 discusses estimation of reduced form VARs for a panel of countries as a Hierarchical Linear Model, Section 3 describes identification of monetary shocks, Section 4 presents results and Section 5 contains conclusions. Details about data and estimation are in the appendix.

2 Estimation

VAR models contain many parameters, and their estimation with short samples, such as those available for the post-communist countries, results in wide error bands and point estimates which are very sensitive to small changes in sample or specification. The strategy employed here to obtain more robust results, is to analyze whole regions (first EMU, then NMS) jointly and exploit the intuition, that parameters of VAR models for individual countries are similar across the region, since all economies in the region are special cases of the same underlying economic model. However, we need to stop short of assuming that all slope coefficients are the same across countries and performing a standard panel estimation. This assumption would only be an approximation, and in dynamic model (such as a VAR) it could seriously distort the results (Pesaran and Smith (1995) show that it results in the inconsistency of the estimator).

The estimation procedure is Bayesian, and uses the Hierarchical Linear Model of Gelman et al. (1995).[3] The idea of similarity is specified as a Normal prior for each country's coefficients, which is centered at the mean which is common for all the region (an exchangeable prior). This prior causes the coefficients to be shrunk towards the common mean. The second stage of the hierarchy consists of the 'hyperprior' about the prior parameters: common mean and the

[2]Other sign restrictions have been used in Faust (1998), Uhlig (2001), Canova and De Nicoló (2002) and Kieler and Saarenheimo (1998). Unlike in the mentioned papers, the combination of zero and sign restrictions enables one to find the desired factorizations of the error variance matrix analytically, avoiding a numerical search procedure.

[3]Classical estimators for heterogeneous panels exist, but are much less efficient: Monte Carlo study in Hsiao et al. (1999) shows that in small samples they perform worse than a variant of Bayesian estimator with the exchangeable prior.

variance of country coefficients around the common mean ('hypervariance'). The Hierarchical Linear Model allows the priors in the second stage of the hierarchy to be noninformative, and therefore the posterior common mean and hypervariance are determined optimally only from the data. Intuitively, more different and more tightly estimated country coefficients increase the posterior probability of large values of the hypervariance. When country coefficients are more similar, or if they differ, but have larger error bounds, hypervariance is more likely to be smaller. Country models which are more tightly estimated receive more weight in the posterior common mean, relative to countries whose estimates are imprecise.

Below, we first distinguish between parameters which are likely to be similar across countries, and those which need not be. So, we apply the exchangeable prior to parameters determining dynamic interrelationships between the endogenous variables, and reactions to common exogenous variables. We specify a noninformative prior for parameters of exogenous variables which are not present for all countries, and for constant terms, which implies that we have country 'fixed effects'.

The following two subsections specify the above prior in the panel VAR setup: first the overall framework, and then the parametrization of the hypervariance. The computation of the posterior is explained in the appendix.

2.1 Panel of VARs as a Hierarchical Linear Model

In what follows, vectors are denoted by lowercase, matrices by uppercase bold symbols, $i = 1 \dots I$ denotes countries, $j = 1 \dots J$ denotes endogenous variables in a VAR, $k = 1 \dots K$ denotes the common right-hand-side variables in the reduced form VAR, $l = 1 \dots L$ denotes lags, $m = 1 \dots M_i$ denotes country specific exogenous variables in the VAR, $t = 1 \dots T_i$ denotes time periods.

For each country in the panel we consider a reduced form VAR model of the form:

$$\boldsymbol{y}_{it} = \sum_{l=1}^{L} \boldsymbol{B}'_{il} \boldsymbol{y}_{i(t-l)} + \boldsymbol{\Delta}'_i \boldsymbol{w}_t + \boldsymbol{\Gamma}'_i \boldsymbol{z}_{it} + \boldsymbol{u}_{it} \tag{1}$$

$\boldsymbol{y}_{it}$ is a vector of J endogenous variables and $\boldsymbol{w}_t$ is a vector of those exogenous variables which are common across countries. We will specify an exchangeable prior about the coefficients of $\boldsymbol{y}_{i(t-l)}$ and $\boldsymbol{w}_t$. The prior will be uninformative for variables in $\boldsymbol{z}_{it}$, which include country specific constant terms and variables which are included for some, but not all countries. Vector $\boldsymbol{u}_{it}$ contains VAR innovations which are i.i.d. $N(0, \boldsymbol{\Sigma}_i)$.

We gather the variables to which the exchangeable prior applies in a vector $\boldsymbol{x}_{it} = \left[\boldsymbol{y}'_{i(t-1)} \cdots \boldsymbol{y}'_{i(t-L)}, \boldsymbol{w}'_t\right]'$. Stacking vertically $\boldsymbol{y}'_{it}, \boldsymbol{x}'_{it}, \boldsymbol{w}'_t$ for all t we obtain the model in terms of data matrices:

$$\boldsymbol{Y}_i = \boldsymbol{X}_i \boldsymbol{B}_i + \boldsymbol{Z}_i \boldsymbol{\Gamma}_i + \boldsymbol{U}_i \tag{2}$$

where $\boldsymbol{Y}_i$ and $\boldsymbol{U}_i$ are $T_i \times J$, $\boldsymbol{X}_i$ are $T_i \times K$, $\boldsymbol{B}_i$ are $K \times J$, $\boldsymbol{Z}_i$ are $T_i \times M_i$ and $\boldsymbol{\Gamma}_i$ are $M_i \times J$. We have: $K = JL + W$, where W is the length of the $\boldsymbol{w}_t$ vector. The coefficient matrix $\boldsymbol{B}_i$ is related to coefficients of (1) by: $\boldsymbol{B}_i = [\boldsymbol{B}'_{i1}, \ldots, \boldsymbol{B}'_{iL}, \boldsymbol{\Delta}']'$.

Let $\boldsymbol{y}_i = \operatorname{vec} \boldsymbol{Y}_i, \boldsymbol{\beta}_i = \operatorname{vec} \boldsymbol{B}_i, \boldsymbol{\gamma}_i = \operatorname{vec} \boldsymbol{\Gamma}_i$.

The statistical model generating the data is assumed to be following:

Likelihood for country i:

$$p(\boldsymbol{y}_i | \boldsymbol{\beta}_i, \boldsymbol{\gamma}_i, \Sigma_i) = \mathrm{N}((\boldsymbol{I}_J \otimes \boldsymbol{X}_i)\boldsymbol{\beta}_i + (\boldsymbol{I}_J \otimes \boldsymbol{Z}_i)\boldsymbol{\gamma}_i, \boldsymbol{\Sigma}_i \otimes \boldsymbol{I}_{T_i}) \tag{3}$$

Country coefficients on the variables in $\boldsymbol{X}_i$ are assumed to be drawn from a normal distribution with a common mean $\bar{\boldsymbol{\beta}}$:

$$p(\boldsymbol{\beta}_i | \bar{\boldsymbol{\beta}}, \lambda, \boldsymbol{L}_i) = \mathrm{N}(\bar{\boldsymbol{\beta}}, \lambda \boldsymbol{L}_i) \tag{4}$$

where λ is an overall prior tightness parameter and $\boldsymbol{L}_i$ is a known, fixed matrix whose construction is discussed below.

Prior for $\bar{\boldsymbol{\beta}}$ and $\boldsymbol{\gamma}_i$ is uninformative, uniform on the real line:

$$p(\bar{\boldsymbol{\beta}}) \propto p(\boldsymbol{\gamma}_i) \propto 1 \tag{5}$$

Alternatively, one could use some informative prior for $\bar{\boldsymbol{\beta}}$, e.g. the Minnesota prior, but, as discussed in Gelman et al. (1995), this is not necessary for the estimation problem to be well posed, and, in order to keep things simple, we do not pursue this possibility here. We also use the standard diffuse prior for the error variance:

$$p(\boldsymbol{\Sigma}_i) \propto |\boldsymbol{\Sigma}_i|^{-\frac{1}{2}(J+1)} \tag{6}$$

Finally, the prior for the overall tightness parameter λ is:

$$p(\lambda | s, v) = \mathrm{IG}_2(s, v) \propto \lambda^{-\frac{v+2}{2}} \exp\left(-\frac{1}{2}\frac{s}{\lambda}\right) \tag{7}$$

where IG_2 denotes the inverted gamma-2 distribution, while s and v are known parameters. For $s > 0$ and $v > 0$ this is a proper, informative prior while $s = 0$ and $v = -2$ results in an improper, noninformative prior.

The model in (3)-(7) defines the structure advocated in the introduction: the countries' dynamic models of variables in $\boldsymbol{Y}_i$ (and possibly some exogenous controls in $\boldsymbol{W}$) are special cases of the unknown underlying model defined by $\bar{\boldsymbol{\beta}}$. Variables in $\boldsymbol{Z}_i$ are those, for which the exchangeable prior would not be reasonable, primarily the country specific constant terms.

The functional form of the prior: combination of normal, uniform, inverted gamma and a degenerate inverted Wishart (for $\boldsymbol{\Sigma}_i$) densities is standard, motivated by computational convenience, so that the prior is conditionally conjugate. The posterior density of the parameters of the model is computed from the Bayes theorem, as a normalized product of the likelihood and the prior. The conditional conjugacy of the prior means that all conditional posterior densities are also normal, inverted gamma and inverted Wishart, which enables convenient numerical analysis of the posterior with the Gibbs sampler.[4]

2.2 Specification of the prior variance

The parametrization of the prior variance for $\boldsymbol{\beta}_i$ is inspired by prior variances in Litterman (1986) and Sims and Zha (1998): it is assumed to be diagonal, with the terms of the form:

$$\frac{\lambda\sigma_{ij}^2}{\sigma_{ik}^2} \tag{8}$$

The ratio of variances reflects the scaling of the variables. As in the above papers, σ^2s are computed as error variances from univariate autoregressions of the variables in question. Therefore, the $\boldsymbol{L}_i$ is computed as:

$$\boldsymbol{L}_i = \mathrm{diag}(\sigma_{ij}^2) \otimes \mathrm{diag}(\frac{1}{\sigma_{ik}^2}) \tag{9}$$

The parameter λ determines the overall tightness of the exchangeable prior. $\lambda = 0$ results in full pooling of information across countries and implies a panel VAR estimation, where all country VAR models are assumed to be identical. On the other hand, as λ grows, country models are allowed to differ more, and become similar to the respective single country estimates. Since the value of λ is unknown, a (possibly uninformative) prior distribution is assumed for it and a posterior distribution is obtained. The reported results are integrated over this posterior distribution. If the posterior inferences conditional on particular values of λ differ in an economically meaningful way, odds ratios for alternative ranges of λ can be computed.

[4]See appendix for details. Gibbs sampler for a similar univariate problem is discussed in more detail in Gelman et al. (1995).

The use of noninformative priors carries the danger of obtaining an improper posterior and rendering the whole problem ill-defined. It is known, that in a hierarchical linear model like (3)-(7) the use of the usual noninformative prior for a variance parameter:

$$p(\lambda) \propto \frac{1}{\lambda} \tag{10}$$

(which obtains when $s = 0$ and $v = 0$) results in an improper posterior (see Hobert and Casella, 1996; Gelman et al., 1995). In this case, the marginal posterior for λ behaves like $1/\lambda$ close to the origin and is not integrable. However, Theorem 1 in Hobert and Casella (1996, p.1464), proved in a similar setup, suggests that the posterior is proper when $s = 0$ and $v = -2$, which corresponds to:

$$p(\lambda) \propto 1 \tag{11}$$

3 Identification of monetary shocks

Identification of the structural model assumes a small open economy with exchange rates flexible enough to react immediately to monetary policy, and monetary policy reacting immediately to the movements of the exchange rate. This assumption remains valid also in presence of managed exchange rates with target bands, like the ERM in the European Union, and some arrangements in the NMS, as long as the rate is not effectively fixed. It is a known empirical regularity (confirmed in the robustness analysis for this paper) that for countries other than the USA, identification schemes that do not allow for immediate response of the exchange rate to the interest rate, and vice versa, produce a 'price puzzle', i.e. an initially positive response of prices to monetary tightening (for more on this subject see e.g. Kim and Roubini (2000)).

The endogenous variables in the VARs are: output, consumer prices, short term interest rate and the exchange rate in national currency units per foreign currency unit, all measured at monthly frequency. No money aggregate is included: It is assumed that the central banks target short term interest rates, and adjust monetary aggregates consistently with this objective. In this setup, interest rates reflect only money supply decisions, and fluctuations of monetary aggregates carry additionally information about money demand. Since identification of money demand is beyond scope of this paper, we conserve degrees of freedom and do not include money aggregates in the specification.

In order not to confuse domestic monetary shocks with the central banks' responses to external developments, specifications include several foreign variables which are treated as exogenous. World developments are captured by current and lagged US Federal Funds Rate, oil and commodity prices. The status of Germany as both regions' locomotive is reflected by including, for all countries, current and lagged German interest rate and two lags German industrial production.

As usually in the identified VAR literature, it is assumed that structural shocks are orthogonal, and thus the covariance matrix of the VAR residuals conveys information about the coefficients of the contemporaneous relationships between endogenous variables. The relationship between the vector of structural shocks $\boldsymbol{\upsilon}_{it}$ and the vector of VAR innovations $\boldsymbol{u}_{it}$ is following:

$$\boldsymbol{G}_i \boldsymbol{\upsilon}_{it} = \boldsymbol{u}_{it} \tag{12}$$

where $\mathrm{var}(\boldsymbol{\upsilon}_{it}) = \boldsymbol{I}_J$ (identity matrix of order J) and $\mathrm{var}(\boldsymbol{u}_{it}) = \boldsymbol{\Sigma}_i = \boldsymbol{G}_i \boldsymbol{G}_i'$. Therefore, the identification involves finding a factorization $\boldsymbol{G}_i$ of the residual covariance matrix that complies with the identifying restrictions.

The identification restrictions[5] adopted to pin down the monetary shock are following:

1. Output and prices do not respond immediately to the monetary policy shock
2. The monetary policy shock is the one which involves a negative comovement of the interest rate and the exchange rate on impact, i.e. interest rate rise is accompanied by exchange rate appreciation.

The remaining shocks are not identified and the triangular form of the upper left block of the matrix reflects a normalization, which has no effect on the impulse responses to the monetary policy shock. The identification restrictions are summarized in the scheme below:

$$\begin{pmatrix} + & 0 & 0 & 0 \\ \bullet & + & 0 & 0 \\ \bullet & \bullet & + & + \\ \bullet & \bullet & - & + \end{pmatrix} \begin{pmatrix} \upsilon_{it1} \\ \upsilon_{it2} \\ \widehat{\upsilon_{it3}} \\ \upsilon_{it4} \end{pmatrix} = \begin{pmatrix} u_{it1} \\ u_{it2} \\ u_{it3} \\ u_{it4} \end{pmatrix} \begin{array}{l} \leftarrow \text{output innovation} \\ \leftarrow \text{price innovation} \\ \leftarrow \text{interest rate innovation} \\ \leftarrow \text{exchange rate innovation} \end{array} \tag{13}$$

where + denote coefficients that are constrained to be positive, 0 - zero restrictions and $\bullet$ - unconstrained coefficients. $\widehat{\upsilon_{it3}}$ is the monetary policy shock. Factorizations satisfying (13) are obtained by rotating the bottom-right block of the Choleski factor of the residual covariance matrix - see the appendix for details.

[5]Restrictions are imposed here on the impulse responses in the first period (inverse of the matrix of structural coefficients), and not on the structural coefficients.

The zero restrictions applied here are standard ones, used in Leeper et al. (1996), Kim (1999), Kim and Roubini (2000) and other papers. The strategy followed in these papers is to impose some more zero restrictions: to assume that monetary authorities don't react immediately to output and price developments. Then the model becomes (over)identified and can be estimated by maximum likelihood. It is then expected, that resulting estimates of the relationships between interest rate and exchange rate are 'reasonable', i.e. as in assumption 2 above. When multiple maxima of the likelihood function exist, the ones satisfying the implicit sign restriction are chosen, which is justified by a Bayesian reasoning (Kim, 1999, footnote 13, p.395).

However, the zero restrictions on the authorities' response to output and prices may not hold exactly: While it is true that official data on output and prices are compiled with delay, it could be argued that the central bankers can have access to quick business community surveys, price surveys, and certainly are able to identify quickly major shocks, like big strikes or floods. The sign restrictions proposed here are an attractive alternative.

4 Data and Samples

We analyze two panels of countries: five euro-area countries (EMU5) and four New Member States (NMS4) from the Central and Eastern Europe.[6] The goal of the paper is to compare the two regions within one unified framework. On the one hand, we want to make the analysis representative and include as many countries as possible. On the other hand, to ensure comparability of the results across regions, we want to maintain the same specifications and identification schemes. This leads to the choice of the assumption of open economy with a flexible (i.e. not fixed) exchange rate, which maximizes the sizes of both panels.

The EMU5 panel consists of Finland, France, Germany, Italy and Spain. Austria, Belgium and Netherlands were excluded because of their quasi-fixed exchange rate against the D-Mark[7], and Ireland because of the lack of monthly CPI data. In Greece, interbank interest rates are not available before 1998, only the Central bank rate is reported. Portugal proved to be an outlier because of dramatic swings

[6] Romania and Bulgaria were considered as potential members of this group, although strictly speaking they are only expected to join the EU in 2007. On the other hand, we did not consider Malta and Cyprus, which are both economically and geographically distinct.

[7] We follow Mojon and Peersman (2001) who also consider them separately for this reason.

of its interest rates in the wake of the 1992 European Monetary System crisis. Therefore, both countries are considered only in robustness checks.

The NMS4 panel consists of Czech Republic, Hungary, Poland and Slovenia. Bulgaria and the Baltic countries were excluded because they had currency boards. In Romania and Slovakia market interest rates vary widely and, for much of the sample, independently of the central bank interest rates. This suggests that the standard model of monetary management, which underlies this analysis, where the central bank manages market interest rates by setting its instrument interest rate, has not been firmly in place. Another possibility is that these countries experienced big shocks to money demand which were not accommodated by the central bank. In either case, the identification of monetary policy shocks adopted here might not be appropriate. Therefore, Slovakia and Romania are considered only in robustness checks.

The sample periods for the NMS countries span the second half of 1990-s up to second half of 2004 and differ for each country, depending on when the post-transition exchange rate control was relaxed. The information on the chronology of the exchange rate regimes was taken from Anzuini and Levy (2004), Table 8 and from Ganev et al. (2002), Section 3 and Table 1. The details about the samples are in the appendix. For the euro-area countries we consider samples of similar lengths as for the NMS countries, covering second half of the 1980's up to 1998 (the start of the EMU). As a summary, we present here the results for a longer sample spanning 1985 (1) - 1998 (12). These results are similar to those for the shorter samples, but free of some features which were deemed not robust.

The data is monthly. The endogenous variables: output, prices, interest rates and exchange rates are measured respectively by log of the Index of Industrial Production (IIP), log of the Consumer Price Index (CPI), short term market interest rate (r-mkt) and the log of the exchange rate in national currency units per SDR. The SDR is a standard basked of main currencies, and it provided an intermediate choice between a US dollar exchange rate, which is used often, but may be influenced by some US specific events, and country specific baskets of most relevant currencies, which are less comparable. Most data (as well as those for the exogenous variables: Federal Funds Rate, oil prices, non-fuel commodity prices, German interbank interest rate and German industrial production) are taken from the IMF IFS database, and some from the Eurostat. See the appendix for details.

The interest rate of 0.1 corresponds to 10% (1000 basis points). The variables other than the interest rate are logs of indexes that assume the value 1 in December 1995. The basic specification contains six lags of the endogenous variables and lags zero and one of the exogenous variables. The exception is German industrial production (included as an exogenous variable for countries other than Germany): it is assumed that foreign central banks observe it with a lag, so lags one and two are included. Shorter lag length of the exogenous variables is chosen to conserve the degrees of freedom.

5 Results

We approximate the posterior distribution of the estimated coefficients using the Gibbs sampler[8] and compute impulse responses over the horizon of 40 months.

5.1 'Mean' impulse responses for the regions

Figure 1 presents the 5th, 50th and 95th percentiles of the posterior distribution of impulse responses to a one standard deviation monetary shock, implied by the 'mean' model for each panel ($\bar{\boldsymbol{\beta}}$).

The immediate (period 0) responses of all variables reflect the identifying assumptions: 1) in the month of the shock the output and prices are unaffected and 2) the interest rate raises and the exchange rate falls (appreciates). The uncertainty band for the impact behavior of the interest rate and exchange rate ranges from the 5th to the 95th percentile of all the range where the sign restrictions are satisfied.

The identified MP shock is associated with a median interest rate increase of 40 basis points in the EMU5 and almost 80 bp in the NMS4. The median initial appreciation is respectively 1% and 1.5%. The interest rate increase is reversed after about one year, while the appreciation persists for about 2 years. The economies respond with a transitory output decline and a possibly permanent reduction of the price level.

The finding that interest rate movements in the NMS4 were on average twice stronger than those in the EMU5 is not surprising: First, output growth rates, inflation levels and interest rates in the post-communist countries tended to be higher than in the Western Europe, which was likely to generate higher variance of shocks. Second, the central banks in the region may have believed that, because of the low financial depth in these countries, monetary policy is not very effective

[8]We generate 2000 draws from the posterior, after discarding the initial 4000 burn-in draws.

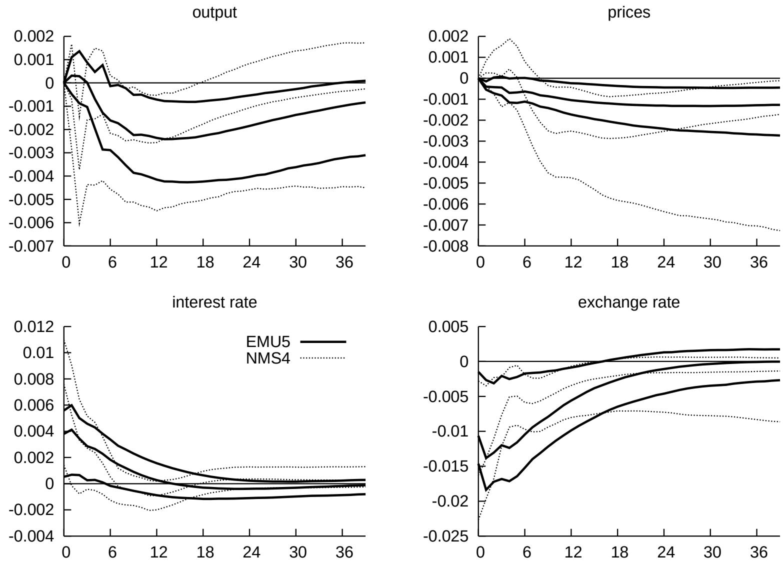

Figure 1: Mean impulse responses for the EMU5 and the NMS4: median, 5th and 95th percentiles of the posterior distribution

there (Ganev et al., 2002), and therefore their policy actions require more vigorous interest rate movements.

A look at the variance decompositions (figure 2) suggests that the higher monetary shocks in NMS4 contributed importantly to the variability of output and especially prices in that region. According to the median of the distribution, around 10% of the variability of output is attributed to monetary shocks in both countries. In case of prices, the shares are about 20% in the EMU5 and 30% in NMS4. As usual in the VAR literature, there is a wide uncertainty about the exact figures. The posterior distributions for output overlap significantly, but monetary shocks are more likely to be responsible for a greater share of variance of prices in the NMS4 than in the EMU5.

In order to control for the different size of shocks we standardize the impulse responses, to make them correspond to the same size of the interest rate shock in the first month. We take the EMU5 shock as a benchmark. Therefore, we scale down the NMS4 impulse responses, so that the average of the impact and first month response of the interest rate is the same as in the EMU5 panel. (We take the average of impact and first lag effects in order to neutralize partly the different dynamics of interest rates in both panels). Figure 3 presents the standardized output and price responses, and figure 4 shows the probability that the standardized NMS4 response is weaker than that of the EMU5

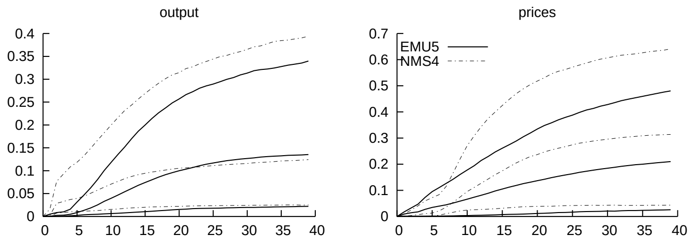

Figure 2: Share of monetary shocks in the variance decompositions for the NMS4 and the EMU5 panels: median, 5th and 95th percentiles of the posterior distributions

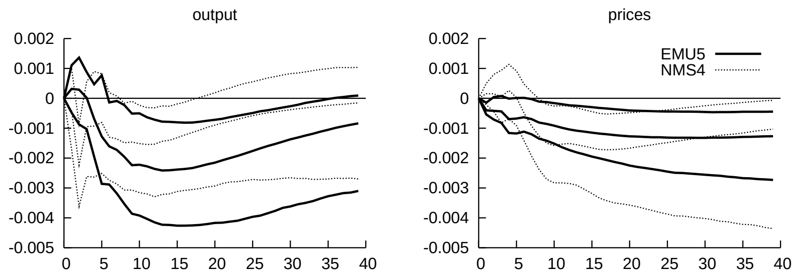

Figure 3: Mean impulse responses of output and prices in the EMU5 and the NMS4 corresponding to the same size interest shock in the first month

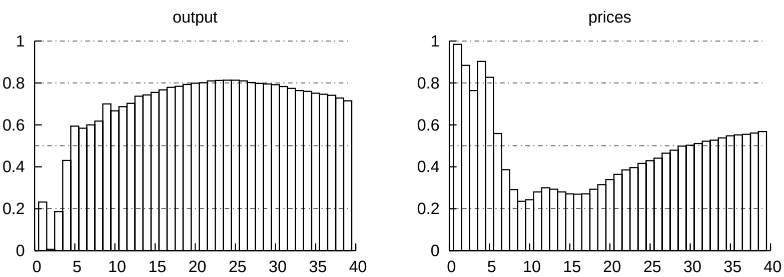

Figure 4: Posterior probability that the standardized response of the NMS4 (from figure 3) panel is weaker than that of the EMU5 panel

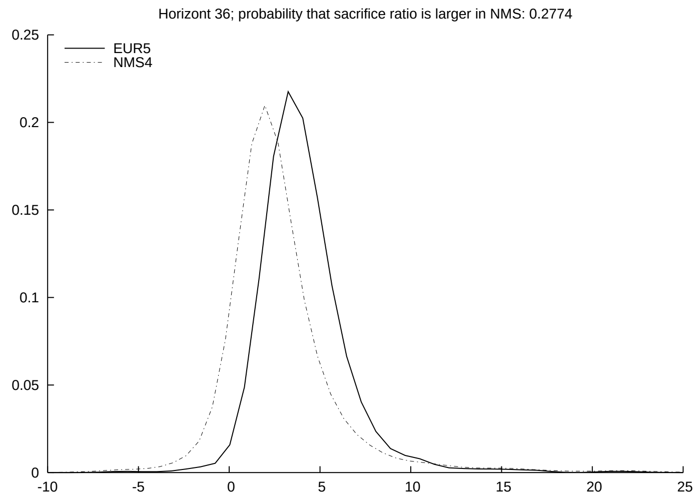

Figure 5: Posterior distribution of the sacrifice ratio (36 months horizon)

panel.

The standardized impulse responses of both regions are quite similar, their 90% probability regions are mostly overlapping. In particular, as in the variance decompositions, we don't find a straightforward confirmation of the belief that the monetary policy is less effective in the NMS4, because of the low financial depth of these countries. It is true that output, in the medium term, responds rather weakly. Also, the price responses are more lagged in the NMS4, and need over 6 months to become significant. However, after this delay, they react more vigorously than in the EMU5.

The observation that output responses tend to be weak, while those of prices quite strong, suggests that the sacrifice ratio facing the NMS4 central bankers could be lower. This possibility is examined closer by comparing the posterior distributions of the sacrifice ratios for both panels, calculated as in Cecchetti and Rich (2001). Their comparison suggests that the sacrifice ratio in the NMS4 might indeed be lower, although the posterior probability of this statement is only 72% with the 36 months horizon.[9]

Central banks in the NMS have shorter track records, and probably enjoy less credibility than their Western European counterparts. If this is the case, their monetary policy has less impact on agent's

[9]This probability exceeds 80% with shorter horizons. The reason why the sacrifice ratios become more similar for longer horizons is that the price decline gets reversed in the long run. The very long run behavior of prices may be, however, less reliably estimated than the medium run behavior.

expectations, which results in longer lags in the response of prices.

The longer lags of price response in the NMS4 panel can be complementarily explained by models of learning in an evolving setup (see e.g. Evans and Honkapohja, 2001). Initially, because of undeveloped financial markets, Central Bank discount rates mattered very little for inflation. As financial markets and credit activity grows, so does the importance of the interest rates, and agents revise estimates of their impact on inflation, and readjust their expectations. However, the nature of learning under uncertainty implies that they adapt their models only partially. As a result, the impact of the Central Bank policies is always initially underestimated, and takes full effect with a delay, after expectations adjust.

However, the most unexpected aspect of the results in figures 3 and 4 is that, after a few months' lag, monetary policy seems to have quite a strong impact on prices in the NMS. This is in spite of the fact, that in the NMS4 indicators of the size of the financial systems, such as the ratio of financial assets/liabilities, or stock market capitalization, to GDP, are lower by a factor of 2 to 5 in comparison with the euro-area average (see Anzuini and Levy, 2004). Apparently, we need to go beyond the simple rule of thumb, that monetary policy is less effective in less financially developed countries, when comparing the Central-Eastern and the Western Europe.

The possibility of lower sacrifice ratios in the NMS suggest other theories, which may be relevant here: The NMS4 have more volatile and, on average, higher inflation rates. Lucas (1973) argues that in an imperfect information model, in countries where aggregate demand fluctuates more, agents adjust their prices more than their outputs. Ball et al. (1988)'s reasoning is, that in higher inflation countries, agents need to adjust their prices more often, and so there is less stickiness. Both models imply, that in such countries aggregate supply curve is steeper, and, consequently, the sacrifice ratio lower, and this is confirmed in their cross-country studies.

There are reasons to believe, that the NMS economies should be responsive to monetary policy: First, the exchange rate channel might be stronger. The NMS4 economies are more open. Moreover, they have less established brands where monopolistic competition is important, so their exports can be more sensitive to exchange rates.

Second, with less developed financial markets, agents may find it more difficult to hedge against the monetary policy changes. Anzuini and Levy (2004) find that the agents in the NMS4 have mostly unhedged foreign debt, and are very exposed to the foreign exchange rate risk. Third, firms in NMS4 find it harder to obtain credit abroad, when domestic credit conditions are tight.

Fourth, even with small institutional financial markets, the prevailing interest rates might still matter for economic decisions and transactions, such as the trade credits, or reinvestment of profits. The NMS4 have a high volume of trade credit (see Anzuini and Levy, 2004) and, since they are catching-up economies, they have more investments compared with GDP than the euro area (Suppel, 2003), which is mostly financed by reinvesting profits.

Summarizing, the story behind figure 3 could be following: In the NMS, the policy tightening by the central bank is less credible, and so has little initial effect on expectations, and thus on pricing decisions. The effect of higher interest rates on output is weaker than in the EMU, because of the small financial markets, but not much weaker, in light of the arguments listed above. In addition, this modest aggregate demand contraction is translated more efficiently into prices, because the last are less sticky.

5.2 Heterogeneity within EMU and NMS

Heterogeneity of the panels is best reflected by the posterior distribution of the overall tightness parameter λ. As follows from the statistical structure in equations (3)-(7), probability mass concentrated on low values of λ means that the posterior distributions of country VAR models are close to each other. When, on the contrary, the posterior distributions of parameters for individual countries tend to differ, higher values of the λ are more likely, and get more posterior support.

In the two panels estimated here, the likelihood turns out to be very informative about λ and the support of the posterior distribution is very narrow. This can be assessed by comparing the shape of the impulse responses for e.g. the 5th and the 95th centiles of the simulated distribution of λ: the graphs are almost indistinguishable. Therefore, the data favors a certain intermediate amount of cross-country information pooling, preferring it both to complete homogeneity, and to independence of the individual country models.

The posterior density of the overall tightness parameter λ in each of the panels is presented in figure 6. The NMS4 panel turns out to be more heterogeneous.

Figures 7 and 8 present impulse responses to a monetary shock for each of the analyzed countries. The considered samples are short: for individual countries, they imply only 3.1 observations per estimated parameter in the NMS4 and 2.4 for the Czech Republic, and maximally no more than 5 in the euro area. For the NMS4 countries these are maximum samples under flexible exchange rate regime. In the standard, individual country estimation, results based on such short data

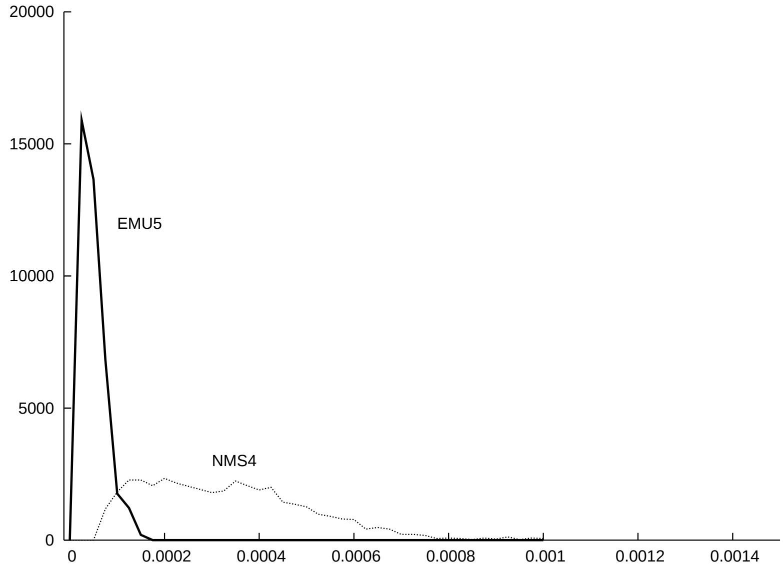

Figure 6: Simulated posterior distribution of λ for the euro-area panel and for the NMS4 panel

would have to be treated with much caution. The results presented here are more reliable, because in the computation of the posterior, the country data is optimally augmented with the information for all other countries in the panel.

Consistently with what the posterior distribution of the tightness parameter λ is suggesting, the impulse responses for EMU5 look more homogeneous than those for NMS4. In the EMU5 panel, Germany is the clear outlier, with the weakest monetary shocks (around 15 bp) and much uncertainty about output and price responses, which renders them insignificant. In Finland, France and Italy the median interest rate shock is around 40 bp, and 50 bp in Spain, which becomes insignificant after about one year. Exchange rate appreciates similarly, by 1-1.5% in all countries. The peak of output response comes after one year, at -0.2% in France, Italy and Spain, and in Finland output reaction is strongest, at -0.3%. Prices fall gradually by about 0.1% in Finland and France, and by about 0.2% in Italy and Spain, which suggests that the latter two may have lower sacrifice ratios, which could be linked with their higher inflation rates

In the NMS4 panel the median interest rate tightening is strongest in Poland (about 90 bp), somewhat weaker in Czech Republic and Slovenia (around 60 bp), and weakest in Hungary, at 40 bp, which is similar as in the EUR5 countries. Exchange rate appreciates by roughly as much as in the EUR5 countries, possibly somewhat stronger (the comparison is blurred by different dynamics: in the NMS the exchange rate appreciates most on impact, while in the EMU the peak response comes after one month), except in Czech Republic, where the appreciation is clearly stronger, at more than 2%. Output falls by about 0.2% in all countries, except Czech Republic where the contraction reaches -0.3%. Prices fall by about 0.2% in Hungary, 0.25% in Slovenia and 0.3% in Poland, and stabilize at the new level after about one year. Czech Republic is an exception, because prices fall by almost 0.4% within the first year, but seem to increase back afterwards. Overall, in Poland, which is an inflation targeter, with largely uncontrolled exchange rate, monetary policy is most volatile, and has most impact. Hungary, where the exchange rate is kept within the narrowest band (among the analyzed countries), sees least impact of monetary policy. This is consistent with the exchange rate channel being crucial for monetary transmission in the region.

5.3 Robustness

The results have been checked for robustness to changing the country composition of the panels, sample periods, specification of the VARs, control variables and the identification scheme. The main conclusions go through under those experiments.

Given the short sample size, and the overparametrization of the VAR model, individual countries results are sensitive to changes in the sample period. However, the conclusions for the mean model are generally robust to changing the sample or removing any of the countries from the panel. As mentioned before, market interest rates for Slovakia and Romania hardly follow the central bank interest rates, so these countries were skipped in the basic estimation. When they are included, results for Slovakia are odd since output responds positively, but mean responses are unaffected. Responses for Romania are reasonable, but the average size of the interest rate shock is very large, more than 500 basis points (Romanian market interest rates are very volatile) and this increases the size of the mean interest rate shock to about 250 bp. Other impulse responses are barely affected. Similar situation concerns Portugal, which is an outlier in the euro-area panel, because the unusual swings of its interest rates in the 1992/1993. During the EMS crisis, the Portuguese Central Bank was for some time fending off speculative attacks on the escudo by interest rates increases, which needed to be more dramatic than in other countries affected by the crisis, partly because of the relatively small volume of the market. The resulting interest rate swings were per-

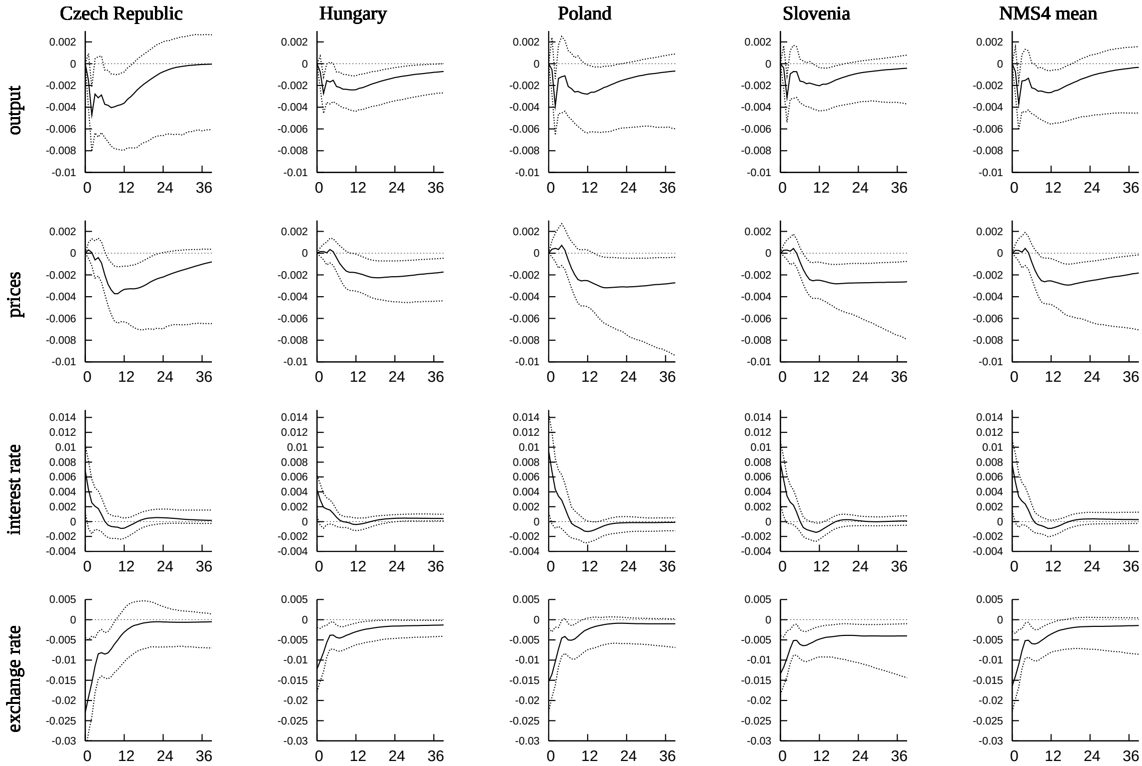

Figure 7: Impulse responses in the EMU5 panel

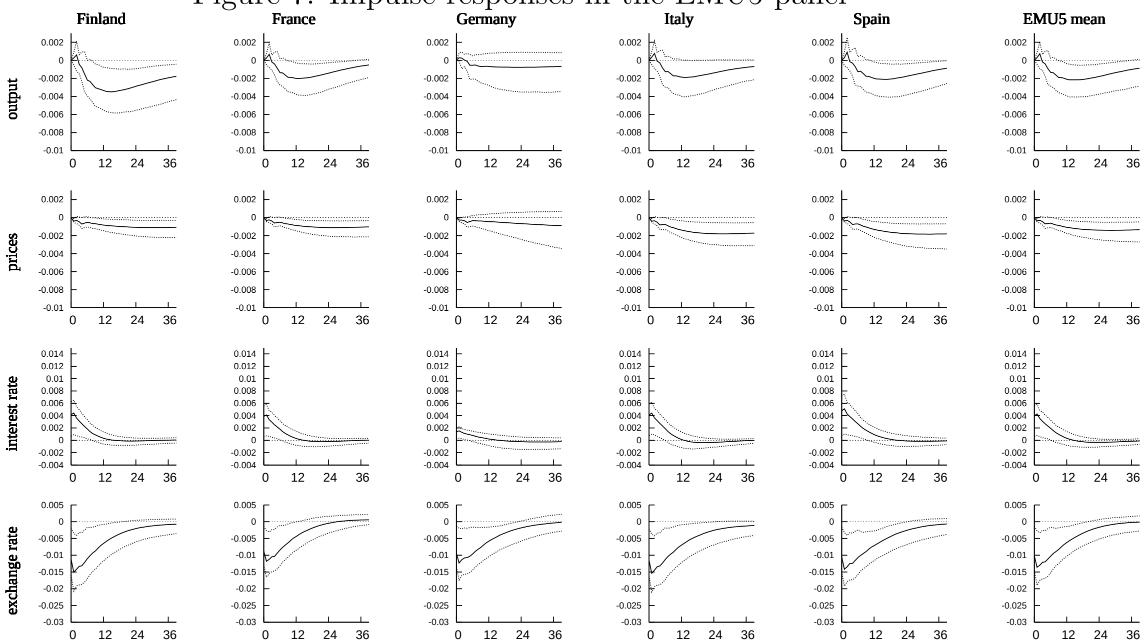

Figure 8: Impulse responses in the NMS4 panel

ceived as temporary anomalies and did not have usual real effects. Therefore, including Portugal in the panel results in a larger size of the interest rate impulses, while barely affecting other variables.

For the NMS4 panel, there is little scope for varying the sample sizes, as they are already very short, but the results are similar for one year shorter samples. For the euro-area panel, we try 9 year samples (roughly equal to the typical sample size for the NMS4 countries) spanning the 1980s until 1998. 9 year samples between 1985 and 1998 produce similar results, which are best characterized by those for the longer 1985-1998 sample. For 9 year samples starting in the first half of the 1980s impulse responses are often insignificant. Apparently, the model lacks features necessary to explain data from early 1980s, but we assume that this does not hamper the comparison performed in this paper.

In another set of experiments, Central Bank interest rates were used instead of the market interest rates. Most of the VAR studies assume, as does this paper, that the Central Bank targets interbank market rates, and the transmission from this intermediate target to the economy is studied. When central bank interest rates are used, comparisons with both the basic NMS4 panel, and the full one including Slovakia and Romania, lead to similar conclusions.

Replacing the SDR exchange rate with the USD exchange rate makes almost no difference for both sets of countries.

If the model estimated without any control variables (only constant terms and lags of endogenous variables), the output and price responses in both panels become more delayed, but deeper in the medium run, compared with the case with all controls included. The comparison of the responses, however, remains unaffected. The contemporaneous German interest rate makes most of the difference, adding further lags does not make a difference. Similarly, adding more lags or changing the set of world control variables often affects individual country responses, but the comparison of the mean responses are unaffected. After some experimentation, it was deemed best to stick with the maximal set of controls and avoid the risk of misspecifying the model. It is hoped that the estimation procedure can 'average out' the noise introduced by possibly excessive number of control variables.

Finally, the analysis was repeated with the recursive identification (i.e. assuming that monetary authorities react with a lag to the exchange rate developments). This identification results in a 'price puzzle': price response is initially positive. For the euro area the 'price puzzle' is not very significant, lasts only about one year, output responds negatively, but the exchange rate seems to depreciate. For the NMS4 panel the positive price response is very strong and output also responds positively.

6 Conclusions

This paper makes one of the first systematic comparisons of the responses to monetary shocks in Western Europe and in the New Member States of the EU. The responses of the NMS4 (Czech Republic, Hungary, Poland and Slovenia) turn out to be broadly similar to those in the EMU, but with interesting differences (albeit estimated with significant uncertainty): monetary shocks tend to be stronger, and generate a more delayed, but strong price level response, possibly at lower output cost.

These results suggest that, when considering the differences between the Central-Eastern and the Western Europe, we need to go beyond the simple rule of thumb, that monetary policy is less effective in less financially developed countries. Some of the structural features of the NMS financial systems (less possibilities for hedging, harder access to foreign financial markets), and their export orientation (strong exchange rate channel), may be amplifying the effects of monetary shocks on aggregate demand. Strong effect on prices, with possibly lower sacrifice ratios, that we find in the NMS, are consistent with the findings of Lucas (1973) and Ball et al. (1988), that in economies where the aggregate demand fluctuates more, and inflation is higher, the aggregate supply curve tends to be steeper, and prices less sticky.

Conclusions from the empirical analysis of this paper for the prospects of the EMU accession are not obvious, because of the Lucas critique. VARs are useful for establishing stylized facts about past monetary policy effects, but we don't know to what extent monetary transmission will change in the wake of the EMU accession. Several mechanisms may be at play:

After joining the EMU, the exchange rate channel, which is important for the NMS4 currently, will largely disappear, and domestic firms will gain easier access to wider financial markets. This will considerably weaken the responsiveness of the NMS to monetary policy. The result may be similar to the scenario predicted by Dornbush and observed within the present EMU: that interest rates that are optimal for the European core are too low for the peripheral, poorer and faster growing countries. As a result, the peripheries observe persistently higher inflation rates and may end up developing financial and property markets bubbles.

Overall, however, the results in this paper downplay the importance of the structurally determined weakness of monetary transmission in the NMS. Instead, they bring to the front the issue of longer response lags. If the observed longer lags in price reaction result from lower credibility of the NMS4 central banks, the credibility problem may be solved overnight by joining the monetary union. If learning under uncertainty is the issue, it could be speeded up by a change towards a less uncertain environment.

Finally, the results of this paper provide the following argument to the proponents of the EMU accession: Given the long lags with which the transmission mechanism operates in the NMS4, and the ultimate strength of its effect, it should be harder for the NMS4 central banks to run an effective stabilizing monetary policy. The independent monetary policy is more likely to be a source of additional variability and giving it up could end up being beneficial in the long run.

Appendix

A Conditional posteriors for the Gibbs sampler

In the model defined by equations (3)-(7) the joint posterior is $\propto$

$$\prod_i |\boldsymbol{\Sigma}_i|^{-\frac{T_i}{2}} \exp\left(-\frac{1}{2}\sum_i (\boldsymbol{y}_i - \tilde{\boldsymbol{X}}_i\boldsymbol{\beta}_i - \tilde{\boldsymbol{Z}}_i\gamma_i)'(\boldsymbol{\Sigma}_i^{-1}\otimes \boldsymbol{I}_{T_i})(\boldsymbol{y}_i - \tilde{\boldsymbol{X}}_i\boldsymbol{\beta}_i - \tilde{\boldsymbol{Z}}_i\gamma_i)\right)$$
$$\lambda^{-\frac{IJK}{2}} \exp\left(-\frac{1}{2}\sum_i (\boldsymbol{\beta}_i - \bar{\boldsymbol{\beta}})'\lambda^{-1}\boldsymbol{L}_i^{-1}(\boldsymbol{\beta}_i - \bar{\boldsymbol{\beta}})\right)\prod_i |\boldsymbol{\Sigma}_i|^{-\frac{J+1}{2}}\lambda^{-\frac{v+2}{2}}\exp\left(-\frac{1}{2}\frac{s}{\lambda}\right) \tag{14}$$

Define all data, $Y \equiv \{\boldsymbol{Y}_1, \ldots, \boldsymbol{Y}_I, \boldsymbol{X}_i, \ldots, \boldsymbol{X}_I, \boldsymbol{Z}_1, \ldots, \boldsymbol{Z}_I\}$ and the set of parameters, $\Theta \equiv \{\boldsymbol{\beta}_1, \ldots, \boldsymbol{\beta}_I, \gamma_1, \ldots, \gamma_I, \boldsymbol{\Sigma}_1, \ldots, \boldsymbol{\Sigma}_I, \bar{\boldsymbol{\beta}}, \lambda\}$.

Conditional posterior for $\boldsymbol{\beta}_i$ is:

$$p(\boldsymbol{\beta}_i|Y, \Theta\backslash\{\boldsymbol{\beta}_i\}) = \mathrm{N}(\boldsymbol{D}_i^{-1}\boldsymbol{d}_i, \boldsymbol{D}_i^{-1}) \tag{15}$$

where

$$\boldsymbol{D}_i = \boldsymbol{\Sigma}_i^{-1}\otimes \boldsymbol{X}_i'\boldsymbol{X}_i + \lambda^{-1}\boldsymbol{L}_i^{-1}$$
$$\boldsymbol{d}_i = (\boldsymbol{\Sigma}_i^{-1}\otimes \boldsymbol{X}_i')\,\mathrm{vec}(\boldsymbol{Y}_i - \boldsymbol{Z}_i\boldsymbol{\Gamma}_i) + \lambda^{-1}\boldsymbol{L}_i^{-1}\bar{\boldsymbol{\beta}}$$

Conditional posterior for $\boldsymbol{\gamma}_i$ is:

$$p(\gamma_i|Y, \Theta\backslash\{\gamma_i\}) = \mathrm{N}(\boldsymbol{F}_i^{-1}\boldsymbol{f}_i, \boldsymbol{F}_i^{-1}) \tag{16}$$

where

$$\boldsymbol{F}_i = \boldsymbol{\Sigma}_i^{-1} \otimes \boldsymbol{Z}_i'\boldsymbol{Z}_i$$
$$\boldsymbol{f}_i = (\boldsymbol{\Sigma}_i^{-1} \otimes \boldsymbol{Z}_i') \operatorname{vec}(\boldsymbol{Y}_i - \boldsymbol{X}_i\boldsymbol{B}_i) = \operatorname{vec}(\boldsymbol{Z}_i'(\boldsymbol{Y}_i - \boldsymbol{X}_i\boldsymbol{B}_i)\boldsymbol{\Sigma}_i^{-1})$$

Conditional posterior for $\bar{\boldsymbol{\beta}}$ is:

$$p(\bar{\boldsymbol{\beta}}|Y, \Theta\backslash\{\bar{\boldsymbol{\beta}}\}) = \mathrm{N}(\boldsymbol{G}_i^{-1}\boldsymbol{g}_i, \boldsymbol{G}_i^{-1}) \tag{17}$$

where

$$\boldsymbol{G}_i = \lambda^{-1}\sum_i \boldsymbol{L}_i^{-1}$$
$$\boldsymbol{g}_i = \lambda^{-1}\sum_i \boldsymbol{L}_i^{-1}\boldsymbol{\beta}_i$$

Conditional posterior for $\boldsymbol{\Sigma}_i$ is:

$$p(\boldsymbol{\Sigma}_i|Y, \Theta\backslash\{\boldsymbol{\Sigma}_i\}) \propto |\boldsymbol{\Sigma}_i|^{-\frac{T_i+J+1}{2}} \exp\left(-\frac{1}{2}\operatorname{tr}\boldsymbol{\Sigma}_i^{-1}\boldsymbol{U}_i'\boldsymbol{U}_i\right)$$

or

$$p(\boldsymbol{\Sigma}_i|Y, \Theta\backslash\{\boldsymbol{\Sigma}_i\}) = \mathrm{IW}(\boldsymbol{U}_i'\boldsymbol{U}_i, T_i) \tag{18}$$

where IW denotes the Inverted Wishart distribution (Bauwens et al., 1999, p.305).

Conditional posterior for λ is:

$$p(\lambda|Y, \Theta\backslash\{\lambda\}) \propto \lambda^{-\frac{IJK+v+2}{2}} \exp\left(-\frac{1}{2}\frac{(s + \sum_i(\boldsymbol{\beta}_i - \bar{\boldsymbol{\beta}})'\boldsymbol{L}_i^{-1}(\boldsymbol{\beta}_i - \bar{\boldsymbol{\beta}}))}{\lambda}\right)$$

or

$$p(\lambda|Y, \Theta\backslash\{\lambda\}) = \mathrm{IG}_2\left(s + \sum_i(\boldsymbol{\beta}_i - \bar{\boldsymbol{\beta}})'\boldsymbol{L}_i^{-1}(\boldsymbol{\beta}_i - \bar{\boldsymbol{\beta}}), IJK + v\right) \tag{19}$$

where IG_2 denotes the inverted gamma-2 distribution (Bauwens et al., 1999, p.292).

Simulating the posterior with the Gibbs sampler consists of randomly drawing parameters in Θ from (15)-(19), always conditioning on all most recently drawn parameters. See e.g. Gelman et al. (1995) for a detailed discussion of the Gibbs sampler.

B Imposing the sign restrictions

The sign restrictions in (13) can be viewed as the priors for the nonzero coefficients of the $\boldsymbol{G}$ which are uniform on the whole real line, on the positive or on the negative part of the real line. The requirement $\boldsymbol{G}\boldsymbol{G}' = \boldsymbol{\Sigma}$ constrains the coefficients to certain intervals, but otherwise all factorizations of the $\boldsymbol{\Sigma}$ are observationally equivalent, i.e. they result in exactly the same value of the likelihood function. Therefore, within this family, the posterior distribution coincides with the prior. The snug is that the coefficients of the $\boldsymbol{G}$ are linked by a nonlinear relationship and they cannot be simultaneously uniform on their admissible intervals.

Technically the sign restrictions are applied in a manner following Kieler and Saarenheimo (1998) and Canova and De Nicoló (2002), obtaining one factorization from another by means of a rotation matrix. The difference is that, thanks to the combination of sign restrictions with some zero restrictions, the resulting search for admissible rotations can be performed analytically, avoiding the computationally intensive numerical search technique of the above papers.

For any factorization $\boldsymbol{G}^*$ such that $\boldsymbol{G}^*\boldsymbol{G}^{*'} = \boldsymbol{\Sigma}$, and any orthogonal matrix $\boldsymbol{D}$, $\boldsymbol{G}^{**} = \boldsymbol{G}^*\boldsymbol{D}$ is also a factorization, i.e. satisfies $\boldsymbol{G}^{**}\boldsymbol{G}^{**'} = \boldsymbol{\Sigma}$. All orthogonal matrices (which correspond to orthogonal linear transformations) are products of sequences of rotations and reflections. Start from the Choleski decomposition of the $\boldsymbol{\Sigma}$. If the zeros in the first two rows are to be preserved, only rotations of the last two columns are allowed. The restriction of diagonal elements to be positive allows us to disregard reflections. Therefore, all matrices satisfying the zero restrictions can be obtained as:

$$\boldsymbol{G}(\theta) = \mathrm{Chol}(\boldsymbol{\Sigma}) \times \mathrm{Rotation}(3, 4, \theta) \tag{20}$$

where Chol() denotes the Choleski decomposition and Rotation(x,y,θ) is the matrix that rotates columns x and y by angle θ. Writing the above equation in detail:

$$\boldsymbol{G}(\theta) = \begin{pmatrix} c_{11} & 0 & 0 & 0 \\ c_{21} & c_{22} & 0 & 0 \\ c_{31} & c_{32} & c_{33} & 0 \\ c_{41} & c_{42} & c_{43} & c_{44} \end{pmatrix} \times \begin{pmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & \cos(\theta) & -\sin(\theta) \\ 0 & 0 & \sin(\theta) & \cos(\theta) \end{pmatrix}$$

When multiplying out the above matrices, the four sign restrictions on the lower right submatrix of $\boldsymbol{G}$, spelled out in equation (13), imply the system of four inequalities:

$$\begin{aligned}
c_{33}\cos\theta &> 0 \\
-c_{33}\sin\theta &> 0 \\
c_{43}\cos\theta + c_{44}\sin\theta &< 0 \\
-c_{43}\sin\theta + c_{44}\cos\theta &> 0
\end{aligned}$$

The above system can be solved for the rotation angle θ. The solution depends on the term c_{43}:

$$\theta \in \left(-\frac{\pi}{2}, \arctan\left(-\frac{c_{43}}{c_{44}}\right)\right) \text{ when } c_{43} > 0, \text{ and} \tag{21}$$

$$\theta \in \left(\arctan\left(\frac{c_{44}}{c_{43}}\right), 0\right) \text{ when } c_{43} < 0 \tag{22}$$

Going through the found range of rotation angles (and postmultiplying the Choleski decomposition of the residual variance by the resulting rotation matrices) we find all the matrices $\boldsymbol{G}$ satisfying the postulated zero and sign restrictions.

In reporting the results, one would like to integrate them over the posterior distribution of θ, which, on the admissible interval, coincides with the prior (since θ doesn't change the value of the likelihood function). The prior can be inferred from the prior for the elements of $\boldsymbol{G}$, but here we stumble on the mentioned problem: $p(\theta) \propto \cos(\theta)$, which corresponds to the uniform distribution of $\boldsymbol{G}_{(3,3)}$, results in highly skewed distributions of $\boldsymbol{G}_{(3,4)}$, $\boldsymbol{G}_{(4,3)}$ and $\boldsymbol{G}_{(4,4)}$, etc. As a compromise, we report results integrated over the *uniform* distribution of θ on its admissible interval, which produces moderately skewed distributions of all parameters. Results obtained with other candidate distributions turned out to be very similar and no conclusions are affected.

Therefore the computation of the posterior is following: for each draw of the residual variance matrix (obtained from the Gibbs sampler) 1) the Choleski decomposition is found, 2) the admissible range for θ is computed from the formula (21) or (22), 3) a random number is drawn from the uniform distribution on the computed range for θ, 4) matrix $\boldsymbol{G}$ is obtained with formula (20).

C Data sources and estimation periods

Data on individual countries: sample periods and IFS series codes.

Country	Begin	End	Obs	Industrial production	Prices	Interest Rate	Exchange rate
Czech Rep.	1997 (6)	2004 (10)	89	93566..CZF...	93564...ZF...	93560B..ZF...	935..AA.ZF...
Hungary	1995 (4)	2004 (10)	115	94466..CZF...	94464...ZF...	94460C..ZF...	944..AA.ZF...
Poland	1995 (6)	2004 (6)	109	96466..BZF...	96464...ZF...	96460B..ZF...	964..AA.ZF...
Slovenia	1995 (1)	2004 (11)	119	96166..CZF...	96164...ZF...	96160B..ZF...	961..AA.ZF...
Finland	1985 (1)	1998 (12)	168	172 66..I	172 64...	172 60B..	172 ..RB.
France	1985 (1)	1998 (12)	168	132 66..I	132 64...	132 60B..	132 ..RB.
Germany	1985 (1)	1998 (12)	168	134 66..I	134 64...	134 60B..	134 ..RB.
Italy	1985 (1)	1998 (12)	168	136 66..I	136 64...	136 60B..	136 ..RB.
Spain	1985 (1)	1998 (12)	168	184 66..I	184 64...	184 60B..	184 ..RB.

International data

Variable	IFS code	IFS name
oil prices	00176AAZZF...	petroleum:average crude price
non fuel commodity prices	00176NFDZF...	non fuel commodities index
federal funds rate	11160B..ZF...	federal funds rate

Notes: All data comes from the IMF International Financial Statistics (IFS) database. The data for the CEE countries and international data was downloaded from http://www.imf.org on February 24th, 2005. The data for the EU countries come from the IFS CD-ROM.

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