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

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Abstract

This paper estimates responses to monetary shocks for several of the current members of the EMU (in the pre-EMU sample) and for the Central and East European (CEE) countries, along with the mean response in each of the groups. The problem of the short sample, especially acute in the case of the CEE, is mitigated by using a Bayesian estimation which combines information across countries. The estimated responses are similar across regions, but there is some evidence of more lagged and deeper price responses in the CEE economies. If this results from structural differences, premature entering the EMU would cause problems, but if from credibility issues, entering the EMU would be desirable. Also, with longer response lags, conducting a stabilizing monetary policy should be difficult in the CEE currently, and giving it up might not be a big loss after all.

JEL classification: C11, C15, C33, E40, E52

Keywords: monetary policy transmission, Structural VAR, Bayesian estimation, exchangeable prior

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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 bearing on monetary transmission. If responses to monetary shocks are significantly heterogeneous, and it is for reasons that do not disappear within the monetary union, conducting the 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 then average inflation.

Examples of papers discussing the impact of structural characteristics of the European economies on their monetary transmission are Dornbusch et al. (1998), Guiso et al. (1999), Mihov (2001), Ehrmann et al. (2003). These papers look for indicators of interest sensitivity of output, size, health and structure of the banking sector, stock market capitalization and other, and try to relate them 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) (Ballabriga et al. (1999) find asymmetries among European economies which may, but need not, be related to monetary transmission). 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 Central and East European (CEE) countries, which joined the European Union in 2004, and which are legally obliged to adopt the euro some time afterwards. For example, Ganev et al. (2002) review the studies conducted in the CEE central banks, and add their own empirical analysis. They express the consensus view 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 is missing so far, is an explicit comparison of responses to monetary shocks across the two regions: CEE and the Western Europe. Such comparison can be expected to be more meaningful and interesting than intra-regional comparisons, 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 regions despite short data series and in a unified framework. The comparison yields interesting results: First, the monetary shocks in the CEE are associated with more persistent movements of the interest rate and the exchange rate. Second, in spite of all the structural differences between the regions, responses to their respective monetary shocks are broadly similar (the differences in responses are usually not significant at standard significance levels). However, with posterior probability in excess of 80% the CEE prices actually respond stronger, although after a longer lag then in the euro-area countries. Alternative interpretations of these results are proposed, and their implications for the EMU accession by the CEE countries are considered.

The principal obstacle in the study of the CEE region are the short available data series. The strategy employed here to mitigate this problem is to introduce to the estimation the prior intuition that parameters of VAR models for individual countries may be similar, since they are all special cases of the same underlying economic model. This is achieved by estimating VARs for a group of countries as a Hierarchical Linear Model: a Bayesian prior is specified, that the coefficients for each unit are randomly drawn from a distribution with a common mean (an exchangeable prior). The second stage of the hierarchy is formed by a prior about the parameters of this distribution, which may be uninformative. This estimation strategy avoids the heterogeneity bias in a dynamic panel (Pesaran and Smith, 1995) and the results of Hsiao et al. (1999) suggest that it should have a small sample performance that is superior to available classical alternatives.

An early treatment of the estimation of a linear model with the exchangeable prior is in Lindley and Smith (1972). This paper follows the more modern approach of Gelman et al. (1995) which takes advantage of the simulation based numerical methods that have become available since then. This allows to incorporate uncertainty about prior parameters and test hypotheses about them and their influence on the results.

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 (2000). This finding suggests, that it should also increase reliability of a structural analysis. However, it has not been used for Structural VARs, except for a parallel paper of Canova and Pappa (2003), which uses a different technique of working with the similar prior.

The second, methodological contribution of this paper is the proposed identification of the monetary shocks. They are identified using minimal assumptions: 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. The latter is a sign (inequality) restriction. Such restrictions have been applied among others in Uhlig (2001) and Faust (1998). This paper's approach is closest to that of Canova and De Nicoló (2002) and Kieler and Saarenheimo (1998) in its use of the rotation matrix. However, unlike in the mentioned papers, the combination of zero and sign restrictions enables one to find the admissible range of rotation angles analytically, avoiding a numerical search procedure.

The structure of the paper is following: the second section discusses estimation of reduced form VARs as a Hierarchical Linear Model, the third: identification of the structural model, the fourth presents the results, the fifth discusses the observed differences between the Eastern and the Western Europe, and the sixth contains conclusions. Details about the data and estimation samples are in the appendix.

2 Estimation

2.1 The statistical model

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

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

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

where $\boldsymbol{y}_{it}$ is a vector of J endogenous variables, $\boldsymbol{w}_t$ is a vector of exogenous variables which are common across countries and $\boldsymbol{z}_{it}$ is a vector of other exogenous variables which may or may not differ across countries. We will specify an exchangeable prior about the coefficients of $\boldsymbol{y}_{i(t-l)}$ and $\boldsymbol{w}_t$, while the prior will be uninformative for country specific variables in $\boldsymbol{z}_{it}$. Vector $\boldsymbol{u}_{it}$ contains VAR innovations which are i.i.d. $N(0, \boldsymbol{\Sigma}_i)$.

The model in terms of data matrices is:

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

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:

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

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:

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

We 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:

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

where IG_2 denotes the inverted gamma 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 is not 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 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 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 (see e.g. Casella and George, 1992).

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.

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 the propriety of the posterior 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 CEE, 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 then 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. It is assumed that the central banks target the short term interest rate, so no monetary aggregate is included.

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 $\boldsymbol{v}_{it}$ and the vector of VAR innovations $\boldsymbol{u}_{it}$ is following:

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

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

The identification restrictions[1] 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:

$$\begin{pmatrix} + & 0 & 0 & 0 \\ \bullet & + & 0 & 0 \\ \bullet & \bullet & + & + \\ \bullet & \bullet & - & + \end{pmatrix} \begin{pmatrix} v_{it1} \\ v_{it2} \\ \widehat{v_{it3}} \\ v_{it4} \end{pmatrix} = \begin{pmatrix} u_{it1} \\ u_{it2} \\ u_{it3} \\ u_{it4} \end{pmatrix} \tag{13}$$

where + denote coefficients that are constrained to be positive, 0 - zero restrictions and $\bullet$ - unconstrained coefficients. $\widehat{v_{it3}}$ is the monetary policy shock.

[1] 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 and Roubini (2000) and many 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, and it is expected that resulting estimates of the relationships between interest rate and exchange rate are 'reasonable', i.e. as in assumption 2 above.

The problem with this approach is that the zero restrictions, on which the identification hinges, are contentious. While it is true that official data on output and prices are compiled with delay, 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. Several recent papers (Faust, 1998; Uhlig, 2001; Canova and De Nicoló, 2002) have criticized the zero restrictions and suggested to use sign restrictions, which are more grounded in economic theory.

4 Results

The model is estimated[2] for two panels of countries: seven euro-area countries and four CEE countries. Since the identification scheme assumes an open economy with a flexible (i.e. not fixed) exchange rate, the countries and sample periods with fixed exchange rate regimes were excluded. The sample periods for the CEE countries span the second half of 1990-s up to early 2004 and differ for each country, depending on when the 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 CEE 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,

[2]The posterior is simulated with the Gibbs sampler: 1000 draws are made after initial 1000 discarded draws.

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. The interest rate of 0.1 corresponds to 10% (1000 basis points). The variables other then the interest rate are logs of indexes that assume the value 100 in December 1995. The basic specification contains six lags of the endogenous variables. Lags 0 and 1 of the US Federal Funds Rate, log of the US IIP, log oil prices and log non-fuel commodity prices serve as the common exogenous controls (to which the exchangeable prior extends), which capture the world developments. The model includes also country specific constant terms and the contemporaneous German interest rate for countries other than Germany.

The maximum euro-area panel consists of Finland, France, Germany, Greece, Italy, Portugal and Spain. (Austria, Belgium and Netherlands were not considered because of their quasi-fixed exchange rate against the D-Mark, 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. Portugal proved to be an outlier because of dramatic swings of its interest rates in the wake of the 1992 European Monetary System crisis. Therefore, both countries are considered only in robustness checks.

The maximum CEE panel consists of Czech Republic, Slovakia, Hungary, Poland, Slovenia and Romania (Bulgaria and the Baltic countries were not considered 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 suffered 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, the basic results are presented for the Czech Republic, Hungary, Poland and Slovenia, while Slovakia and Romania are considered only in robustness checks.

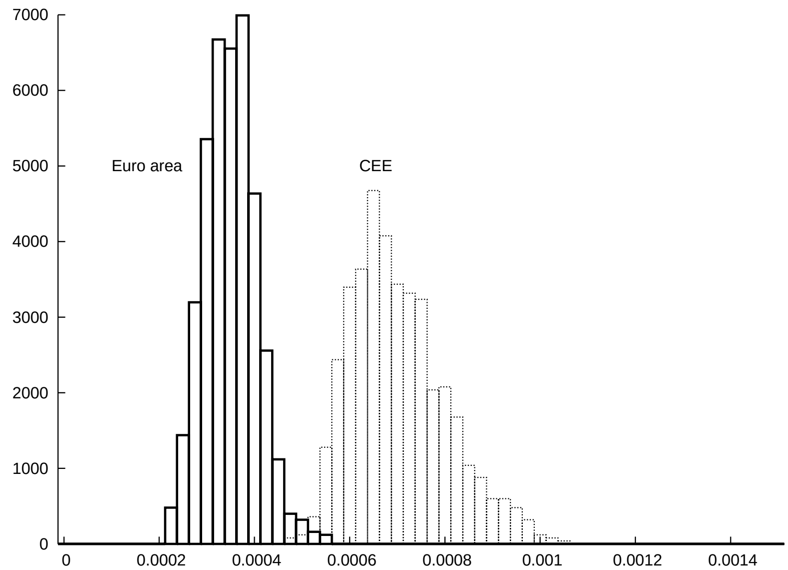

Figure 1: Simulated posterior distribution of λ for the euro-area panel (uninformative prior) and for the CEE panel (informative prior)

4.1 Homogeneity of the two panels

Homogeneity 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.

The posterior density of the overall tightness parameter λ in each of the panels is presented in figure 1.

In the estimation for the euro area, the prior for λ is noninformative. Nevertheless, the likelihood turns out to be very informative about this parameter 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, and using the posterior odds ratio we would reject both the hypothesis of complete homogeneity, and of independence of the individual country models.

The CEE panel turns out to be much more heterogeneous: when the prior is noninformative, the posterior mean of the tightness parameter λ is 6 times as large as that for the euro-area data. However, a priori we could believe that the CEE region is as homogeneous as the euro area (we certainly think that the regions may differ a lot between them, but there are no obvious reasons to believe that the CEE is more heterogeneous than Western Europe). The prior distribution for the λ that reflects this belief is just the posterior distribution of λ for the euro-area panel. For computational convenience, we fit an inverted gamma distribution to that posterior distribution, and use it as the prior for the CEE panel. Compared with the diffuse prior, the informative prior for λ has almost no effect on the mean impulse responses, but it results (as could be expected) in narrower error bands. With the informative prior, the posterior mean of λ is still 2 times higher then for the euro-area panel (as can be seen in figure 1).

4.2 'Mean' impulse responses for the panels

The impulse responses to a one standard deviation monetary shock are calculated for 40 months. The plots display the median, the 5th and the 95th centiles of the posterior distribution, found by Monte Carlo simulation of the posterior. Figure 2 presents the impulse responses implied by the 'mean' model for each panel ($\bar{\boldsymbol{\beta}}$). Impulse responses of individual countries, which are of separate interest themselves, are presented in the following section. They all share the same qualitative features.

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 an interest rate increase by 40 basis points on average in the euro area and almost 80 bp in the CEE. The initial appreciation is on average by 1% (maximally 2%) in both regions. Both the interest rate increase and the appreciation are reversed after one year. The economies respond with a transitory output decline and a possibly permanent reduction of the price level.

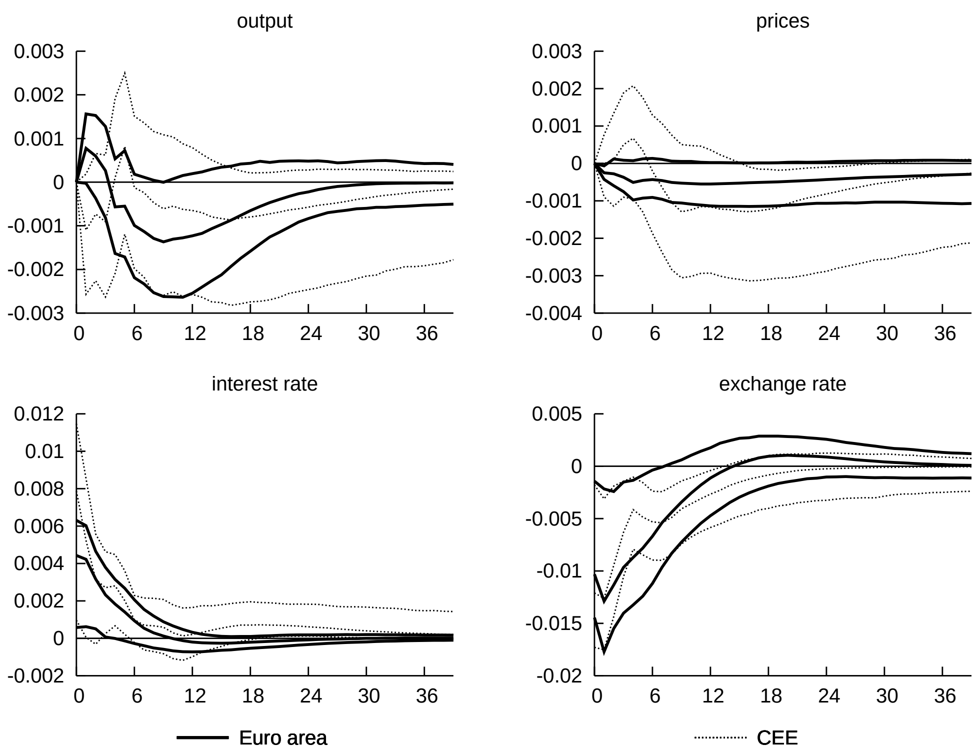

Figure 2: Mean impulse responses for the CEE and the euro area

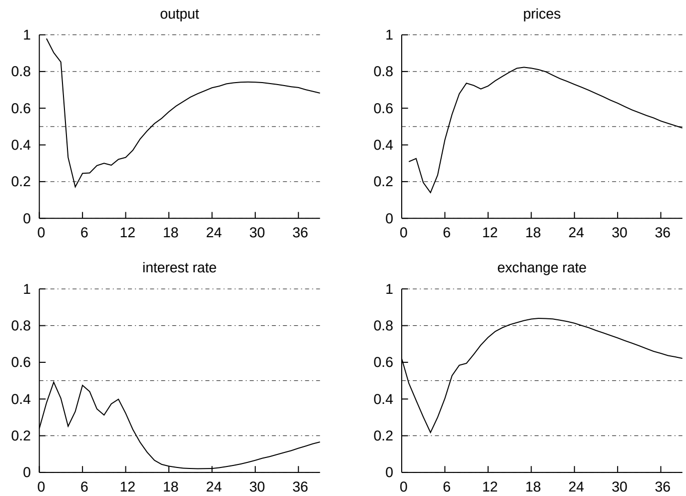

Figure 3: Posterior probability that the response of the CEE panel is deeper (more negative) then that of the euro-area panel

The responses of both regions are generally similar and their uncertainty bands mostly overlap, but there are some systematic differences. To quantify the posterior uncertainty about the differences in responses, in figure 3 we plot, for each variable an period, the *posterior probability that the response of the CEE panel is deeper (more negative) then that of the euro-area panel*. As often in the VAR analysis, the difference is usually not significant at standard significance levels. However, the following observations have some posterior support in excess of 80%:

1. Interest rate and exchange rate responses are more persistent in the CEE.
2. In the short run, prices responses are more significant in the euro area.
3. In the medium run, prices fall more in the CEE.

The validity of observation 1 can be read from the lower panels of the figure 3. Probability of the interest rate in the CEE being lower then in the euro area is never high, and for most of the second year after the shock it is lower then 10%, while the probability that the CEE exchange rate is lower (has appreciated more) is above 80% in the

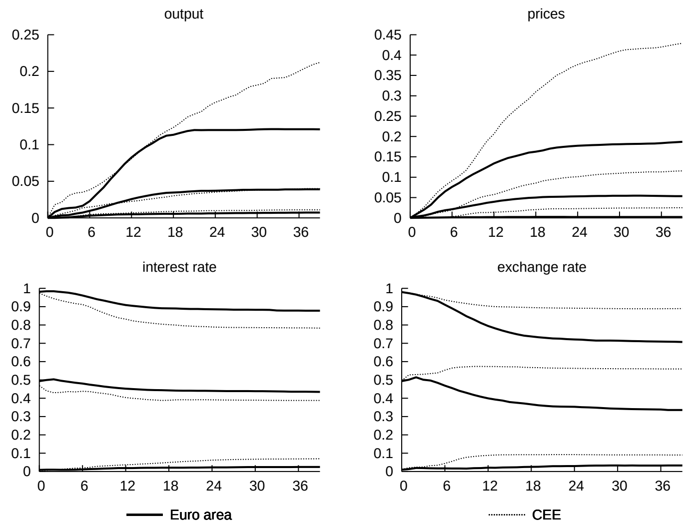

Figure 4: Share of monetary shocks in the variance decompositions for the CEE and the euro-area panels: median, 5th and 95th percentiles of the posterior distributions

second year. Observations 2 and 3 about price responses receive the posterior support of around 80%, and are very robust to specification and sample changes.

The comparison of output responses is less clear cut and, in general, not robust to specification and sample changes. This may be related to the fact that we have more noisy measures of output, then we have of the other variables. The conclusion of the VAR exercise seems to be that output responses are of similar magnitude in both panels.

The analysis of impulse responses is complemented with that of variance decompositions, plotted in figure 4. The posterior distributions overlap significantly, but monetary shocks tend to be responsible for a greater share of variance of prices and exchange rates in the CEE.

4.3 Results for individual countries

Figures 5 and 6 present impulse responses to a monetary shock for each of the analyzed countries. Consistently with the results for the posterior tightness of the exchangeable prior (λ), the impulse responses for the CEE panel are somewhat more heterogeneous then for the euro-area panel.

Interest rate responses are clearly weakest in Germany, and somewhat stronger in Spain then in other countries. Spain has also the strongest and most significant price and output responses. Output responses of Germany and Italy are not significant. Price responses are not significant for Finland, France and Germany, although they are generally correct, i.e. most of the posterior probability mass is on the negative responses. For some other subsamples and sets of exogenous controls we obtain more significant responses, but never much larger in magnitude.

In the CEE panel, the impulse responses are strongest and most significant for the Czech republic and Slovenia. In the case of Hungary, the output response is insignificant, while in the case of Poland the price response is insignificant, both with very wide uncertainty bands. Interest rate shocks are found to be strongest in Poland, and weakest in Hungary. The individual country responses tend to look more in line with the theory e.g. when a smaller set of exogenous control variables is used. However, the comparison of the mean responses, which is in the focus of this paper, is not affected by such changes.

The considered samples are short, and for individual countries imply only 3.1 observations per estimated parameter in the CEE (and 2.4 for the Czech Republic), and maximally no more then 5 in the euro area. For the CEE countries these are maximum samples under flexible exchange rate regime. Given the short data, individual country models may not be very reliable. However, in theory they should be more reliable then with a usual single country estimation, because in the computation of the posterior, the country data is augmented with the information for all other countries in the panel.

4.4 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 coun-

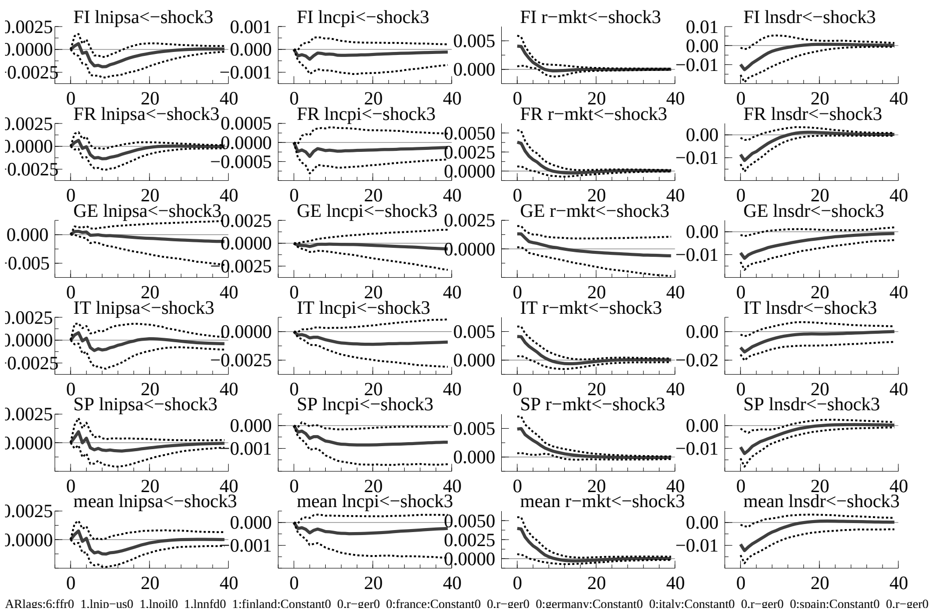

Figure 5: Impulse responses for the euro-area panel

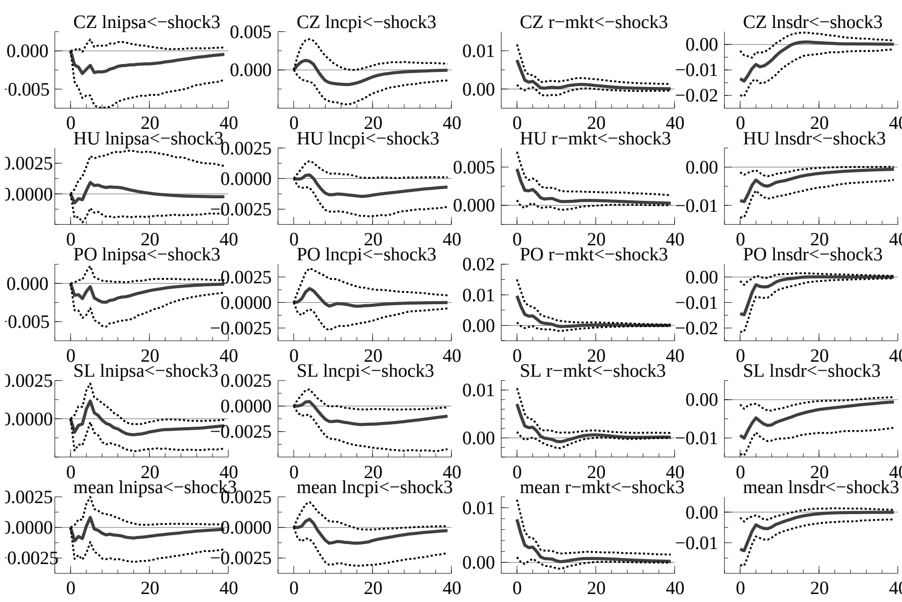

Figure 6: Impulse responses for the CEE panel, informative prior for λ

tries 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 then in other countries affected by the crisis, partly because of the relatively small volume of the market. The resulting interest rate swings were perceived 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 CEE 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 CEE 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 CEE 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 CEE panel the positive price response is very strong and output also responds positively.

5 Differences in monetary transmission between the East and the West

The finding of weaker responses in the CEE in the short run is consistent with the argument, that countries with smaller financial systems, in relation to their GDP, are expected to have a weaker monetary transmission. In the CEE, 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).

On the other hand, there are some reasons to believe that the CEE economies should be more responsive to monetary policy:

- First, the exchange rate channel might be stronger. The CEE 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 (I thank Bartosz Maćkowiak for this point). Similarly, Anzuini and Levy (2004) find that the agents in the CEE have a lot of unhedged foreign debt, and are very exposed to the foreign exchange rate risk.

- Third, 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 CEE 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 then the euro area (Suppel, 2003), which is mostly financed by reinvesting profits.

The bottom line is that it is not obvious if the financial and economic structures of the CEE economies imply a weaker or stronger monetary transmission, compared with the euro area.

The common view, expressed e.g. in Ganev et al. (2002) seems to be that the factors that make monetary transmission weaker should prevail. This paper makes the first (to my knowledge) explicit comparison and confirms this view as regards the short run. However, there is some measure of evidence that in the medium run, monetary policy in the CEE is actually quite effective. After some time, the stronger and/or more protracted interest rate hikes and exchange rate appreciations end up having their effect.

Apart from the structural characteristics of the economies, the differences in monetary transmission can be related to the credibility of the central bank policies. In this respect, the crucial difference between the East and the West is the short history of market economy, and short track record of the central banks in the former.

If the CEE central banks enjoy less credibility, economic agents adjust prices more sluggishly. Therefore contractionary monetary shock will have more output cost in the short run, and will affect prices only in the longer term. Moreover, the central bank has to make more vigorous and protracted movements of its policy instruments, in order to convince the doubting agents of its intentions. The results found here provide support for this scenario: while there is no clear evidence of output reacting more in the short run, we find prices adjusting more sluggishly and more persistent interest and exchange rate changes.

The lagged reaction observed in the CEE panel could 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.

6 Conclusions

This paper makes one of the first systematic comparisons of the responses to monetary shocks in the Western Europe and in the CEE. The responses of the CEE countries turn out to be broadly similar, but with interesting differences (albeit estimated with significant uncertainty): monetary shocks tend to be more persistent, and generate a more delayed but stronger prices response.

The empirical findings of this paper shed new light on the general belief, that due to the lower financial depth, monetary transmission is weaker in the CEE than in the Western Europe. As far as the short term effects are concerned, this view is corroborated. Consistently with it, monetary policy is more decisive, with more persistent shocks. In the longer term, such policy turns out to have a strong impact on prices.

Lessons regarding the desirability of the EMU accession depend on the interpretation of these results. If the more lagged price responses result from underdevelopment of the financial markets of the CEE economies, this suggests possible problems when the CEE countries join the EMU prematurely. The argumentation is following: First, we find strong responses in CEE conditional on probably stronger and more persistent monetary shocks generated by the current central banks. Fine-tuned policies of the ECB may have a weaker and even more lagged effect on the CEE economies. Second, as a consequence of joining the EMU, the exchange rate channel, which may be playing important role currently, will be drastically weakened, rendering the CEE economies even less responsive to monetary policy. This makes likely the scenario predicted by Dornbush and observed within the present EMU: that interest rates that would match the European core are too low for the peripheral, poorer and faster growing countries. The peripheries observe persistently higher inflation rates and may end up developing financial and property market bubbles.

The problem with this interpretation is that it assumes, that the CEE central banks are able to compensate for the structurally motivated weakness of responses by a conducting a more 'stubborn' policy. But if, for structural reasons, a 80bp discount rate increase does not make much difference for the economy, why would a sustained 20bp difference do the trick? This, however, would be perfectly plausible in the credibility interpretation of the difference with the West.

If the observed differences result from lower credibility of the CEE central banks, the implication for the EMU accession is quite opposite: the credibility problem could be solved overnight by joining the monetary union, provided that the ECB will enjoy the same credibility in the East as in the West. If learning under uncertainty is the issue, it could be speeded up by a discrete 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 CEE, and the ultimate strength of its effect, it should be extremely hard for the CEE central banks to run a conscious 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.

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A 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:

$$\boldsymbol{G}(\theta) = \text{Chol}(\boldsymbol{\Sigma}) \times \text{Rotation}(3, 4, \theta) \tag{14}$$

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}:

$$\theta \in \left(-\frac{\pi}{2}, \arctan\left(-\frac{c_{43}}{c_{44}}\right)\right) \text{ when } c_{43} > 0 \text{, and} \tag{15}$$

$$\theta \in \left(\arctan\left(\frac{c_{44}}{c_{43}}\right), 0\right) \text{ when } c_{43} < 0 \tag{16}$$

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 (15) or (16), 3) a random number is drawn from the uniform distribution on the computed range for θ, 4) matrix $\boldsymbol{G}$ is obtained with formula (14).

B Data and estimation periods

B.1 CEE panel

B.1.1 Czech Republic

Sample: 1997 (6) - 2004 (1) (80 observations)

output	OECD	772027KSA	Industrial production ISIC C-E sa
prices	OECD	775241K	CPI All Items
interest rate	OECD	776225D	3-month PRIBOR
exchange rate	IFS	935 ..RB.	Principal Rate, Koruny, per SDR

B.1.2 Hungary

Sample: 1995 (4) - 2004 (1) (106 observations)

output	OECD	802027KSA	Industrial production ISIC C-E sa
prices	IFS	944 64...	Consumer Prices
interest rate	IFS	944 60C..	Treasury Bill Rate
exchange rate	IFS	944 ..RB.	Official Rate, Forint, per SDR

B.1.3 Poland

Sample: 1995 (7) - 2004 (1) (103 observations)

output	OECD	812027KSA	Industrial production ISIC C-E sa
prices	OECD	815241K	CPI All Items
interest rate	IFS	964 60B..	Money Market Rate
exchange rate	IFS	964 ..RB.	Official Rate, Zlotys, per SDR

B.1.4 Slovenia

Sample: 1995 (1) - 2003 (10) (106 observations)

output	IFS	961 66..C	Industrial Production Seas. Adj
prices	IFS	961 64...	Consumer Prices
interest rate	IFS	961 60B..	Money Market Rate
exchange rate	IFS	961 ..RB.	Official Rate, Tolars, per SDR

B.2 Euro area panel

Common sample: 1985 (1) - 1998 (12) (168 observations, or subsamples of 108 observations)

B.2.1 Finland

output	IFS	972 66..I	Industrial Production Seas. Adj
prices	IFS	972 64...	Consumer Prices
interest rate	OECD	646225D	3-MONTH HELIBOR
exchange rate	IFS	172 ..RB.	Official Rate, Natl Currency, per SDR

B.2.2 France

output	IFS	132 66..I	Industrial Production Seas. Adj
prices	IFS	132 64...	Consumer Prices
interest rate	IFS	132 60C..	Treasury Bills:3 Months
exchange rate	IFS	132 ..RB.	Official Rate, Natl Currency, per SDR

B.2.3 Germany

output	IFS	134 66..I	Industrial Production Seas. Adj
prices	IFS	134 64...	Consumer Prices
interest rate	IFS	134 60BS.	Interbank Deposit Rate (Imf-sdr)
exchange rate	IFS	134 ..RB.	Market Rate, Natl Currency, per SDR

B.2.4 Italy

output	IFS	136 66..I	Industrial Production Seas. Adj
prices	IFS	136 64...	Consumer Prices
interest rate	IFS	136 60B..	Money Market Rate
exchange rate	IFS	136 ..RB.	Market Rate, Natl Currency, per SDR

B.2.5 Spain

output	IFS	184 66..I	Industrial Production Seas. Adj
prices	IFS	184 64...	Consumer Prices
interest rate	IFS	184 60B..	Call Money Rate
exchange rate	IFS	184 ..RB.	Market Rate, Natl Currency, per SDR