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

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Abstract

This paper compares impulse responses to monetary policy shocks in the euro area countries before the EMU and in the New Member States (NMS) from central-eastern Europe. We mitigate the small-sample problem, which is especially acute for the NMS, by using a Bayesian estimation that combines information across countries. The impulse responses in the NMS are broadly similar to those in the euro area countries. There is some evidence that in the NMS, which have had higher and more volatile inflation, the Phillips curve is steeper than in the euro area countries. This finding is consistent with economic theory.

JEL classification: C11, C32, C33, E40, E52

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

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1 Introduction

Ever since the seminal paper by Sims (1980), which introduced Vector Autoregressions (VARs), there has been interest in comparing impulse responses to monetary policy shocks estimated for different countries.[1] These comparisons were motivated by the study of robustness, and tended to stress similarities across countries. The interest in potential cross-country differences in monetary transmission was boosted prior to the creation of the economic and monetary union (EMU) in Europe, because such differences are highly undesirable within currency unions.[2] It was pointed out that significant differences exist in some countries' structural characteristics, including among potential EMU members. Understanding how structural differences affect economies' responses to monetary policy is especially relevant in the context of prospective currency unions, but is also of great importance for monetary economics and policy in general.

Examples of papers discussing the impact of European economies' structural characteristics on the monetary transmission are Dornbusch et al. (1998); Cecchetti (1999); Guiso et al. (1999); Mihov (2001) and Ehrmann et al. (2003). These papers first look at indicators of interest sensitivity of output, size, health and structure of the banking sector, stock market capitalization and other. They then relate them, by theoretical reasoning, to the strength of monetary transmission. Results of this type of analysis are often ambiguous, as different characteristics sometimes have conflicting implications, and their relative quantitative importance is uncertain. The ultimate judgment has to come from macroeconomic data, usually analyzed with a Structural VAR technique. Papers in this line of research naturally fall into two categories: those that find significant and interpretable differences among the examined countries, and those that do not. Examples of the first group are Ramaswamy and Sløk (1998); Cecchetti (1999) and Mihov (2001). Papers in the second category, such as Kieler and Saarenheimo (1998) and Mojon and Peersman (2001), find that whatever asymmetries in monetary transmission may exist among EU countries, they are not strong enough to be robustly

[1]Sims (1980) compared impulse responses to monetary and other innovations in the US and Germany. Other examples of studies applying common identifying assumptions to several countries are Sims (1992); Kim (1999); and Kim and Roubini (2000); as well as papers quoted below.

[2]As discussed e.g. in Dornbusch et al. (1998), the concern is that the burden of disinflation would fall disproportionately on some countries, while others would have to accept higher than average inflation. This would make conducting common monetary policy politically difficult. This problem is independent from the long-debated question of whether current and potential future EMU member countries constitute an optimal currency area.

detected in the available data.[3]

The EU New Member States (NMS) from the central-eastern Europe are, in many important respects, quite different from the old EU members. It is well known that their financial systems, measured by total financial assets or stock market capitalization, are much smaller relative to GDP than those of the old EU member states (see e.g. Angeloni et al., 2005). It is reasonable to expect that in such conditions monetary policy may have a weaker impact on the economy. Central banks in the NMS also have shorter track records, which may make it more difficult to affect agents' expectations, potentially resulting in longer price response lags.

On the other hand, even with small institutional financial markets, prevailing interest rates still matter for many economic decisions and transactions, such as trade credits or the reinvestment of profits. In the context of other countries, there is also research suggesting that financial system imperfections may actually strengthen monetary transmission (see e.g. Kashyap and Stein, 2000).

Another potentially important feature of the NMS is that they have had higher average inflation rates. This is shown in Table 1 (the choice of countries and sample periods is justified in section 3). Economic theories predict that this should have implications for monetary transmission. New-Keynesian models which allow price stickiness to be determined endogenously predict that in a higher inflation environment agents adjust their prices more often, so there is less price stickiness (see Ball et al., 1988; Dotsey et al., 1999). With less price stickiness, the Phillips curve is steeper, and, consequently, the output cost of disinflation (the 'sacrifice ratio') is lower. This is confirmed empirically in Ball et al. (1988) cross-country study. In practice, the level and the standard deviation of inflation are positively correlated (which is confirmed in Table 1). In the Lucas (1973) imperfect information model, more volatile inflation makes agents, ceteris paribus, adjust prices more than their output. Therefore, the Lucas model also predicts that the NMS should have a steeper Phillips curve.

VAR studies of monetary transmission have also been extended to the NMS. Examples are Elbourne and de Haan (2006); Gavin and Kemme (2004) and Anzuini and Levy (2007). Ganev et al. (2002) and Coricelli et al. (2006) contain surveys of this literature. Most of these works concentrated on comparisons between individual NMS, and it seems that few robust lessons

[3]These papers study VAR impulse responses to national monetary policy shocks. In an alternative approach, Ciccarelli and Rebucci (2006) estimate responses to common monetary policy shocks under fixed exchange rates, thus mimicking the hypothetical situation inside a monetary union. This approach does not use VARs, as in VARs exchange rates are determined endogenously, and shocks that leave them fixed are difficult to construct.

Table 1: Annual CPI inflation in the East and the West.

	Czech Rep.	Hungary	Poland	Slovenia
1995	9.2	28.3	28.1	13.5
1996	8.8	23.6	19.8	9.8
1997	8.5	18.3	15.1	8.4
1998	10.6	14.2	11.7	7.9
1999	2.1	10.0	7.3	6.1
2000	3.9	9.8	10.1	8.9
2001	4.7	9.2	5.5	8.4
2002	1.8	5.3	1.9	7.5
2003	0.1	4.6	0.8	5.6
2004	2.8	6.8	3.6	3.6
2005	1.8	3.6	2.1	2.5
2006	2.5	3.9	1.1	2.5
2007	2.9	7.9	2.4	3.6
median	2.7	9.2	5.5	6.8
std.dev.	2.8	7.8	8.4	2.5

	Finland	France	Italy	Portugal	Spain
1987	4.1	3.3	4.7	9.3	5.2
1988	5.1	2.7	5.1	9.7	4.8
1989	6.6	3.5	6.2	12.6	6.8
1990	6.1	3.4	6.5	13.4	6.7
1991	4.1	3.2	6.3	10.9	5.9
1992	2.6	2.4	5.1	8.9	5.9
1993	2.1	2.1	4.5	6.5	4.6
1994	1.1	1.7	4.0	5.2	4.7
1995	1.0	1.8	5.2	4.1	4.7
1996	0.6	2.0	4.0	3.1	3.6
1997	1.2	1.2	2.0	2.2	2.0
1998	1.4	0.6	2.0	2.7	1.8
median	2.4	2.2	4.9	7.7	4.8
std.dev.	2.1	0.9	1.5	3.9	1.6

Source: IMF International Financial Statistics. Medians and standard deviations are calculated for all periods presented except for the Czech Republic (years 1998-2007) and Slovenia (years 1997-2006). These samples coincide with the samples used later in the paper.

emerge regarding qualitative differences across regions.

What has been largely missing so far is an explicit comparison of impulse responses to monetary policy shocks across the two regions - central-eastern and western Europe - estimated with a consistent methodology. Such a comparison can be expected to be more meaningful and interesting than the many intra-regional comparisons performed so far, as the structural differences between these regions dwarf those within them. This paper fills this gap by using an econometric technique which makes it possible to robustly estimate responses for both regions despite short data series.

The comparison yields interesting results: First, in spite of the structural differences between the regions, the impulse responses of output and prices to monetary policy shocks are similar in the NMS and in the euro area countries. We do not find support for the relative ineffectiveness of monetary policy in the NMS. Uncertainty bands for NMS price responses are much wider and they include the possibility of even stronger effects of monetary policy than in the euro area countries. Second, we find that, consistently with the predictions of economic theories, the NMS may have steeper Phillips curves, so that they face a lower output cost of disinflation (sacrifice ratio). Furthermore, we find that in later subsamples, when inflation in the NMS has abated, price responses become weaker and monetary transmission more similar to that in the euro area. We conclude that the structural weakness of monetary transmission in the NMS is quantitatively less important than often believed, and that the effect of higher and more volatile inflation rates in the sample may be important.

The principal obstacle in the study of the former centrally planned economies are the short available data series. To mitigate this problem, we perform a Bayesian estimation with the prior (called the exchangeable prior), which conveys the intuition that parameters of the 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 the posterior which pools information across countries, ensuring efficient use of the scarce data. The classical discussion of exchangeable priors for linear regression models can be found in Lindley and Smith (1972). The present paper adapts to VARs the formulation of Gelman et al. (2003), called the Hierarchical Linear Model.

Exchangeable priors have had a number of applications in econometric studies. The closest paper is Canova (2005), who uses the exchangeable prior in the estimation of VARs for Latin American countries to study the transmission of US shocks. Another related work is Ciccarelli and Rebucci (2006), who use the exchangeable prior in the estimation of a small multi-country time-varying model to study heterogeneity and time-variation of responses to monetary policy in the run-up to the EMU (see footnote 3). Other applications of the exchangeable prior are Canova and Marcet (1995), who study income convergence in panels of countries and regions; Zellner and Hong (1989), who find that the exchangeable prior improves the out-of-sample forecasting performance of time series models; and Canova and Ciccarelli (2004), who use it in their forecasting time-varying VARs.

In a related line of research, Gavin and Kemme (2004) use the posteriors from the OECD economies as priors for monetary VARs for the NMS and find that this prior improves their forecasting performance. The use of developed country data as a source of prior for the former centrally planned economies had been pioneered by Leamer and Taylor (1999). In contrast to these studies, in the present paper, priors for the central-eastern European countries are exchangeable only within the region, and are unaffected by the information from western Europe.

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

2 Econometric model

VAR models contain many parameters and with short samples, such as those available for the former centrally planned economies, they produce 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 jointly whole regions (first euro area, then NMS) 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 a dynamic model such as a VAR, it could seriously distort the results (Pesaran and Smith (1995) show that it results in an inconsistent estimator).

The estimation procedure is Bayesian and uses the Hierarchical Linear Model of Gelman et al. (2003).[4] The idea of similarity is formalized as a

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

Gaussian prior for each country's coefficients, which is centered at a common mean for 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: the common mean and the variance of country coefficients around the common mean (hypervariance). The Hierarchical Linear Model allows to specify the priors in the second stage of the hierarchy as noninformative, and let the data speak about the posterior common mean and hypervariance, given the assumed likelihood and prior structure. Intuitively, more different and more tightly estimated country coefficients increase the posterior probability of large hypervariance values. 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 likely to be similar across countries and those that need not be similar. We apply the exchangeable prior to parameters that determine dynamic interrelationships between endogenous variables, and reactions to some of exogenous variables. We specify a noninformative prior for the coefficients of other exogenous variables (those which seem to have most heterogeneous effects on countries) and for constant terms, which implies that we have country fixed effects.

The following two subsections specify the prior in the panel VAR setup: first the overall framework, and then the parameterization of the hypervariance. The third subsection explains the identification of monetary policy shocks.

2.1 Panel of VARs as a Hierarchical Linear Model

In what follows, vectors are denoted by lowercase bold symbols, matrices by uppercase bold symbols, $c = 1 \dots C$ denotes countries, $l = 1 \dots L$ denotes lags, $t = 1 \dots T_c$ denotes time periods.

For each country c in the panel we consider a reduced form N-dimensional VAR model of the form:

$$\boldsymbol{y}_{ct} = \sum_{l=1}^{L} \boldsymbol{B}'_{cl} \boldsymbol{y}_{c(t-l)} + \boldsymbol{\Delta}'_c \boldsymbol{w}_t + \boldsymbol{\Gamma}'_c \boldsymbol{z}_{ct} + \boldsymbol{u}_{ct} \tag{1}$$

$\boldsymbol{y}_{ct}$ is a vector of N endogenous variables and $\boldsymbol{w}_t$ is a vector of W exogenous variables which are common across countries. We specify an exchangeable prior about the coefficients of $\boldsymbol{y}_{c(t-l)}$ and $\boldsymbol{w}_t$. The prior for the coefficients of

$\boldsymbol{z}_{ct}$, which include, for example, country-specific constant terms, is noninformative. The vector $\boldsymbol{u}_{ct}$ contains VAR innovations which are i.i.d. $N(0, \boldsymbol{\Sigma}_c)$.

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

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

where $\boldsymbol{Y}_c$ and $\boldsymbol{U}_c$ are $T_c \times N$, $\boldsymbol{X}_c$ are $T_c \times K$, $\boldsymbol{B}_c$ are $K \times N$, $\boldsymbol{Z}_c$ are $T_c \times M_c$ and $\boldsymbol{\Gamma}_c$ are $M_c \times N$. The coefficient matrix $\boldsymbol{B}_c$ is related to coefficients of (1) by: $\boldsymbol{B}_c = [\boldsymbol{B}'_{c1}, \ldots, \boldsymbol{B}'_{cL}, \boldsymbol{\Delta}'_c]'$.

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

The statistical model generating the data is assumed to be as follows. The likelihood for country c has the form:

$$p(\boldsymbol{y}_c | \boldsymbol{\beta}_c, \boldsymbol{\gamma}_c, \Sigma_c) = \mathrm{N}((\boldsymbol{I}_N \otimes \boldsymbol{X}_c)\boldsymbol{\beta}_c + (\boldsymbol{I}_N \otimes \boldsymbol{Z}_c)\boldsymbol{\gamma}_c, \boldsymbol{\Sigma}_c \otimes \boldsymbol{I}_{T_c}) \tag{3}$$

Country coefficients on the variables in $\boldsymbol{X}_c$ are assumed to be drawn from a normal distribution with a common mean $\bar{\boldsymbol{\beta}}$ and a variance $\boldsymbol{\Lambda}_c$ (which may be country specific):

$$p(\boldsymbol{\beta}_c | \bar{\boldsymbol{\beta}}, \boldsymbol{\Lambda}_c) = \mathrm{N}(\bar{\boldsymbol{\beta}}, \boldsymbol{\Lambda}_c) \tag{4}$$

$\boldsymbol{\Lambda}_c$ is discussed in detail in the next subsection.

The prior for $\bar{\boldsymbol{\beta}}$ and $\boldsymbol{\gamma}_c$ is noninformative, uniform on the real line:

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

Alternatively, one could use some informative prior for $\bar{\boldsymbol{\beta}}$, e.g. the Minnesota prior, but this is not necessary for the estimation problem to be well posed, and we do not pursue this possibility here. In fact, it is interesting if the exchangeable prior can substitute the atheoretical Minnesota prior in ensuring that VAR's impulse responses are reasonable in spite of the large number of parameters estimated with few observations. We also use the standard diffuse prior for the error variances:

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

The model (3)-(6) defines the structure advocated in the introduction: the countries' dynamic models of variables in $\boldsymbol{Y}_c$ and exogenous controls in $\boldsymbol{W}$ are special cases of the unknown underlying model defined by $\bar{\boldsymbol{\beta}}$.

The functional form of the prior, the combination of normal, uniform, inverted gamma and a degenerate inverted Wishart (for $\boldsymbol{\Sigma}_c$) densities, is standard, motivated by computational convenience, so that the prior is conditionally conjugate. The posterior density of the model's parameters 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 or inverted Wishart, which enables convenient numerical analysis of the posterior with the Gibbs sampler.[5]

2.2 Specification of the prior variance

What remains to be specified are $\boldsymbol{\Lambda}_c$ - the variances of $\boldsymbol{\beta}_c$ around the common mean $\bar{\boldsymbol{\beta}}$. We would like to incorporate the prior uncertainty with respect to these variances and learn from the data how heterogeneous the countries are. However, we need some constraints on these $NK \times NK$ variances to make the estimation meaningful. In this paper, they are tightly parameterized in a way inspired by the specification of the variance of the Minnesota prior, as discussed in Litterman (1986) or Doan (2000). The off-diagonal terms are set to zero. In the VAR for country c, the coefficient of variable k (where $k = 1...K$ runs over lags of endogenous variables and common exogenous controls) in equation n has a variance equal to

$$\mathrm{var}(\boldsymbol{\beta}_c(k, n)) = \lambda \frac{\hat{\sigma}_{cn}^2}{\hat{\sigma}_{ck}^2} \tag{7}$$

As implied by this formula, all variances are scaled by the common parameter λ. However, some coefficients in $\boldsymbol{\beta}_c$'s are large, and some small. Specifying simply a single variance of λ may imply that some coefficients are allowed to differ from the common mean by a fraction of their own size, and other by orders of magnitude. To adjust for this, as in the Minnesota prior, we scale each coefficient's variance by a factor which adjusts for the size of the coefficient. For a variable denoted by i, $\hat{\sigma}_i$ is computed as the estimated standard error in the univariate Lth-order autoregression for this variable. This standard error captures the scale of unexpected movements of the variable. The same ratio of $\hat{\sigma}_i^2$'s is used in the Minnesota prior, since, as Litterman (1986) argues, what is relevant for the size of a coefficient is the relative size of unexpected movements of the involved variables.

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 $\boldsymbol{\beta}_c$ are assumed to be identical.

[5]The Gibbs sampler for a similar univariate problem is discussed in detail in Gelman et al. (2003).

On the other hand, as λ grows, country models are allowed to differ more and become similar to the respective single country estimates. In the first case the effective number of coefficients we estimate is NK (plus coefficients in $\boldsymbol{\gamma}_c$'s), and in the second case it goes to CNK (plus coefficients in $\boldsymbol{\gamma}_c$'s). We suppose that there is some intermediate range for λ which delivers a reasonable balance between fitting individual countries data on the one hand, and constraining the specification to make the estimates tighter on the other. However, magnitudes of VAR coefficients are hard to interpret: we just do not know what it means in economic terms that two VAR coefficients differ by, say, 1%. Therefore, it is difficult to elicit a meaningful subjective prior for λ.

Fortunately, the problem of specifying noninformative priors for such hypervariances has been studied extensively in the statistical literature. Alternative specifications are discussed e.g. in Gelman et al. (2003) and more recently in Gelman (2006). One convenient form of the prior, which delivers conditional conjugacy, is the inverted gamma density:

$$p(\lambda|s,v) = IG_2 \propto \lambda^{-\frac{v+2}{2}} \exp\left(-\frac{1}{2}\frac{s}{\lambda}\right) \tag{8}$$

The standard noninformative prior for variances, $p(\lambda) \propto 1/\lambda$, which obtains when $v = 0$ and $s = 0$, results in an improper posterior in this setup. The frequently used 'weakly informative' prior with very small but strictly positive values of s and v is shown to be sometimes quite informative in unintended ways, and the results may strongly depend on the chosen values. In the case when there are more than 5 units from which the variance around the common mean is inferred, Gelman (2006) recommends simply using the uniform prior for the standard deviation (which translates into the prior for variance $p(\lambda) \propto \lambda^{-\frac{1}{2}}$), and this recommendation is followed here.

Note that, because λ applies to all $\boldsymbol{\beta}_c$, the effective number of units from which it is estimated is very large: it is not C, the number of countries, but CKN, the number of coefficients in $\boldsymbol{\beta}_1 \dots \boldsymbol{\beta}_C$. Therefore, the sample will contain plenty of information about this parameter, and any truly weakly informative prior will be dominated. It is easy to show that the conditional posterior for λ has the inverse gamma form:

$$\begin{aligned} p(\lambda|Y,\boldsymbol{\beta}_1 \dots \boldsymbol{\beta}_C, \bar{\boldsymbol{\beta}}) \propto \lambda^{-\frac{CNK+v+2}{2}} \\ \exp\left(-\frac{1}{2}\frac{s + \sum_{c=1}^{C}\sum_{k=1}^{K}\sum_{n=1}^{N}(\boldsymbol{\beta}_c(k,n) - \bar{\boldsymbol{\beta}}(k,n))^2/\frac{\hat{\sigma}_{cn}^2}{\hat{\sigma}_{ck}^2}}{\lambda}\right) \end{aligned} \tag{9}$$

The marginal posterior for λ has a nonstandard form, and it has a nonzero value at 0, as the data can never completely rule out the homogeneous case (see Gelman et al., 2003).

The main difference between this paper's approach and that of Canova (2005), who works with a similar model, in that he does not use a prior for his parameters, which play the same role as λ, nor does he analyze their posterior distribution. Rather, he conditions on their values found by maximizing the predictive density in a subsample. He is working with a single panel, while the focus of this paper is to compare two panels. Here we want to properly account for the posterior uncertainty about the heterogeneity within each panel in order to make meaningful comparisons of heterogeneities across panels. Also, he does not adjust the prior variance to account for different magnitudes of coefficients, like we do in equation (7).

2.3 Identification of monetary policy shocks

Identification of the structural model assumes a small open economy with the exchange rate flexible enough to react immediately to monetary policy, and monetary policy that reacts immediately to exchange rate movements. This assumption remains valid also in the presence of managed exchange rates with target bands, like the European Union's ERM, and in the arrangements in the NMS in our sample, 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 the exchange rate to respond immediately to the interest rate, and vice versa, tend to produce a price puzzle, i.e. an initially positive response of prices to monetary policy tightening (for more on this subject see e.g. Kim and Roubini, 2000).

The endogenous variables in the VARs are: a measure of real output, the consumer price level, the short-term interest rate and the exchange rate in national currency units per foreign currency unit, all measured at monthly frequency (more details about the series and their transformations are given in the next section and in the appendix). A money aggregate is not included in the baseline specification: 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 monetary aggregate fluctuations also carry information about money demand. Since identifying money demand is beyond the scope of this paper, we conserve degrees of freedom and do not include money aggregates in the baseline specification. We also follow the common practice of not including fiscal variables, which is justified in monetary VARs as long as fiscal policy is passive (in the sense of Leeper, 1991). This is no more controversial for the NMS than for the euro area countries. It is reasonable to assume that in both regions and both sample periods, monetary policy focused on maintaining low inflation in a manner independent of fiscal policy.

In order not to confuse domestic monetary policy shocks with the central banks' responses to external developments, specifications include several foreign variables which are treated as exogenous. These include world-wide variables (commodity prices and the US Fed Funds rate) and German variables (interest rate, output and the exchange rate), reflecting Germany's importance as an export market and the Bundesbank's leading role in the region.

As is standard in the identified VAR literature, we assume that structural shocks are orthogonal, which means that 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 $\boldsymbol{\upsilon}_{ct}$ and the vector of VAR innovations $\boldsymbol{u}_{ct}$ is as follows:

$$\boldsymbol{G}_c \boldsymbol{\upsilon}_{ct} = \boldsymbol{u}_{ct} \tag{10}$$

where $\mathrm{var}(\boldsymbol{\upsilon}_{ct}) = \boldsymbol{I}_N$ (identity matrix of order N) and $\mathrm{var}(\boldsymbol{u}_{ct}) = \boldsymbol{\Sigma}_c = \boldsymbol{G}_c \boldsymbol{G}_c'$. Therefore, the identification involves finding an appropriate factorization $\boldsymbol{G}_c$ of the residual covariance matrix. The monetary policy shock is the only shock identified here. We identify it with the following two assumptions:

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.

These assumptions are summarized in the scheme below:

$$\begin{pmatrix} + & 0 & 0 & 0 \\ \bullet & + & 0 & 0 \\ \bullet & \bullet & + & + \\ \bullet & \bullet & - & + \end{pmatrix} \begin{pmatrix} \upsilon_{ct1} \\ \upsilon_{ct2} \\ \widehat{\upsilon_{ct3}} \\ \upsilon_{ct4} \end{pmatrix} = \begin{pmatrix} u_{ct1} \\ u_{ct2} \\ u_{ct3} \\ u_{ct4} \end{pmatrix} \begin{array}{l} \leftarrow \text{output innovation} \\ \leftarrow \text{price innovation} \\ \leftarrow \text{interest rate innovation} \\ \leftarrow \text{exchange rate innovation} \end{array} \tag{11}$$

where + denotes coefficients that are constrained to be positive, 0 denotes zero restrictions and $\bullet$ denotes unconstrained coefficients. $\widehat{\upsilon_{ct3}}$ is the monetary policy shock. The triangular form of the upper left block of the $\boldsymbol{G}_c$ matrix reflects a normalization, which has no effect on the impulse responses to the monetary policy shock.

Assumption 1 is standard in the VAR literature. Assumption 2 is justified by a standard mechanism: an exogenous interest rate increase by the central bank makes local currency assets more attractive and immediately leads to capital inflows, and appreciation of the currency. An opposite comovement is also theoretically possible. The key to the alternative mechanism is that higher interest rates increase the cost of debt servicing. When the risk of default is sufficiently high initially or when the interest rate increase is sufficiently large, country's assets may actually become less attractive, and a capital outflow and an exchange rate depreciation would follow. Blanchard (2004) studies an episode in Brazil in which this second mechanism may have switched on. We cannot rule out the presence of such episodes in the analyzed sample, but on average such perverse effects of monetary policy are not a concern of both region's central banks.

The negative comovement of the interest rate and the exchange rate distinguishes monetary policy shocks from all other shocks, which involve the opposite comovement. Other shocks, which may have both internal and external origin, cause capital to flow e.g. into the country, bidding up the exchange rate and bidding down interest rates, or forcing the central bank to lower the interest rates to prevent further currency appreciation.[6]

The precise elasticities of interest rates to exchange rates and vice versa, conditional on both shocks, are not known, and the sign restriction discriminates between shocks only approximately. This is captured in the appropriately wide uncertainty bands around impulse responses.

Technically, factorizations satisfying (11) are obtained by multiplying the Choleski factor of the residual covariance matrix by matrices of the form

$$\begin{pmatrix} \boldsymbol{I}_2 & 0 \\ 0 & \boldsymbol{V} \end{pmatrix}$$

where $\boldsymbol{V}$ is a 2×2 random orthonormal matrix, until a draw is found that complies with the restrictions. The approach follows Uhlig (2005) and, as discussed there, from the Bayesian point of view it amounts to multiplying the prior discussed in the previous section by a uniform prior on the space of orthonormal matrices times an indicator function, selecting the orthonormal matrices which deliver the restriction.

[6]Many central banks implicitly target some Monetary Conditions Index (MCI), i.e. a weighted average of the interest rate and the exchange rate. Investment banks' analyses and IMF country reports routinely analyze monetary policy stance with the help of MCI's, even though few central banks admit to targeting them explicitly.

3 Empirical Results

3.1 The baseline specification

We analyze two panels of countries: five euro area countries (EA5) and four of the New Member States (NMS4) of the European Union. The reference point for the NMS countries are the EA countries in the years 1987-1998, i.e. *in the period before the euro adoption.* The goal of the paper is to compare the two regions within the same econometric framework, and therefore we include only countries with a flexible (i.e. not fixed) exchange rate, to which the same identification scheme is applicable.

The EA5 panel consists of Finland, France, Italy, Portugal and Spain. Austria, Belgium and the Netherlands were excluded because of their quasi-fixed exchange rate against the D-mark[7], Ireland because of the lack of monthly CPI data, and Greece because of the lack of interbank interest rates for much of the sample. Finally, Germany is excluded because of its special status as both regions' economic locomotive, with the leading role of the Bundesbank. Instead, German variables are included as exogenous controls for all countries.

The NMS4 panel consists of the Czech Republic, Hungary, Poland and Slovenia. Bulgaria and the Baltic countries were excluded because they have had currency boards. In Romania and Slovakia, for much of the sample, market interest rates fluctuate strongly and 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.

The sample periods for the NMS4 countries span the period from the second half of the 1990s up to late 2007 and differ for each country, depending on when the post-transition exchange rate controls were relaxed. The information on the chronology of the exchange rate regimes and other developments was taken from various issues of IMF Country Reports. The sample for the Czech Republic starts in January 1998, with the introduction of inflation targeting (this avoids the Czech currency crisis of 1997). The sample for Hungary begins in April 1995, and for Poland in May 1995, when both countries introduced crawling band regimes. The sample for Slovenia

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

starts in May 1997, when it introduced a new monetary framework, which is considered to be a regime shift. For the euro area countries, we consider a sample similar in length to that of the NMS, covering 1987 to 1998 (to the introduction of the euro).

The data is monthly. The endogenous variables - output, prices, interest rates and exchange rates - are measured respectively by the log of the index of industrial production, log of the consumer price index, the short-term market interest rate and the log of the exchange rate in national currency units per euro. Before 1998 the exchange rate is that of the D-mark multiplied by the official conversion rate of the D-mark to the euro. Two groups of exogenous variables are included to control for world-wide developments and developments in Germany. World factors are represented by the Federal Funds rate, oil prices and non-fuel commodity prices (converted to euros). German developments are represented by the index of industrial production in Germany, the interest rate on the German interbank market and the euro(D-mark)/USD exchange rate. Data have mostly supported the specification in which the coefficients of the world variables are shrunk with the exchangeable prior and the coefficients of German variables are excluded from the exchangeable prior. Data sources are provided in the appendix.

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: 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.

3.2 Impulse responses

We simulate the posterior distribution of the coefficients using the Gibbs sampler.[8] Figure 1 presents the 5th, 50th and 95th percentiles of the posterior distribution of impulse responses to a one standard deviation monetary policy shock, in the mean model for each panel. This is the model implied by $(\bar{\boldsymbol{\beta}})$,

[8] The results are based on 1,000 draws from the posterior obtained from a sequence of 1,010,000 draws, from which we discard the first 10,000, and save every 1000th draw from the remaining sequence. Draws are very autocorrelated, and therefore saving only every 1000th was deemed to be optimal for storage considerations. We establish convergence using the Geweke (1992) diagnostics for convergence of a Markov chain implemented in the coda R package (Plummer et al., 2007). Sequences initialized with the country-by-country point estimates and those initialized with perturbed fully pooled estimates converge to the same distributions.

and the mean residual variance, which for each draw is computed from the residuals $\boldsymbol{U}_c$ of all countries. Responses are shown for three years (36 months) after the shock.

The immediate (period 0) responses of all variables simply reflect the identifying assumptions: 1) in the month of the shock, output and prices are unaffected, and 2) the interest rate rises and the exchange rate falls (appreciates). The identified monetary policy shock is associated with a median interest rate increase of 45 basis points in the NMS4 and almost 60 basis points in the EA5. The median initial appreciation is similar in both panels, around 1.2%. The interest rate increase is reversed after about one year, while the appreciation persists for about 1.5 years. The economies respond with a transitory output decline and a possibly permanent reduction of the price level.

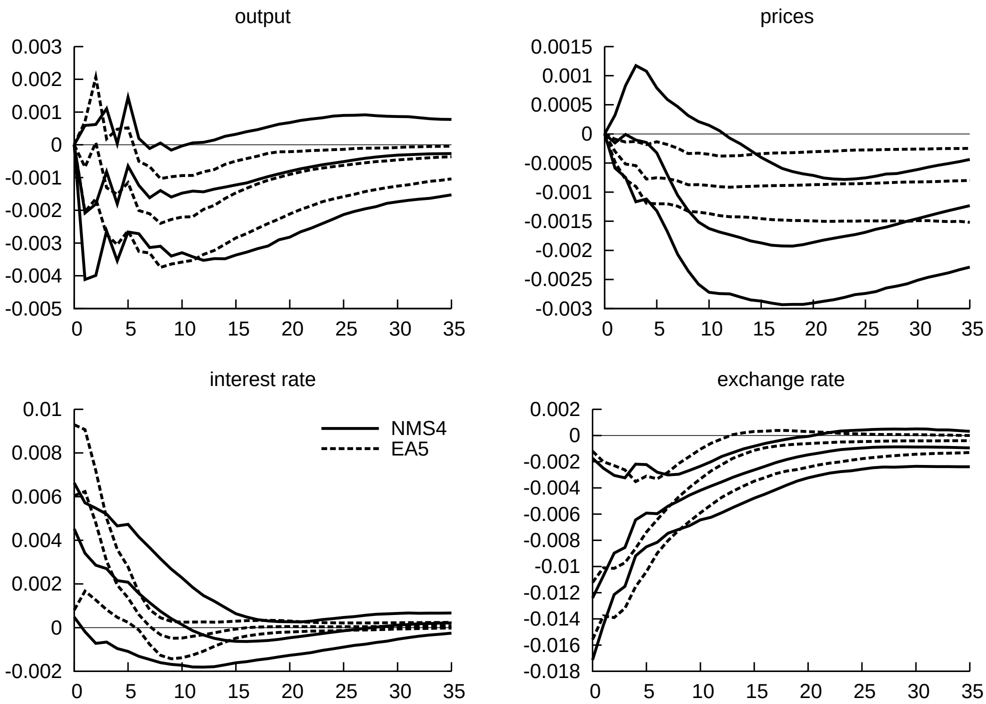


Figure 1: Mean impulse responses to monetary policy shocks for the NMS4 and EA5: median, 5th and 95th percentiles of the posterior distribution

Overall, the impulse responses are similar. Output responses to a contractionary monetary policy shock are negative and transitory (they last about two years), and price level responses are negative and very persistent.

The uncertainty bands largely overlap. It is reassuring that, in spite of the differences in the financial structures and often stressed peculiarities of the transition from the centrally planned economy, basic economic models seem to be applicable in both regions. Impulse responses of output and prices in the NMS4 are estimated with less precision. The uncertainty bands for the NMS4 include, for horizons longer than six months, the possibility of much stronger price responses than in the EA5. In the short run, the comparison is reversed, and the point estimates of the NMS4 price responses are weaker, but this comparison is less statistically significant than the comparison for longer horizons. The median price level response to a standard monetary policy shock after 18 months is 0.2% in the NMS4, and only less than 0.1% in the EA5. The median output response is at most 0.2% in both panels.

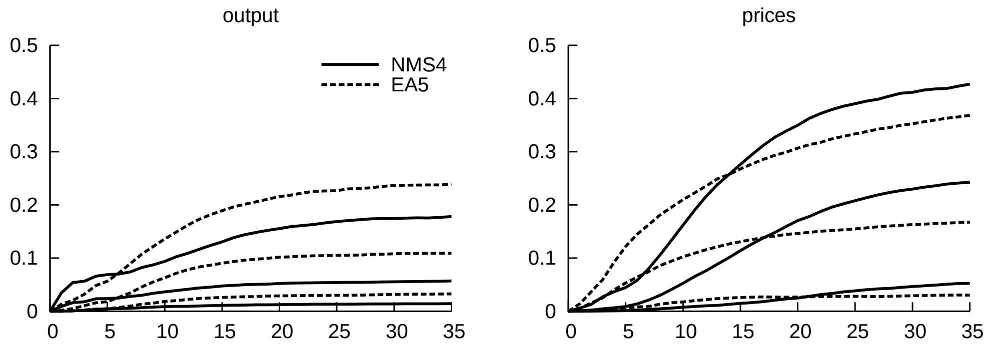


Figure 2: Shares of monetary policy shocks in variance decompositions of output and prices in the NMS4 and EA5: median, 5th and 95th percentiles of the posterior distribution

Figure 2 presents contributions of monetary policy shocks to forecast variance of output and prices in both regions, for horizons up to three years. According to the median estimates, at the horizon of 3 years, monetary policy shocks account for about 24% of the variance of prices in the NMS4, compared with 17% in the EA5. For output, monetary policy shocks contribute about 6% of the variance in the NMS4 and 11% in the EA5. These shares are quite high by the standards of the literature. We interpret this as an indirect support for the employed identification scheme: shocks satisfying our assumptions are indeed present in the data, and may account for an important share of its variability.

The observation that output responses tend to be similar, while those of prices are stronger in the NMS4, suggests that the output cost of disinflation facing the NMS4 central bankers might be lower. We examine this closer by comparing across regions the output cost of constraining price level growth with an exogenous monetary policy tightening. We calculate this cost as follows:

$$\text{output cost of disinflation (at horizon H)} = \frac{\frac{1}{H}\sum_{h=1}^{H}\psi_h^y}{\psi_H^p}$$

where ψ_h^x denotes the response of variable x (where x is either y - output, or p - price level) to a monetary policy shock after h periods. While interpreting this statistic, it is best to think of impulse responses as deviations from the unconditional forecast (i.e. a forecast computed under the assumption of no shocks occurring). The numerator shows the average output loss in the H months after a standard contractionary monetary policy shock, compared with the scenario without the shock. The denominator shows the difference of the price level observed H months after the shock, and the price level which would have been observed in the absence of the shock. [9]

Figure 3 presents the posterior distribution of such cost of disinflation for the two-year horizon. This figure shows that in the NMS4, for the price level to grow by one percent less than in the unconditional forecast in a two year period, the authorities need to consent to the output path which is about 0.5 of a percent below the unconditional forecast on average during these two years. In the EA5 such an output cost will amount to around 1.5 percent.

Another finding in Figure 1 is that a standard monetary policy shock is associated with a stronger movement of the interest rate in the EA5 countries. We do not emphasize this finding, because, as shown below in the individual country results and in the section on robustness, the size of the average monetary policy shock may be sensitive to the sample.

Figure 4 presents impulse responses to a monetary policy shock for each of the analyzed countries. In this subsection we focus on the impulse responses obtained with the exchangeable prior, whose 5th and 95th percentiles are represented by the continuous line. (Impulse responses obtained without the exchangeable prior, represented by gray areas, are discussed in the next subsection.) Country impulse responses obtained with the exchangeable prior differ mainly by the scale of the shock, which is country specific. Their shapes are similar and close to those of the mean impulse responses, suggesting that

[9]Sacrifice ratios calculated from VARs are studied in Cecchetti and Rich (2001). Their calculation is based on a VAR with the second difference of price level, which makes it possible to define the sacrifice ratio as the output cost of a *permanent* reduction in inflation. The VAR in this paper does not impose a unit root in the inflation process, nor even in the price level process, so the output cost of affecting prices needs to be defined differently.

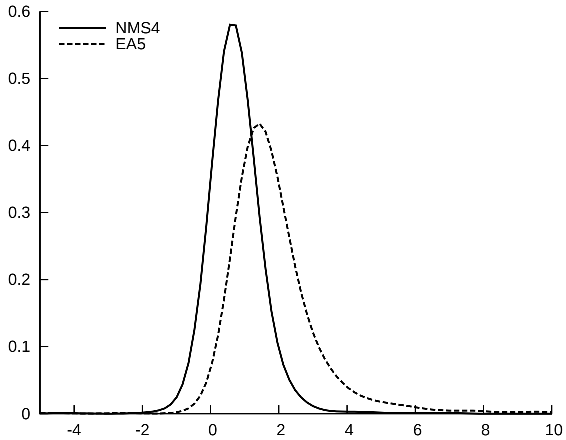


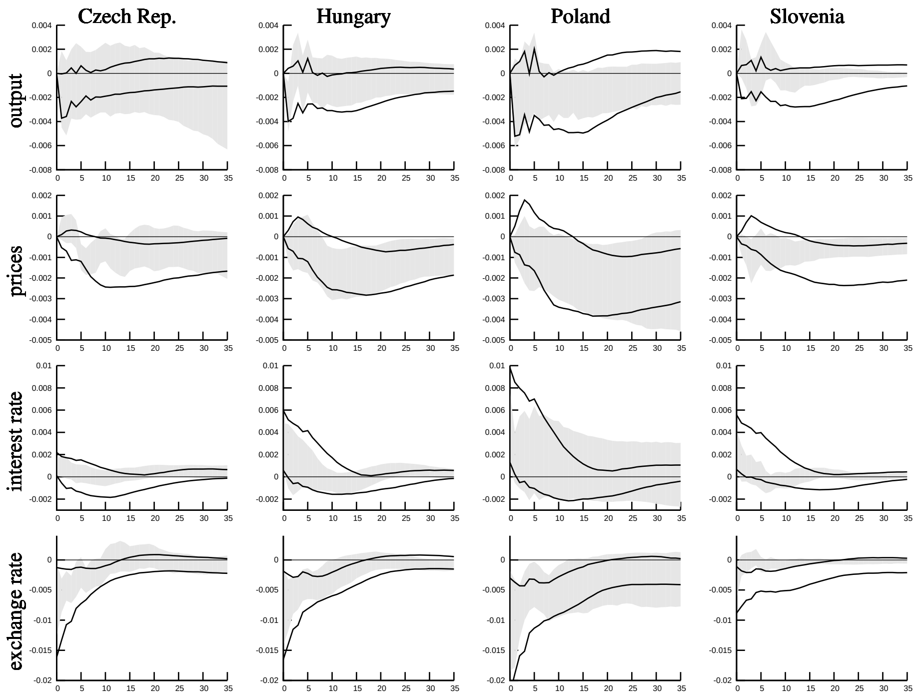
Figure 3: Posterior distribution of the output cost of disinflation (24 months horizon)

the posterior is relatively close to the panel VAR specification, but some caveats to this conclusion will be discussed in the next subsection.

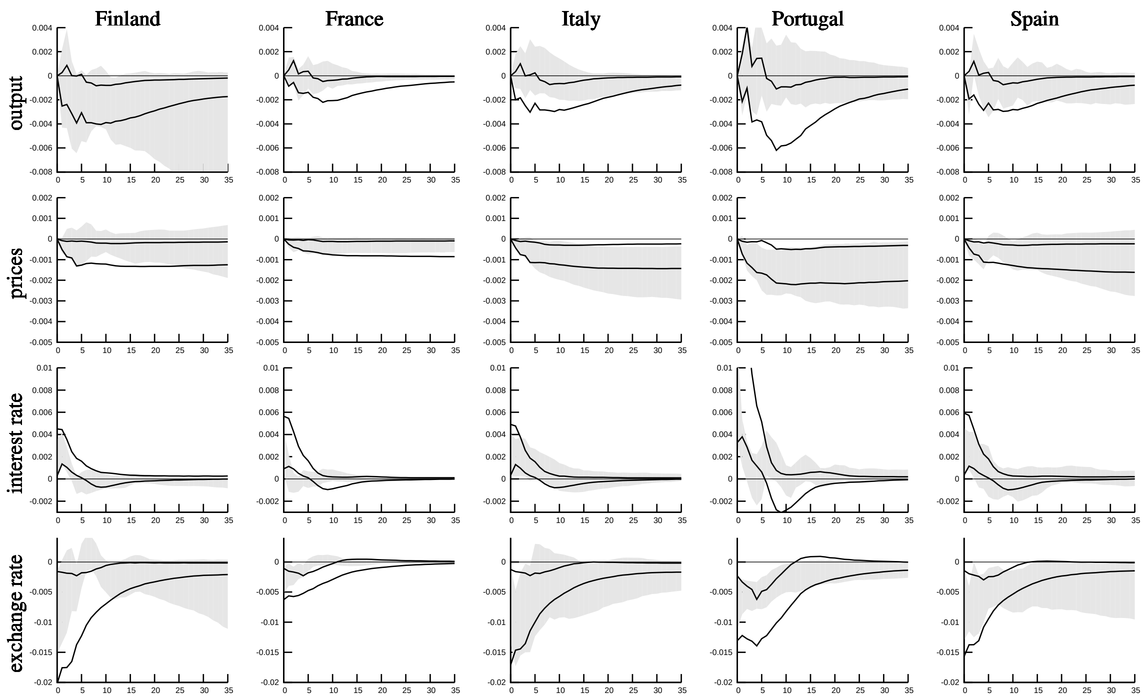
The amplitudes of monetary policy shocks vary widely. In the NMS4 panel, the 95th percentile of the immediate interest rate response is as much as 100 basis points in Poland, and only 20 basis points in the Czech Republic. The 5th percentile of the immediate exchange rate response is as much as 2% in Poland, and less than 1% in Slovenia. The response of prices in the Czech Republic is most statistically significant: the uncertainty bands put very little weight on positive initial price responses. In other countries price responses have considerable probability mass on both sides of zero in the first months. In Slovenia, they are also weakest. Impulse responses of Hungary and Poland are most similar in shape to the mean responses, which is intuitive, since they have the longest samples and thus, are likely to have the highest weight in the posterior.

In the EA5, Portugal is an outlier in terms of the high volatility of its interest rate movements and, in the medium term, its output responses. However, when we repeat the estimation dropping Portugal, the mean size of the interest rate is much lower, but all the remaining comparisons with the NMS4 are unaffected. In Portugal and France, exchange rates move little relative to the interest rate movements. Otherwise, the qualitative features of the impulse responses are similar.

The first lesson from these results is that, when comparing 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



A. NMS4 panel



B. EA5 panel

Figure 4: Impulse responses to monetary policy shocks in the NMS4 and the EA5, 5th and 95th percentiles of the posterior. Continuous line: results with the exchangeable prior; shaded regions: country-by-country estimation.

countries. After a few months' lag, a standard monetary policy shock is likely to have an even stronger impact on prices in the NMS4 than in the EA5. The precise comparison the effect of *equal* monetary policy shocks is tricky to conduct, because the monetary policy shock involves a simultaneous movement of two variables: the interest rate and the exchange rate, and their relative response is different. However, it is clear from Figure 1 that whether we normalize the shocks to involve the same movement of the interest rate, or of some linear combination of the interest rate and the exchange rate, the effect on prices is, in the medium term, stronger in the NMS4.

Second, the output cost of disinflation is lower in the NMS4 than in the EA5. This is consistent with the prediction of economic theory, that the Phillips curve is steeper in countries with higher and more volatile inflation.

3.3 Effects of the exchangeable prior

In this subsection we compare the posterior obtained with the exchangeable prior to the posteriors estimated under two limiting cases: no pooling (country-by-country estimation) and full pooling (panel VAR estimation). First, to characterize the tradeoffs involved in shrinking the coefficients, we look at the indicators of model fit and overparameterization proposed in Spiegelhalter et al. (2002) (see also Gelman et al., 2003, pp.179-184). Then we comment on the effects of scaling prior variances in equation (7), and report the posterior distribution of λ. Finally, we discuss how impulse responses to monetary policy shocks estimated with no pooling and with full pooling differ from those estimated with the exchangeable prior.

Table 2 reports statistics for model comparison introduced by Spiegelhalter et al. (2002). The first statistic is the expected deviance $\hat{D}_{\text{avg}}$, which measures the fit of the model to the data on average over the posterior distribution.[10] The expected deviance is related to the mean squared error, and therefore the smaller it is, the better is the fit, and it is not comparable across samples. For both the NMS4 and the EA5 panels, the fit, as measured by this statistic, is best for the country-by-country estimation (-14,114 and -20,384 respectively) and worst for the panel VAR (-13,792 and -19,827 respectively). The exchangeable prior is in between: it has a better fit than the panel VAR, but a worse fit than the models estimated individually for countries.

[10] Deviance is defined as $D(y, \theta) = -2 \log p(y|\theta)$, where y denotes data and θ denotes parameters. For the normal model the deviance is proportional to the mean squared error. The expected deviance is computed as the deviance averaged over the posterior distribution of the parameters.

Table 2: Summary statistics for alternative model specifications

	D_{avg}	parameter count†	p_D	DIC
NMS4				
country-by-country	-14114	632	633.1	-13481
pooled	-13792	272	263.4	-13529
exchangeable prior	-13819	632	278.4	-13540
exchangeable prior (no scaling)	-13797	632	266.3	-13531
EA5				
country-by-country	-20384	790	786.2	-19597
pooled	-19827	310	300.8	-19526
exchangeable prior	-19959	790	327.5	-19631
exchangeable prior (no scaling)	-19960	790	339.1	-19621

†The parameters for each country were counted as follows: coefficients of lagged dependent variables = 4 equations × 4 variables × 6 lags; coefficients of exogenous controls to which the exchangeable prior applies = 4 equations × 3 variables × 2 lags; coefficients of exogenous controls to which the exchangeable prior does not apply = 4 equations × 3 variables × 2 lags + 4 constant terms; unique terms in the 4×4 variance matrix = 10.

The second statistic is the effective number of parameters p_D. For comparison, we also provide the simple count of parameters explicitly entering the likelihood function. The simple count of parameters in a Bayesian model can easily exceed the number of observations, but this does not mean that such a model has negative degrees of freedom, because the prior information also carries information about parameters. p_D is a way to measure the number of free parameters effectively estimated from the data. When the prior is noninformative, p_D approximately corresponds to the parameter count (the exact equality holds in simple special cases, e.g. in a linear normal model with known variances), but when the prior is informative, it will be lower (and need not be an integer). In a hierarchical model this measure depends on the 'focus' on the level of parameter hierarchy, and here we focus on the parameters of the individual countries' VARs (not on the regional means). See Spiegelhalter et al. (2002) for a detailed discussion and derivation of this measure.

As shown in Table 2, the estimates of the effective number of parameters for the models with the flat prior (the country-by-country specification and the fully pooled specification) are indeed quite close to the actual count of the parameters. The estimate p_D is particularly close to the parameter count in the country-by-country estimation for the NMS4 (633.1 vs 632). In the remaining cases the estimated effective number of parameters is somewhat smaller then the parameter count (263.4 vs 272, 786.2 vs 790, and 300.8 vs 310). In both regions, the effective number of parameters in the model with the exchangeable prior (278.4 for the NMS4 and 327.5 for the EA5) is much smaller than the parameter count (632 and 790), and close to that in the panel VAR (272 and 310). This confirms the conclusion from comparing the individual country impulse responses in Figure 4, that the a-posteriori optimal strength of shrinkage is quite strong, and quite close to assuming homogeneity within panels.

The third statistic, the Deviance Information Criterion (DIC) of Spiegelhalter et al. (2002), summarizes the tradeoff between the fit and overparameterization. The DIC is simply the sum of the expected deviance and the effective number of parameters, and obviously the smaller it is, the better the model. It is designed to pick the model with the best out-of-sample predictive power in the included countries. We have seen that the exchangeable prior allows us to achieve a reasonable fit while estimating effectively less than half of the parameters needed in the unconstrained estimation. As shown in the last column of Table 2, in both regions this gives it an advantage in terms of the DIC.

An alternative to the model-comparison criteria discussed here would be to introduce informative priors and compute Bayesian odds ratios for alternative models. In the presence of good prior information this would certainly be advantageous, but in the present context, one would have to resort to using 'weakly informative' priors in many dimensions. This tends to make odds ratios sensitive to the particular choice of the 'weakly informative' priors and, because of this, is discouraged by some authors (see e.g. Gelman et al., 2003, example on pp.185-186).

A separate line in the table reports, for each region, results for the exchangeable prior used without taking into account the scaling terms from equation (7), which reflect the magnitudes of coefficients. That is, equation (7) is replaced with the relation: $\mathrm{var}\left(\boldsymbol{\beta}_c(k,n)\right) = \lambda$. This could, potentially, make a large differences for the results. The omitted scaling terms $\frac{\hat{\sigma}_{cn}^2}{\hat{\sigma}_{ck}^2}$ are equal to 1 for the coefficients of own lags of a variable, but, in our data, $\hat{\sigma}^2$ of e.g. oil price exceeds those of interest rates by orders of magnitude. In both the NMS4 panel and the EA5 panel, this alternative specification produces a lower DIC than in the baseline specification, but the difference is small. In the NMS4 it leads to a somewhat stronger posterior shrinkage, and in the EA5 to a weaker posterior shrinkage than in the model with scaling. This lack of consistent pattern is intuitive, because scaling does not matter

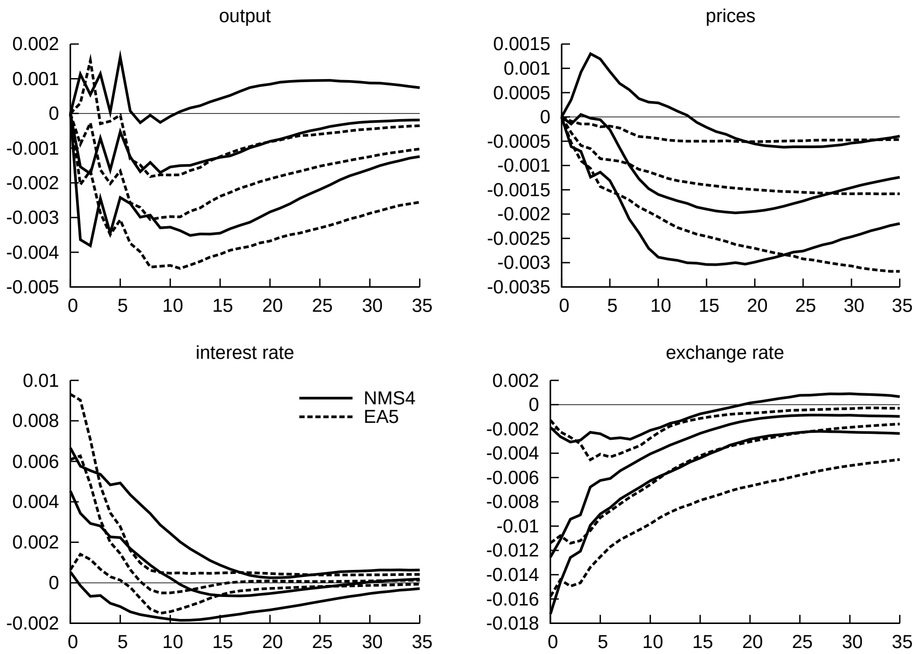


Figure 5: Mean impulse responses to monetary policy shocks for the NMS4 and EA5: median, 5th and 95th percentiles of the posterior distribution. No scaling of λ.

in the two extreme cases: no pooling and full pooling. Consequently, if the scaling in the prior differs from the scaling in the data generating process, the posterior degree of shrinking can end up being either stronger or weaker than in the data generating process.

Figure 5 shows impulse responses to a monetary policy shock estimated with this specification of the prior variance. Impulse responses in the NMS4 are hardly affected, but both price and output impulse responses in the EA5 are stronger. Therefore, with this specification we would not conclude that the response of prices in the NMS4 is stronger than in the EA5. However, the comparison of the output costs of disinflation shown in Figure 6 is not affected. Therefore, under this specification the output cost of disinflation is still lower in the high-inflation NMS4 panel than in the low-inflation EA5 panel, as predicted by economic theory.

We return to the baseline specification, using the scaling in equation (7). We have shown how different degrees of information pooling across countries affect fit and overparameterization. Now we turn to studying their effect on the substantive conclusions that can be drawn about monetary transmission.

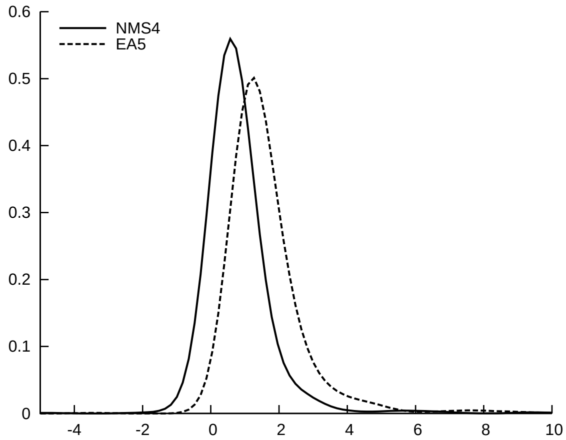


Figure 6: Posterior distribution of the output cost of disinflation (24-month horizon). No scaling of λ.

Figure 7 illustrates the dangers of applying excessive shrinking. It compares the euro area panel's mean impulse responses obtained with the exchangeable prior against those obtained with full pooling, i.e. with the panel VAR model. The panel VAR results exhibit much more persistence, which is often an artifact of imposing homogeneity on a heterogeneous panel, as shown in Pesaran and Smith (1995). Most strikingly, the effect of monetary policy shocks on output is permanent in the panel estimation (whereas it dies away in the exchangeable prior estimation), with dramatic implications for the implied output cost of disinflation.[11] Mean impulse responses for the NMS4 panel turn out to be very similar with full pooling and with the exchangeable prior, so they are not reported for brevity's sake.[12]

Figure 4 allows to compare individual countries' impulse responses to monetary policy shocks obtained with the exchangeable prior against those obtained with country-by-country estimation (which are represented with shaded areas). VARs estimated on individual country data are heavily over-parameterized, given the length of the available samples. For individual countries we have only 3.1 to 4.1 observations per estimated parameter in

[11]Interestingly, the VAR results for individual countries in Mojon and Peersman (2001) and for the euro area in Peersman and Smets (2001) show a similar phenomenon: output responses to monetary policy shock die out in all countries except France, but in the estimation on the euro area data they are permanent (Mojon and Peersman, 2001, Graph 1a-b).

[12]However, in the shorter samples studied in the next subsection, also the NMS4 full pooling and exchangeable prior results differ importantly.

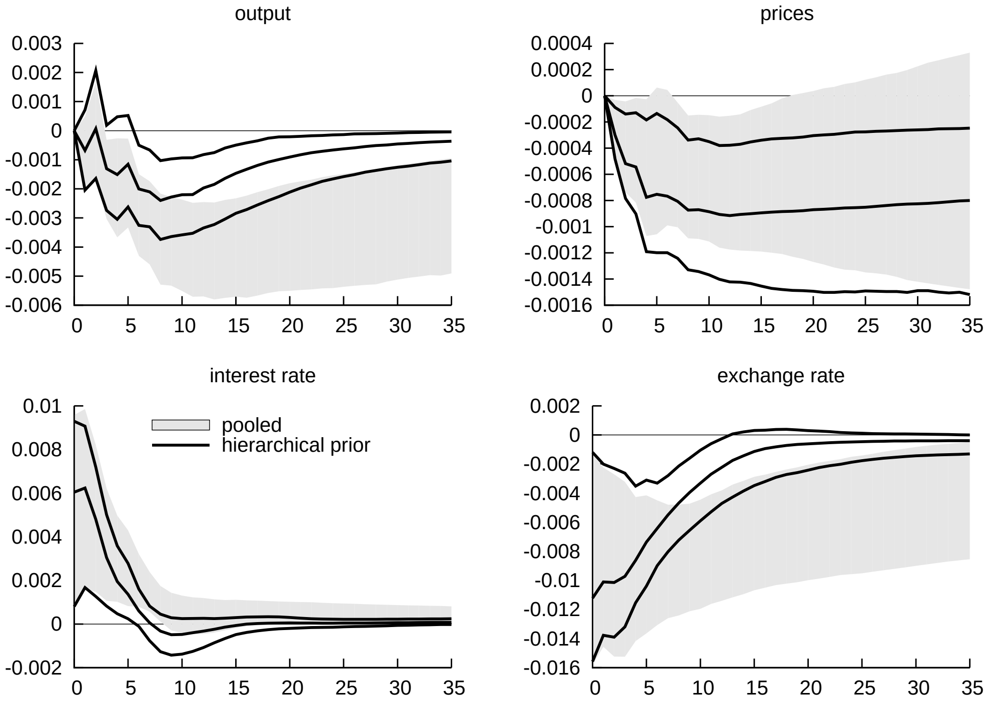


Figure 7: Impulse responses in the EA5 panel: exchangeable prior and full pooling

the NMS4, and 3.9 in the EA5. Results based on such short data have to be treated with caution. As often found in this situation, several impulse responses exhibit cycling behavior which is probably spurious (especially: price response in the Czech Republic, output response in Slovenia, interest rate response in Portugal, and exchange rate response in Finland). Many impulse responses, especially those of output, are insignificant. Finally, at longer horizons the responses sometimes diverge (e.g. output responses in the Czech Republic and Finland), which implies a nonstationary model, or collapse to zero (responses of Slovenia and France) which implies a high degree of certainty that the model is stationary with little persistence. Impulse responses obtained with the exchangeable prior exhibit fewer of these extreme phenomena and greatly facilitate comparisons across panels.

The strength of shrinkage is governed by the overall tightness parameter λ. Figure 8 presents the posterior densities of the square root of this parameter in each of the panels. Taking the tightness of the well known Minnesota prior as the benchmark, the posterior mass is concentrated on rather low val-

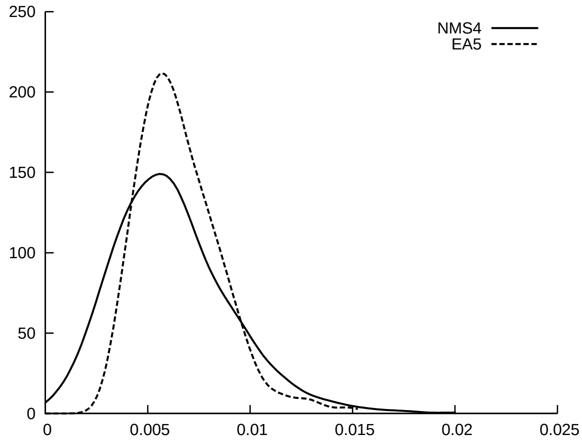


Figure 8: Posterior distribution of $\sqrt{\lambda}$ in the NMS4 panel and in the EA5 panel

ues.[13] This might give an impression that the posterior is close to the panel VAR model, and indeed this is what we see comparing impulse responses of individual countries. However, the size of $\sqrt{\lambda}$ is tricky to interpret. In large time series models with roots near unity, small differences in many individual coefficients can results in large differences in impulse responses. As shown in Figure 7, in the EA5 panel excessive shrinking distorts the impulse responses significantly. Therefore, the low posterior probability of $\sqrt{\lambda}$ smaller than 0.0025 in the EA5 panel (seen in Figure 8) is important for the results.

An alternative to using the flat prior for λ endorsed by Gelman (2006), is to use 'weakly informative' priors, but some care is needed in this case. Natural 'weakly informative' priors for λ are of inverse-gamma form with small scale and degrees of freedom. We have found that priors $\mathrm{IG}_2(\epsilon, \epsilon)$ with $\epsilon \leq 0.001$ are indeed neutral, and produce similar results as the flat prior. However, although inverse gamma densities with low degrees of freedom have a relatively fat right tail, their left tail is thin and, as a result, they are actually very informative in ruling out values in the vicinity of zero.[14] Therefore,

[13] $\sqrt{\lambda}$ corresponds to the 'overall tightness parameter' of the Minnesota prior. The RATS manual (Doan, 2000) suggests, as a benchmark, values of this parameter of 0.2 to 0.1. In contrast, most of the posterior mass of $\sqrt{\lambda}$ lays below the value of 0.01, so the posterior is more than ten times tighter. However, the Minnesota prior is centered at the random walk model which is a very crude description of the data. In contrast, the posterior λ shows the tightness around the posterior mean model $\bar{\boldsymbol{\beta}}$ which is fitted to the analyzed data and which is random itself.

[14] This is a different problem with 'weakly informative' inverse gamma densities than

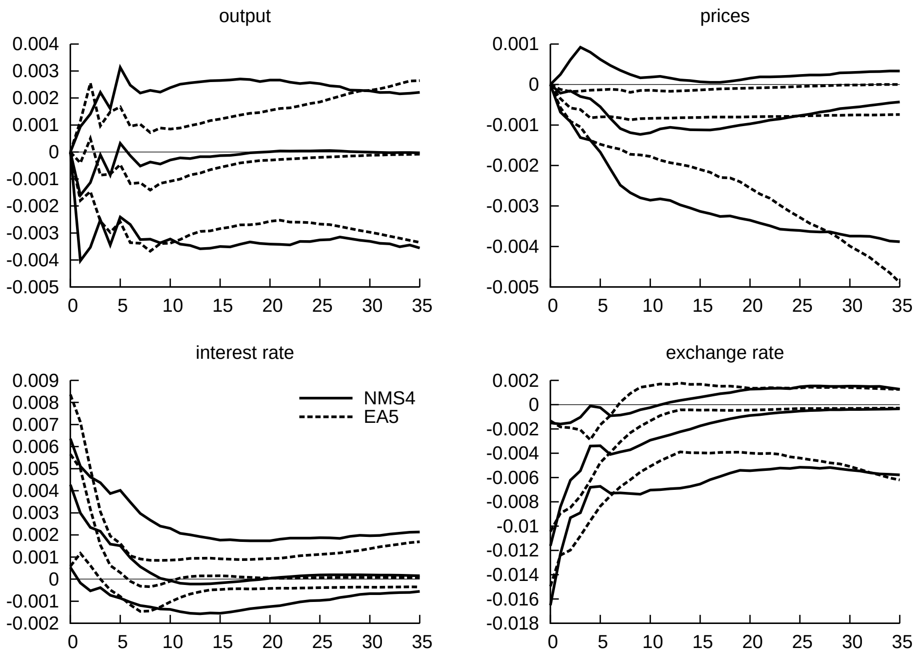


Figure 9: Mean impulse responses to monetary policy shocks for the NMS4 and EA5: median, 5th and 95th percentiles of the posterior distribution. Prior $p(\lambda) = \mathrm{IG}_2(0.1, 0.1)$.

e.g. the prior $\mathrm{IG}_2(0.1, 0.1)$ produces posteriors of $\sqrt{\lambda}$ which support much less shrinkage than in the benchmark case. They are peaked at 0.05 and they have almost no mass below 0.04. The resulting mean impulse responses are presented in Figure 9. Their uncertainty bands are wider and no significant differences between NMS4 and EA5 can be seen.

The previous discussion gives an idea of the sensitivity of the results to informative priors for λ. Informative priors that support strong shrinkage do not affect the NMS4 results much, but bias the EA5 results towards those in Figure 7. This sharpens the difference in the output cost of disinflation across regions, observed in the benchmark specification, and reinforces our conclusions. On the other hand, informative priors that support weak shrinkage result in less significant results, in which any differences between NMS4 and EA5 are blurred.

that stressed in Gelman (2006). Gelman considers a case where the shrinking parameter is likely to be quite large and is estimated with very few units. In his case, even the right tail of the prior is thin enough to have a strong effect on the results. In our example, the problem is the left tail.

3.4 Robustness checks

In this subsection we report a number of robustness checks. First, we check how the NMS results change across the high- and low-inflation subsamples. Then we study the importance of the most volatile sample periods for the estimated results. Subsequently, we report the results of other modifications of the baseline specification.

We have found that the output cost of disinflation was lower in the NMS4 and we have interpreted this finding as a result of their higher inflation rates and/or more volatile conditions. If this interpretation is correct, one should expect higher output costs of disinflation in the more recent subsamples, when inflation in NMS4 has been more in line with inflation observed in EA5 (see Table 1). To check this, we repeat the estimation in the subsamples starting in January 1999 and January 2000.

Figure 10 reports responses of output and prices in the NMS4 in the full sample, and in the two shorter subsamples. Full sample results for the EA5 panel are added in each plot for comparison. Interest rate and exchange rate responses, which are not shown, are slightly weaker in the later samples. The initial interest rate response is 38 basis points in the sample starting in 1999 and 34 basis points in the sample starting in 2000. All plots (including those for the EA5) show responses rescaled to a 50 basis points shock. This may improve their comparability, although it should be kept in mind that the NMS shock is associated with a stronger exchange rate movement, relative to the interest rate, than the EA5 shock. Therefore, effectively, it involves tighter monetary conditions.

As seen in Figure 10, in the later samples the NMS output responses become somewhat stronger and price responses become weaker. Samples are short and the uncertainty is high, so that individually these changes are not very statistically significant. However, the distributions of the output cost of disinflation presented in Figure 11 confirm that output cost of disinflation in the NMS4 becomes higher, and similar to that in the EA5. Therefore, this evidence is consistent with output cost of disinflation increasing when inflation falls, and with the NMS4 and EA5 becoming more similar over time.

An important link in the monetary policy transmission mechanism, and a potential source of relevant heterogeneity between EMU and NMS, is the response of long term rates to monetary policy. However, markets for long term bonds were developed only in the new millennium in most of the NMS and 5-dimensional VARs on the resulting shorter samples are estimated with too much uncertainty to allow any meaningful comparisons.

The next question we address is whether the baseline results are not driven by periods of exceptional volatility in financial markets. In general volatile periods are a boon for the econometrician, because they carry precious information: larger movements of the examined variables allow better identification of the coefficients. However, it would be troubling if the results depended exclusively on few special events. The sample used for the EA5 countries, spanning 1987 to 1998, contains the volatile period of the ERM crisis in 1992, as well as the Mexican peso crisis of 1995, which had an impact on some of the western European countries. Eastern European countries were, for their part, strongly affected by the Russian crisis in 1998. Figure 12 presents the posterior medians of the estimated monetary policy shocks. It confirms that most of the largest monetary policy shocks occurred exactly during these volatile periods.

Such time-varying volatility of the shocks would call for more sophisticated methods, such as the VARs with time-varying volatilities of Uhlig (1997) or Markov-switching VARs of Sims and Zha (2006). Their application in conjunction with the exchangeable prior used here is a subject of further research. For the present purposes, we choose a simpler approach and remove the volatile periods from the sample, by truncating the samples and introducing period-specific dummy observations. In dynamic models, such dummying-out does not remove the information from these periods perfectly, because the delayed effect of the shocks in these periods persist for some time, but it certainly limits their effects on the results. In order to ensure that the effect of the volatile periods is strongly suppressed, we used sequences of consecutive dummy observations. Dummying-out is simpler, but of course it involves a greater information loss than efficient use of both volatile and calm periods.

The continuous vertical lines in Figure 12 show periods for which dummy observations were introduced, and the dashed vertical lines in the beginning of the sample show the points up to which the samples were truncated. We will refer to the estimates in those truncated samples with dummy variables included as *calm sample* estimates.

The Russian crisis erupted in August 1998, but was preceded by the Asian crisis, which also affected the Czech economy. Therefore, in the Czech Republic the sample was truncated from the beginning until July 1999. In Hungary and Poland, the half-year period starting in August 1998 was dummied out. Furthermore, in Hungary we remove the large 1996 monetary policy shock, that occurred when the central bank devalued the exchange rate to help implement the Bokros package of fiscal reforms. We remove the first half of 2003 (in January 2003 the forint came under speculative attack; see e.g. IMF Country Report No. 03/124) and the volatile period in 2006 when fiscal uncertainties resulted in pressures on financial markets (see e.g. IMF Country Report No.07/250). In Poland we additionally remove a four-month period in the middle of 2001 when political attacks on the central bank caused uncertainty in financial markets.

For Finland we dummy out the 26-month period starting from March 1991, during which the economy was in crisis and suffered a number of external shocks, from the collapse of the trade with the Soviet Union to speculative attacks (Honkapohja and Koskela, 1999). For France, Italy, Portugal and Spain, we dummy out the neighborhoods of the ERM crisis in the second half of 1992 and early 1993, as well as the Mexican peso crisis in the first half of 1995. In addition, we dummy out a volatile period in Portugal in late 1990/early 1991, and truncate the Spanish sample in the beginning to avoid the volatile first half of 1987.

Table 3 reports how successful we are in eliminating volatile periods. It reports normality tests for the median monetary policy shocks, both the baseline estimates and the calm sample estimates (the periods with dummy variables have zero shocks not included in the computation of the test statistics). Thanks to the inclusion of constant terms to which the exchangeable priors do not apply, shocks always have a mean close to zero. Lack of skewness cannot be rejected for all countries except for Hungary and France. Most importantly for the present discussion, lack of excess kurtosis is rejected at the 5% significance level in all countries except Slovenia. However, as indicated by the statistics in the last two columns, the shocks estimated on the calm samples exhibit no skewness and excess kurtosis disappears in all cases except Portugal.

Figure 13 shows the mean impulse responses estimated in the calm samples (with all the period dummies). Two observations stand out. First, the amplitude of the interest rate and the exchange rate movements conditional on the standard monetary policy shock is affected by the periods included. Without volatile periods, standard monetary policy shocks are associated in both NMS4 and EA5 with a median interest rate increase of about 40 basis points, instead of, respectively, 45 and 60 basis points in the baseline estimation. The median exchange rate appreciation is 1% and 0.6% respectively, compared with 1.2% in both countries in the baseline estimation. Second, impulse responses of output and prices are barely affected, and all observations stressed in the previous subsection (larger uncertainty around the NMS4 responses, stronger median NMS4 price response, and lower output cost of disinflation in NMS4 - see Figure 14) go through.

In another robustness check we tried refining the identification of monetary policy shocks by using the dynamics of the central bank foreign reserves. Sign restrictions discriminate between the monetary policy and other shocks only approximately, because we do not know the precise elasticity of the reaction of the exchange rate to exogenous interest rate movements. As a result,

Table 3: Normality tests for the estimated median monetary policy shocks

	full sample		calm sample	
	chi2(1)	p-value	chi2(1)	p-value
NMS				
Czech Republic				
mean	0.00	0.96	0.00	0.99
skewness	0.10	0.75	3.67	0.06
kurtosis	4.16	0.04	2.20	0.14
Hungary				
mean	0.00	0.97	0.00	0.99
skewness	26.19	0.00	0.99	0.32
kurtosis	98.67	0.00	1.66	0.20
Poland				
mean	0.00	0.98	0.00	1.00
skewness	0.00	0.99	2.15	0.14
kurtosis	4.21	0.04	0.20	0.65
Slovenia				
mean	0.00	1.00	0.00	0.99
skewness	0.72	0.40	1.06	0.30
kurtosis	1.14	0.29	1.21	0.27
EA				
Finland				
mean	0.00	0.98	0.00	0.98
skewness	0.09	0.77	0.03	0.85
kurtosis	20.28	0.00	1.06	0.30
France				
mean	0.00	1.00	0.00	0.98
skewness	40.70	0.00	0.94	0.33
kurtosis	319.29	0.00	2.98	0.08
Italy				
mean	0.00	0.97	0.00	0.98
skewness	0.34	0.56	0.20	0.66
kurtosis	78.72	0.00	0.00	0.99
Portugal				
mean	0.00	0.96	0.00	1.00
skewness	0.03	0.85	0.02	0.90
kurtosis	146.80	0.00	14.22	0.00
Spain				
mean	0.00	0.99	0.00	0.99
skewness	0.08	0.78	0.35	0.55
kurtosis	40.42	0.00	0.07	0.80

Note: The test statistics for the null hypotheses of no skewness and no excess kurtosis are computed as in the Jarque-Bera normality test. Denoting the first centered sample moment of the standardized residuals by (μ_1), third by (μ_3) and fourth by (μ_4), the test statistics are computed, respectively, as $T\mu_3^2/6$ and $T(\mu_4 - 3)^2/24$. The test statistic for the null hypothesis of zero mean is simply $T\mu_1^2$.

the identified shocks are in fact some linear combination of monetary policy shocks and other shocks, which involve a capital flow in the opposite direction. To sift off better these other shocks we add the following identifying assumption. We assume that the response of foreign reserves to an exogenous monetary tightening should be positive, consistently with the higher demand for the local currency it is supposed to generate.

We use this additional assumption in a specification of the VAR with foreign reserves as the fifth variable.[15] VARs with five variables are more heavily overparameterized, and, as shown in Figure 15, the responses of output and prices in both panels become weaker and less significant. The comparison between output costs of disinflation, shown in Figure 16, becomes less stark, but still suggests a steeper Phillips curve in the NMS4. We conclude that the qualitative results stressed in this paper are robust to refining the identification of monetary shocks by using the dynamics of foreign reserves.

A drop in foreign reserves conditionally on a monetary policy shock needs not be a signal of an identification failure. Sometimes central banks actively use foreign exchange market intervention as a monetary policy tool, and the level of exchange rate as one of the operating targets. Then an exogenous policy tightening could be implemented as an increase of the interest rate accompanied by a sale of foreign currency reserves in order to ensure a stronger exchange rate appreciation than otherwise. To see what happens on average in the analyzed samples, we introduce total foreign reserves as the fifth variable, without any constraint on their response to a monetary policy shock. Impulse responses of output and prices (unreported here) remain similar as in the baseline specification, but, as in Figure 15, they are also less significant, due to the more acute overparameterization of the model. The response of foreign reserves, displayed in Figure 17, is insignificant initially, but subsequently becomes positive. We conclude that the dynamics of foreign reserves is consistent with our interpretation of the baseline monetary policy shocks identified in this paper.

The analysis was repeated also with the recursive identification, assuming that monetary authorities react with a lag to the exchange rate developments. As shown in Figure 18, under this identification, the exchange rate appreciates with some delay, exhibiting 'delayed overshooting' leading to the 'forward discount bias puzzle' discussed in Kim and Roubini (2000). Output responses are initially insignificant in both panels, and a 'price puzzle' emerges in the NMS4, i.e. the price response is significantly positive initially. This confirms the findings of price puzzles under recursive identification for

[15]Logged, data from the International Financial Statistics labeled as Total reserves minus gold.

these countries, reported in a number of VAR studies discussed in Coricelli et al. (2006). After about two years, both output and price responses are negative, and the output cost of disinflation is again somewhat lower in the NMS4, although the difference is small and not statistically significant. The two main conclusions from this exercise are as follows. First, the results from the recursive identification point in the same direction as the baseline results of this paper. Second, we confirm the finding of Kim and Roubini (2000) that allowing for simultaneous reaction of interest rates and exchange rates to monetary policy shocks is needed for the proper identification of these shocks.

Finally, when the model is estimated with the baseline identification scheme but without any control variables (only constant terms and lags of endogenous variables), the output response becomes significantly stronger in the EA5 than in the NMS4, and, most importantly, the price response becomes positive in the euro area. We conclude that the exogenous controls are crucial for successfully identifying monetary policy shocks.

4 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 (the Czech Republic, Hungary, Poland and Slovenia) turn out to be qualitatively similar to those in the EA5 (Finland, France, Italy, Portugal and Spain), but with interesting differences (albeit estimated with significant uncertainty): while the output responses are broadly similar, the uncertainty bands for price responses include the possibility of stronger effects than in the EA5.

These results suggest that, when considering the differences between central-eastern and western Europe, we need to go beyond the rule of thumb that monetary policy is less effective in less financially developed countries. Another feature of these economies, which has been missing from the discussion so far, and which may be important for monetary transmission, is the higher inflation level and variance that they have experienced in the 1990s. Economic theory (Ball et al., 1988; Dotsey et al., 1999; Lucas, 1973) predicts that higher (or more volatile) inflation makes prices less sticky and the Phillips curve steeper. Our findings imply a steeper Phillips curve in the NMS4 on the full sample (compared with the EA5), consistently with these predictions. Confirming this interpretation, in the samples excluding the high-inflation 1990s we find that the estimated price responses tend to be weaker and output responses stronger than on the full sample, indicating a flatter Phillips curve.

Among the studied Central and Eastern European countries, Slovenia adopted the euro in 2007 and the remaining countries have committed to joining the EMU in the future. The question of similarity of monetary transmission between central-eastern and western European countries is therefore very relevant. However, some caution is needed in drawing the implications of this paper's findings for the adoption of the euro by the NMS. EMU entry will result in fixing the currently floating exchange rates with the main trading partners, and as a result the exchange rate channel will largely disappear. Also in other respects EMU entry may potentially be an important regime change, and therefore the Lucas critique may apply to the extrapolation of the results from this paper. Consequently, impulse responses reported in this paper should not be treated as forecasts of the responses to common monetary policy shocks after the euro adoption.

Instead, this paper compares monetary transmission in the NMS with that of the current EMU members *before* they adopted the euro, i.e. it provides a comparative assessment of the initial conditions. The caveat is, that for such initial conditions to be informative there should be no major differences in the transition to the EMU regime. The results of this paper do not support the structural weakness of monetary transmission in the NMS as an argument against further EMU expansion. However, the comparisons of this paper may be affected by different inflation levels and volatilities in the sample period.

Appendix A Data sources and estimation periods

Table A1. Database codes of variables.

Variable	code
Endogenous variables	
Industrial production	xxx66..CZF...†
Consumer price index	xxx64...ZF...
Money market interest rate	xxx60B..ZF...‡
Exchange rate: local currency per SDR	xxx..AA.ZF...
Exchange rate: German currency per SDR	134..AA.ZF...
Exogenous controls	
Exchange rate: German currency per USD	134..AE.ZF...
German industrial production	13466..CZF...
German interest rate	13460B..ZF...
Federal funds rate	11160B..ZF...
Petroleum:average crude price	POILAPSP
Non-fuel commodities index	PNFUELW
Robustness checks	
Total reserves minus gold	xxx.1L.DZF...

Notes: †Except Poland: 96466..BZF... ‡Except Hungary, where the money market interest rate was not available and the treasury bill rate, with code 94460C..ZF..., was used instead.

All data come from the IMF Internatonal Financial Statistics, (available at http://www.imf.org) except for commodity prices, which are taken from the IMF Primary Commodity Prices database available at
http://www.imf.org/external/np/res/commod/externaldata.csv
The data were downloaded on February 12th, 2008.

Country exchange rates w.r.t. the German currency (D-mark prior to 1999, euro afterwards) were obtained as the ratio of the rate local currency per SDR to the rate German currency per SDR.

Table A2. Baseline samples

Country	Begin	End	Obs
Czech Republic	1998 (1)	2007 (10)	118
Hungary	1995 (4)	2007 (10)	151
Poland	1995 (6)	2007 (11)	150
Slovenia	1997 (5)	2006 (12)	116
Euro area countries	1987 (1)	1998 (12)	144

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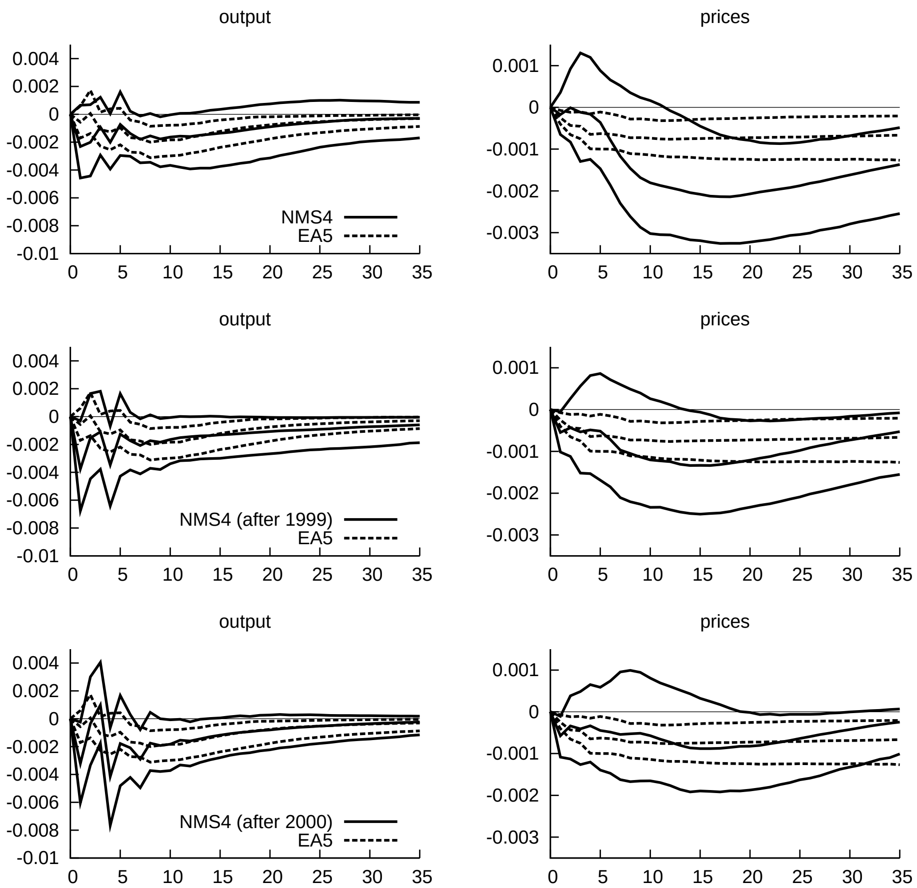


Figure 10: Mean impulse responses to a 50 basis points monetary policy shock for the NMS4 and the EA5: median, 5th and 95th percentiles of the posterior distribution. Full sample, and NMS4 samples starting in 1999 and 2000.

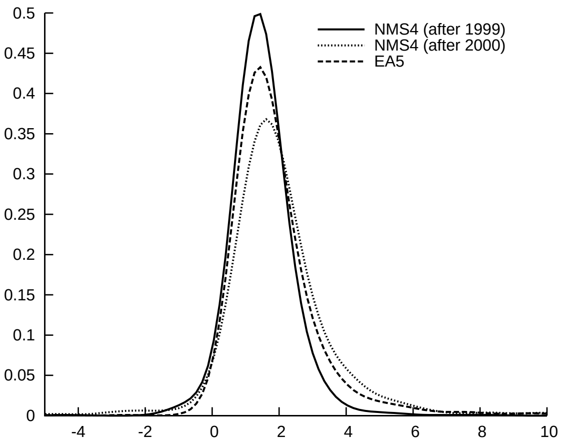
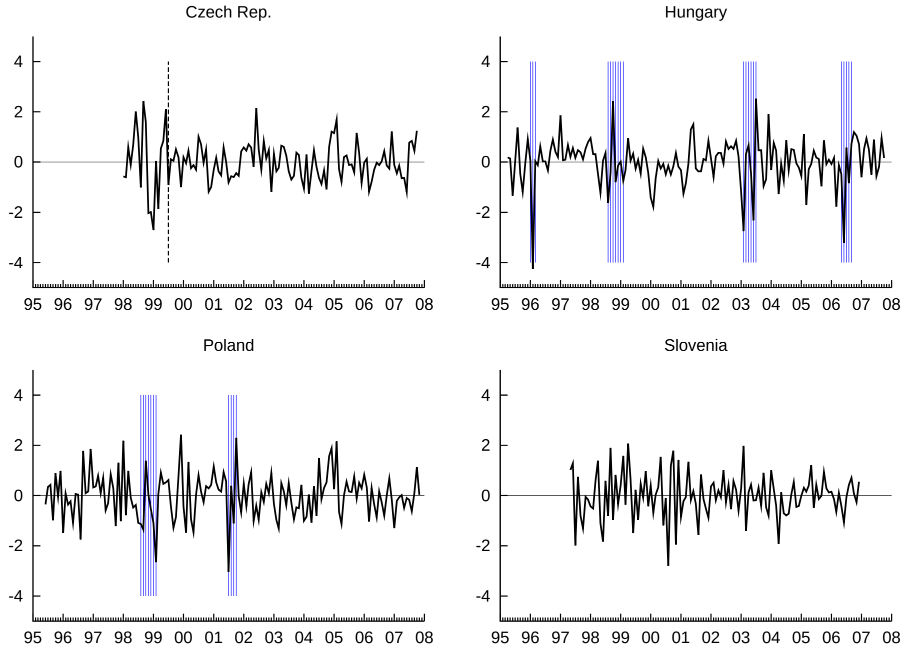
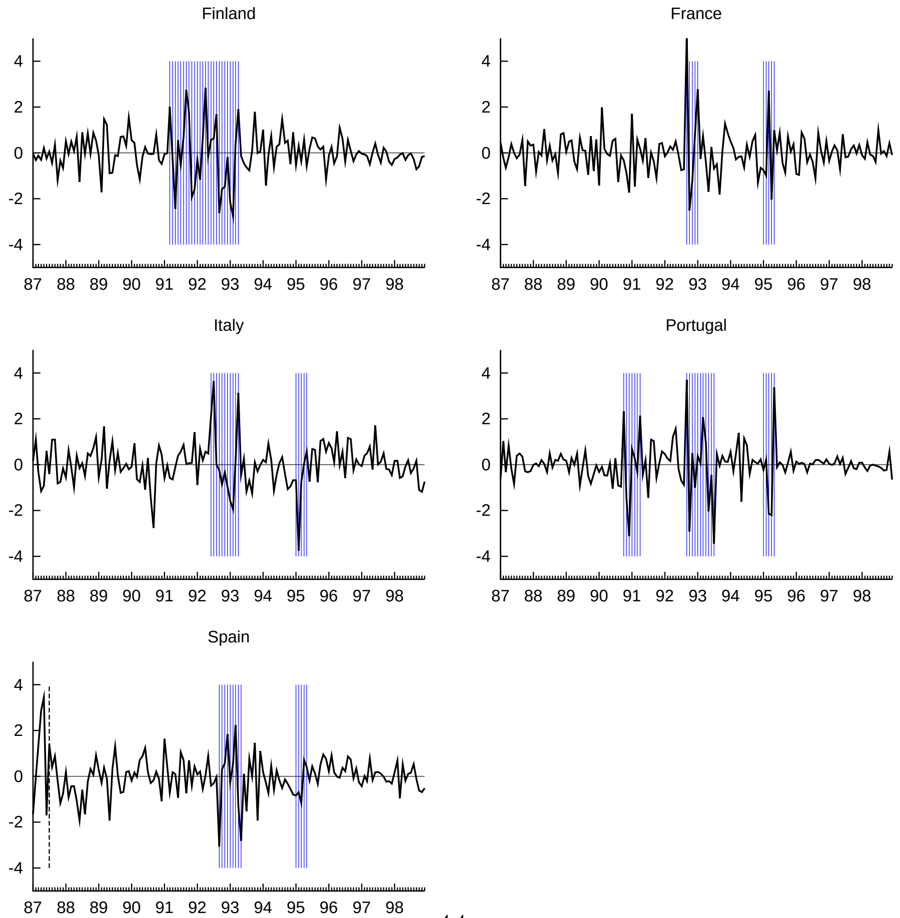


Figure 11: Posterior distribution of the output cost of disinflation (24-month horizon). Samples for NMS4 starting in 1999 and 2000.



A. NMS4



B. EA5

Figure 12: Median estimated monetary policy shocks and dummies for volatile periods.

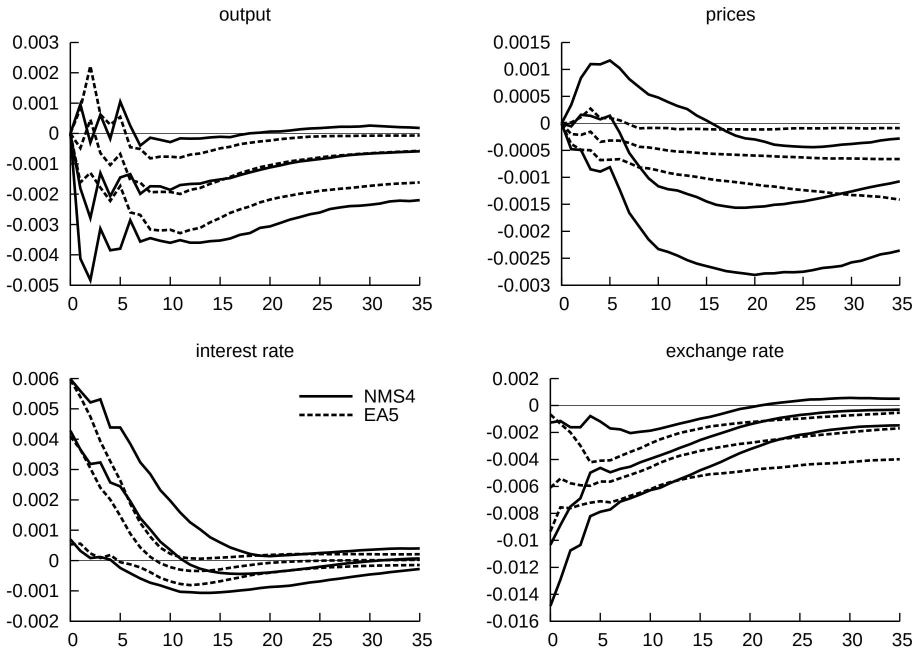


Figure 13: Mean impulse responses to monetary policy shocks for the NMS4 and EA5: median, 5th and 95th percentiles of the posterior distribution. Calm samples.

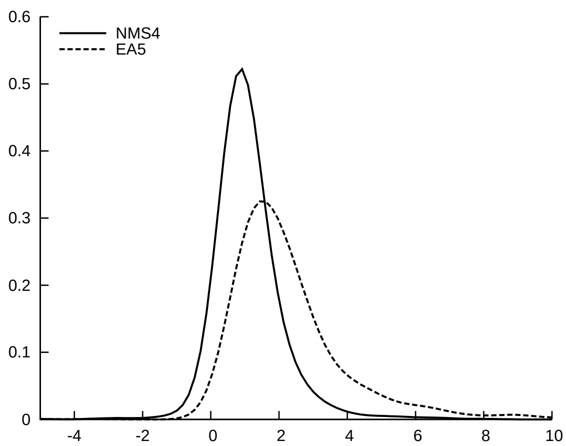


Figure 14: Posterior distribution of the output cost of disinflation (24-month horizon). Calm samples.

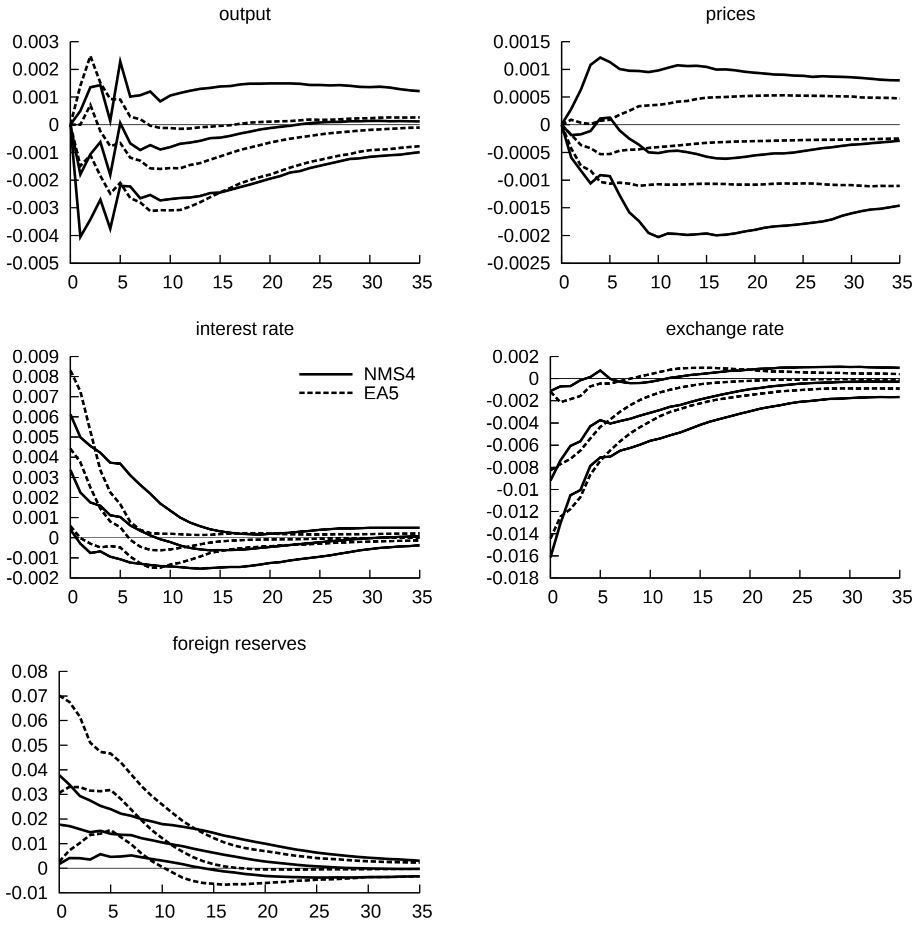


Figure 15: Mean impulse responses to monetary policy shocks for the NMS4 and EA5: median, 5th and 95th percentiles of the posterior distribution. Specification with foreign reserves.

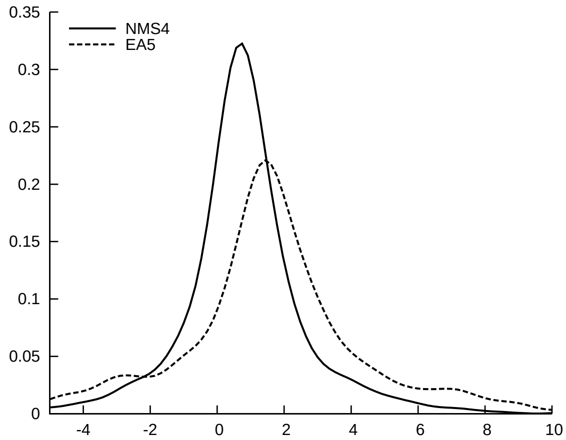


Figure 16: Posterior distribution of the output cost of disinflation (24-month horizon). Specification with foreign reserves.

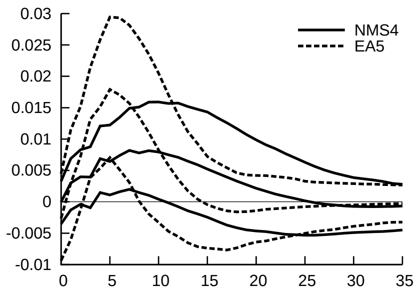


Figure 17: Mean impulse of foreign reserves to a monetary policy shock, when it is unconstrained in the identification.

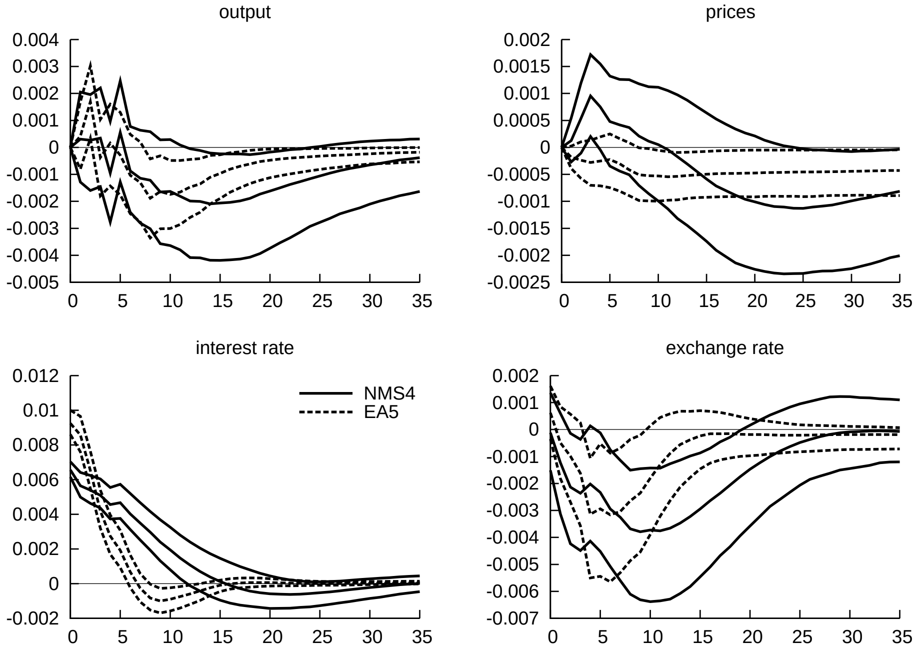


Figure 18: Mean impulse responses to monetary policy shocks for the NMS4 and EA5: median, 5th and 95th percentiles of the posterior distribution. Recursive identification.

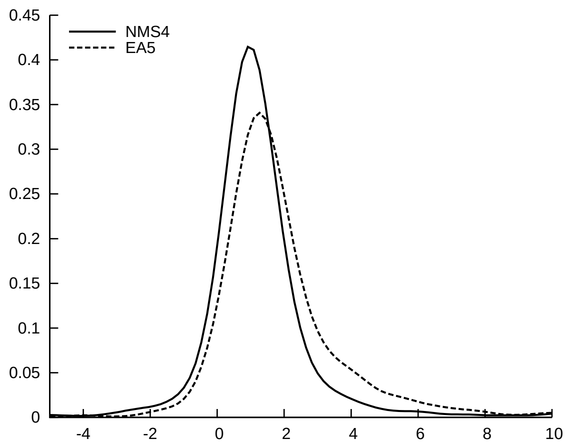


Figure 19: The posterior distribution of the output cost of disinflation (24-month horizon). Recursive identification.