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Sources of Real Exchange Rate Variability in Central and Eastern European Countries: Evidence from Structural Bayesian MSH-VAR Models

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Abstract

This paper investigates the relative importance of cost, demand, financial and monetary shocks in driving real exchange rates in four CEE countries over 2000–2018. A two-country New Keynesian open economy model is used as a theoretical framework. In the empirical part, a Bayesian SVAR model with Markov switching heteroscedasticity is employed. The structural shocks are identified on the basis of volatility changes and named with reference to the sign restrictions derived from the economic model. Main findings are fourfold. First, real and financial shocks have similar contributions to real exchange variability, whereas that of monetary shocks is small. Second, financial shocks amplify exchange rate fluctuations stemming from real shocks. Third, even though the exchange rate gaps change over time, they remain quite similar across CEE countries except for Slovakia. Fourth, Slovakia introduced the euro at the time of a relatively large real overvaluation, which subsided after a lengthy adjustment process.

Keywords: open economy macroeconomics, real exchange rate, real and nominal shocks, Bayesian MS-VAR models, structural VAR models

JEL Classification: F33, C11, F41, E44

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1 Introduction

Real exchange rate fluctuations translate into changes in external competitiveness of an economy, affect an internal allocation of resources between sectors and can undermine financial stability. The adverse effects of exchange rate variability can hinder trade relations by triggering off accusations of currency manipulations as evidenced by the trade conflict between China and the United States or by a rise in the use of protectionist measures in the aftermath of the global financial crisis (GFC) even in the European Union (see e.g., Evenett and Fritz, 2016; Wajda-Lichy, 2014). The real exchange rate variability is even more important in small open economies as they are more deeply integrated into global and/or regional supply chains, like Central and Eastern European (CEE) economies.

Our research objective is to examine the sources of real exchange rate variability in four CEE countries: the Czech Republic, Hungary, Poland and Slovakia. We intend to assess the relative importance of cost, demand, financial and monetary shocks in driving the real exchange rates over time and countries. Additionally, two research questions are raised: (1) Are sources of exchange rate variability similar across time, especially in the pre-GFC and crisis periods? (2) Are exchange rates in different CEE countries driven by similar shocks?

Real exchange rate volatility can be traced down to nominal exchange rate volatility as prominently observed by Mussa (1986) and more recently reminded by Rose (2011) and Krugman (2011). The natural explanation of this stylised fact is that nominal (monetary) shocks are the main driver of the real exchange rate. In fact, Juvenal (2011) calls this view the 'consensus' and backs it with Rogoff's words '[m]ost explanations of short-term exchange rate volatility point to financial factors such as changes in portfolio preferences, short-term asset price bubbles, and monetary shocks' (Rogoff, 1996). She observes, however, that the empirical evidence has not provided much support for this idea. In their seminal paper Clarida and Galí (1994) found that the real exchange rate variability was accounted for mostly by real and not nominal shocks. This observation has been confirmed in many other studies. Recently, Tian and Pentecost (2019) and Chen and Liu (2018) found demand shocks were the most important source of real exchange rate fluctuations in China. Gehrke and Yao (2017) demonstrated that both supply and demand shocks played an important role in driving the U.S. real effective exchange rate fluctuations.

Empirical evidence does not necessarily invalidate the 'consensus'. We think that the relative unimportance of nominal shocks can result from (i) neglecting financial shocks in an empirical specification of the underlying model, and (ii) shifts in the shocks' volatility. There are at least two good reasons to include financial shocks in the analysis. First, the risk that the identified shocks do not correspond to the underlying shocks is mitigated (see e.g., Dąbrowski (2012). This point is also related to the problem of aggregation of multiple shocks. See, e.g., Faust and Leeper (1997))

Second, financial shocks are found to be important in other studies. For example, Gehrke and Yao (2017) demonstrated that the risk premium shocks explain 36%–58% of real exchange fluctuations in the long run and Eichenbaum et al. (2020) found that shocks to the foreign demand for dollar-denominated bonds drive the bulk of exchange rate movements. (See also, Farrant and Peersman (2006) and Peersman (2011) who provide evidence on an important role for risk premium shocks in explaining exchange rate fluctuations in the short run in advanced economies.)

Shifts in the shocks' volatility can further contribute to the gap between the 'consensus' view and empirical evidence. It is because nominal shocks stem from developments in financial markets and their volatility is more likely to be subject to large and abrupt changes than that of real shocks. The analytical framework should, therefore, be flexible enough to capture switches between high and low volatility regimes. For example, Herwartz and Lütkepohl (2014) using VAR models with Markov-switching heteroscedasticity argue that monetary shocks can have quite different characteristics across volatility states, including the forecast error variance components due to these shocks.

Existing evidence on the sources of real exchange rate variability in CEE countries is not unambiguous. In some studies, nominal shocks are found to be the most important driver of exchange rate fluctuations. For example, Borghijs and Kuijs (2004) argue that exchange rates propagate monetary and financial shocks rather than absorb real shocks. More recently, a similar point is made by Shevchuk (2014). In other studies, however, real shocks are found to be the primary source of exchange rate variability. For example, Dąbrowski and Wróblewska (2016) demonstrated that real shocks are at least as important as nominal shocks in Poland and Slovakia. For a similar finding, see, e.g. Arratibel and Michaelis (2014), Stążka-Gawrysiak (2009). Audzei and Brázdik (2015) are in the middle ground as they find that nominal shocks explain slightly more than half of the variability of the real exchange rate in the Czech Republic. More recently, these authors examined a larger set of CEE economies and found that the contribution of financial shocks is about 10% (Audzei and Brázdik, 2018; the contribution of monetary shocks is not reported).

Our contribution to the literature is threefold. First, our approach has firm theoretical foundations. We use a New Keynesian small open economy model borrowed from Engel and West (2006) and Galí and Monacelli (2005) as a theoretical framework, but extend it to allow for financial shocks. Such an extension proves crucial: financial shocks are found to be important, and in some cases even dominant, source of exchange rate variability in CEE countries. Their omission would lead to misleading results on shocks identification in general and the importance of nominal shocks in particular.

Second, to methodologically facilitate the identification of shocks we allow for Markovian breaks in (conditional) covariance matrix of otherwise homoscedastic VAR models, which from an empirical perspective is intended to capture possible shifts in the shocks' volatility. Such an extension proves not only valid in our analysis (although not for every country under study), but also enables us to use this additional statistical information on volatility regime changes to identify the underlying structural shocks. Although the very idea of identification through heteroscedasticity (in simultaneous-equation models) dates back to Rigobon (2003), the concept has been transferred to and developed within the context of SVAR models by Lanne et al. (2010); see also Kilian and Lütkepohl (2017) and references therein. Kulikov and Netšunajev (2013), Woźniak and Droumaguet (2015) and Lütkepohl and Woźniak (2020) developed Bayesian framework for the MSH-VAR models, with the first two papers retrieving the structural shocks from the reduced form (the approach also adopted in our research), while the latter two directly estimating the structural model. Whichever of either Bayesian or 'classical' maximum likelihood statistical setting one chooses, still, as indicated by Herwartz and Lütkepohl (2014), even in the apparent presence of heteroscedasticity it may yet be the case that not all shocks can be successfully isolated. Then, introducing some additional, conventional constraints such as zero or sign restrictions (typically imposed and indispensable in homoscedastic SVARs) may prove useful. Since our results from 'pure' (i.e. without additional restrictions) SVAR models with Markov switching heteroscedasticity allow for only partial identification of shocks, we too resort to such a combined approach in this work.

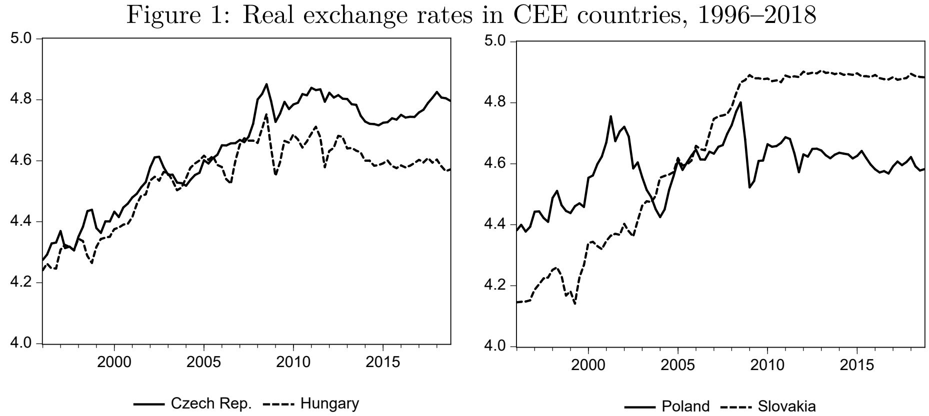
Third, we provide evidence on the importance of real and nominal shocks in driving the real exchange rate variability in CEE countries in 2000–2018 period using detailed results of forecast error variance and historical decompositions. The results are summarised in four findings. First, real and financial shocks have important and similar contributions to real exchange variability, whereas that of monetary shocks is negligible. Second, financial shocks amplify the real exchange rate fluctuations caused by real shocks. Third, the exchange rate deviations from the long-run level change over time, but remain quite similar across CEE countries except for Slovakia. Fourth, the euro was introduced in Slovakia at the time of a relatively large real overvaluation, which subsided after a rather lengthy adjustment process. Using results for CEE countries we also derive two policy implications on the desirability of exchange rate management.

The rest of the paper is structured as follows. In the next section, a stylised view of exchange rate fluctuations in CEE countries is sketched. Section 3 lays out a theoretical model and discusses the reactions of key macroeconomic variables to structural shocks. In Section 4 both the methodology and data used in the empirical analysis are presented. Empirical results are discussed in Section 5. The final section delivers conclusions.

2 A stylised view of exchange rate fluctuations in CEE countries

It is worthwhile to start with the brief inspection of real exchange rate developments in CEE countries. In Figure 1 the logs of real exchange rates in 1996–2018 are depicted. Three observations can be made. First, all CEE currencies appreciated substantially: the average annual rate of real appreciation ranged from 0.9% for the Polish zloty to 3.3% for the Slovak koruna (data sources are described in Section 4.2). This phenomenon is usually explained in the literature with the Balassa-Samuelson effect (see e.g., Egert et al., 2003; Sonora and Tica, 2014). Being middle-income countries at the beginning of the economic transition that started in the early 1990s they recorded relatively fast growth in the following years and were moved by the World Bank to the high-income group in the second half of 2000s.

Figure 1: Real exchange rates in CEE countries, 1996–2018



Second, the appreciation trend was permanently interrupted by the Global Financial Crisis (GFC) in 2008–2009. The whole appreciation can, in fact, be attributed to changes in the pre-crisis period. Indeed, the average annual rates were higher by 1–2 p.p. in 1996q1–2007q2 than in the whole period. After wide gyrations in years 2007–2010, the average annual rates decreased close to zero in the Czech Republic and Slovakia and even shifted to negative territory in Hungary and Poland.

Third, during the GFC the CEE exchange rates fluctuated substantially. In the run-up to the crisis (2007q3–2008q3) CEE currencies appreciated heavily: the average annual rates were at least twice as large as in the pre-crisis period and ranged from about 8% in Hungary to more than 20% in the Czech Republic. These changes were more than compensated by depreciation during the most severe financial turbulence (2008q4–2009q1): the annualised rates of change ranged from −41% to −73%. The only exception was Slovakia with a 10%-appreciation. This overshooting was eliminated in the following five quarters: currencies appreciated from 6% to 14% (again on annual basis) and in Slovakia a small real depreciation of 2% was observed. Natural questions implied by these empirical regularities concern the importance of shocks in driving the exchange rates, especially in the run-up to, during and after the GFC. The research questions are, therefore, as follows: (1) Are sources of exchange rate variability similar across time, in particular in the pre-GFC and crisis periods? (2) Are exchange rates in CEE countries driven by similar shocks?

3 Theoretical framework

A New Keynesian small open economy model developed by Engel and West (2006) is used as a theoretical framework. The key model assumptions fit reasonably well characteristics of CEE economies. These economies are open to trade and financial flows and at the same time both their exports and prices have a negligible contribution to foreign consumption and price level.

The model itself is similar to the framework developed by Galí and Monacelli (2005). (For a well-organised and detailed exposition of the model see also chapter 8 in Galí (2015).) It is a fully microfounded two-country and two-good model of an open economy that consists of four main equations: interest rate rule, Phillips curve, IS relation and uncovered interest rate parity condition. In Engel and West's version, the model includes three shocks: cost, demand and monetary. We extend their approach and include financial shocks as well.

As is conventional in the literature the central bank is assumed to stabilise both prices and output. It is also assumed that the domestic central bank cares about the exchange rate stability, so the exchange rate term enters the interest rate rule:

$$i_t = -\gamma_q q_t + \gamma_\pi E_t \pi_{t+1} + \gamma_y y_t + u_{mt}, \tag{1}$$

where q_t is the real exchange rate (its increase is an appreciation of home currency). Constant and trend terms are omitted in this and other equations. Therefore, the long-term level of the real exchange rate is zero and non-zero q_t's are deviations from the target level. All other variables are defined as differences between domestic and foreign variables (denoted with a superscript h and asterisk, respectively). Thus, i_t is the interest rate differential $i_t^h - i_t^*$, π_t stands for the difference in inflation rates $\pi_t^h - \pi_t^*$ and y_t is the difference between output gaps $y_t^h - y_t^*$. (Following Engel and West (2006) we use an expected inflation along with a current output gap to simplify calculations.) The monetary disturbance u_t^m is assumed to be a stationary AR(1) process driven by monetary shocks. All parameters are positive and γ_π is greater than unity (the Taylor principle).

Prices are sticky and their adjustment is described by the Phillips curve:

$$\mathring{\pi}_t = \beta E_t \mathring{\pi}_{t+1} + \kappa y_t + u_{ct}, \tag{2}$$

where $\mathring{\pi}_t$ is a difference between inflation rates of domestically produced and foreign goods, $\pi_{dt} - \pi_t^*$. Both these inflation rates enter the home country inflation as both goods are bought at a home country: $\pi_t^h = (1-\alpha)\pi_{dt} + \alpha\pi_{ft}$. The parameter α is the weight of imported goods. It can be shown that $\pi_t = \mathring{\pi}_t - \alpha\Delta q_t/(1-\alpha)$. The cost disturbance u_t^c has analogous properties to the monetary disturbance. The parameters β and κ are positive.

The output gap differential can be described as related to the real exchange rate and demand and financial disturbances:

$$y_t = -\theta_q q_t + u_{dt} - \theta_f u_{ft}, \tag{3}$$

where both disturbances are stationary AR(1) processes driven by demand and financial shocks, respectively. Parameters θ's are positive. Equation (3) can be obtained from the dynamic IS relation in an open economy and the UIP condition.

The uncovered interest rate parity condition includes the risk premium term which is interpreted as a financial disturbance:

$$i_t = -(E_t s_{t+1} - s_t) + u_{ft}, \tag{4}$$

where s_t is a nominal exchange rate (its increase is a nominal appreciation of a home currency).

After combining these four equations and the relation between π_t and $\mathring{\pi}_t$ one can write down the system of difference equations in the output gap and inflation differential:

$$\begin{bmatrix} 1 & \kappa\theta \\ 0 & \eta+\gamma \end{bmatrix} \begin{bmatrix} \mathring{\pi}_t \\ q_t \end{bmatrix} = \begin{bmatrix} \beta & 0 \\ \gamma_\pi - 1 & \eta \end{bmatrix} \begin{bmatrix} E_t \mathring{\pi}_{t+1} \\ E_t q_{t+1} \end{bmatrix} + \begin{bmatrix} 1 & \kappa & -\kappa\theta_f & 0 \\ 0 & \gamma_y & -(1+\gamma_y\theta_f) & 1 \end{bmatrix} u_t', \tag{5}$$

where $u_t = \begin{bmatrix} u_{ct} & u_{dt} & u_{ft} & u_{mt} \end{bmatrix}'$. This system can be solved recursively, i.e. first for inflation differential and then for the real exchange rate. (Local uniqueness and stability of the solution require imposing restrictions on parameters in the first two matrices in equation (5). It is assumed that they hold.) The solution for other variables can be constructed. It has the following form:

$$\begin{bmatrix} y_t & q_t & \mathring{\pi}_t & i_t \end{bmatrix}' = \Psi u_t', \tag{6}$$

where Ψ is a matrix of functions of parameters of the model.

Using the signs of parameters in the matrix Ψ it is possible to determine the direction of variable reaction to a given shock. The signs of responses are reported in Table 1. The interpretation of responses is in line with economic intuition. A positive cost

Table 1: Reactions of endogenous variables to structural shocks

Variable	Symbol	Shocks			
		cost	demand	financial†	monetary
Relative output	y_t	–	+	[+]	–
Real exchange rate	q_t	+	+	–	+
Relative inflation‡	$\mathring{\pi}_t$	+	[+]	[+]	–
Interest rate diff.	i_t	+	+	+	+

Notes: † See comments in the main text. ‡ Reactions of π_t are the same as $\mathring{\pi}_t$ except for the reaction to a demand shock which is ambiguous (the direct effect is positive but the impact through the real exchange rate is negative). In brackets restrictions *not* imposed in empirical part.

shock increases relative inflation and makes the central bank to raise the interest rate. Under the assumption that the Taylor principle holds the real appreciation will follow and output response will be negative.

A positive demand shock raises output and sets off equilibrating forces that drive inflation of domestic goods relative to foreign goods up and result in a real appreciation. These two effects make it difficult to say how the differential between home and foreign inflation, π_t will change. The interest rate will unambiguously increase.

Both inflation and output are adversely affected by a positive monetary shock and the domestic currency appreciates.

The responses to a financial shock are less straightforward as they depend, at least to a certain extent, on the weight of the exchange rate stability in central bank policy. In principle, a positive financial shock results in deprecation of domestic currency (the UIP condition). This translates positively into output and inflation and makes the central bank to increase the rate of interest. If, however, the exchange rate is heavily stabilised then the interest rate is raised to a level that prevents nominal depreciation of domestic currency. Thus, output and inflation are adversely affected. The latter reaction results in real deprecation of the domestic currency.

The restrictions on short-term responses derived from the small open economy model are imposed in the empirical part in order to identify structural shocks. Two points should, however, be made. First, following the discussion of responses to a financial shock, restrictions on output and inflation reactions are not imposed. Second, taking into account that the empirical counterpart of relative inflation is closer to a simple inflation differential π_t rather than a difference between inflation rates of domestically produced and foreign goods $\mathring{\pi}_t$, the restriction on a response to demand shock is not imposed.

4 Methodology and data

4.1 Statistical model and inference

We start our analysis with a reduced-form Bayesian n-variate VAR(2) model with Markov-switching heteroscedasticity (MSH-VAR(2)) built for the considered four-dimensional time series (y_t):

$$y_t = A_1 y_{t-1} + A_2 y_{t-2} + \mu + u_t, \quad u_t|S_t \sim iiN(0, \Sigma_{S_t}), \quad t = 1, 2, \ldots, T, \tag{7}$$

where $\{S_t,\ t \in \mathbb{Z}\}$, $S_t \in \{1,2\}$, is a two-state homogeneous and ergodic Markov chain with transition probabilities denoted by p_{ij}, $\forall_{t\in\mathbb{Z}}\ P(S_t = j|S_{t-1} = i) = p_{ij}$, $p_{i1} + p_{i2} = 1$, $p_{ij} \in (0,1)$, $i, j \in \{1,2\}$. Matrices Σ_{S_t} are symmetric and positive-definite. The initial conditions (y_{-1}, y_0) are treated as known and set as pre-sample observations. Matrices A_1 and A_2 collect autoregressive parameters, whereas μ represents the vector of constants.
We impose the following prior distribution for the model's parameters:

The inverted Wishart distributions for the covariance matrices:
$\Sigma_1 \sim iW(2S_\Sigma, n+2)$, $\Sigma_2 \sim iW(S_\Sigma, n+2)$, where the hyperparameter S_Σ is obtained from the training sample covering the period previous to the analysed one. Since the available training samples used here for various countries under study are rather short, we intend not to lose any data points from them. Therefore, we resort to the HP filter to remove their long-run trend. However, for the modelled sample data we employ the method introduced by Hamilton (2018), due to its many advantages (over the HP filter, in particular), discussed in Subsection 4.2. Despite this apparent inconsistency, such an approach largely facilitates our setting reasonable hyperparameters for the priors.
The difference between the priors assumed for Σ_1 and Σ_2 reflects our prior belief according to which the second-state (conditional) variance is lower than the one in the first state.

The matrix normal prior for the matrix A collecting the parameters of the VAR part of the estimated model, i.e. $A' = \begin{pmatrix} A_1 & A_2 & \mu \end{pmatrix}$: $A \sim mN(0, I_4, \underline{\Omega})$, where $\underline{\Omega} = diag(0.2I_4,\ 0.1I_4,\ 0.2)$, so $\underline{\Omega}$ is a diagonal matrix.

The uniform distribution for the transition probabilities:
$p_{ii} \sim U(0,1) = Beta(1,1)$, $i = 1, 2$.

The joint prior distribution is truncated by the stability condition:

$$p(\theta) = p(\Sigma_1, \Sigma_2, A, p_{11}, p_{22}) = p(\Sigma_1)p(\Sigma_2)p(A)p(p_{11})p(p_{22})I_{[0,1)}(|\lambda|_{max}),$$

where θ collects the model's parameters, $\theta = (\Sigma_1, \Sigma_2, A, p_{11}, p_{22})$, $I_{[a,b)}(.)$ is the indicator function of the interval $[a,b)$ and λ stands for the vector of the eigenvalues of the companion matrix for the analysed VAR(2) process (see, e.g., Lütkepohl, 2005):

$$\begin{pmatrix} A_1 & A_2 \\ I_4 & 0_{4\times 4} \end{pmatrix}.$$

Note that our model is similar to the one used by Dąbrowski et al. (2018), but in the present work we do not consider the possibility of (co)integration of the analysed time series. We do not observe any difficulty with fulfilling the stability condition by the data at hand, so it appears that a stationary model describes the dynamics of the considered time series well enough.

The states are identified via an inequality restriction in which we assume that the conditional volatility of the exchange rate in the first state is higher than in the second one.

The reduced-form errors (u_t) are linear combinations of the structural shocks (denoted by ε_t), i.e. $u_t = B\varepsilon_t$. As shown by Lanne et al. (2010), it is possible to utilise switches in covariance matrix (between Σ_1 and Σ_2) of the reduced-form shocks to identify the structural shocks. In the two-state case there always exist matrices B and Λ_2 such that

$$\Sigma_1 = BB',\ \Sigma_2 = B\Lambda_2 B',\ \Lambda_2 = diag(\lambda_{21},\ \lambda_{22},\ \lambda_{23},\ \lambda_{24}), \tag{8}$$

so it is possible to simultaneously diagonalise the covariance matrices. As proved by Lanne et al. (2010), this diagonalisation is locally unique (i.e. up to changing the signs and ordering of B's columns) if the diagonal elements of Λ_2 are all distinct, and so are the structural shocks defined as $\varepsilon_t = B^{-1}u_t$. Therefore, to ensure the global identification we additionally need to impose the shocks' order and to normalise them by setting the sign of one reaction to each shock (i.e. one element in each column of B). In this paper, we put λ_{2i}'s in descending order, whereas the normalisation restrictions are based on the presented economic model underlying this study (see Table 1). These restrictions are the same across the data sets analysed in this paper, and will be discussed later.

It is worth emphasising that structural shocks obtained in this framework feature switching (therefore, time-varying) covariance matrices (I_4 in the first state and Λ_2 in the second one), but the trajectories of the variables' responses remain unchanged. Moreover, as the matrix Λ_2 is diagonal, the structural shocks are orthogonal also in the second state, with $\lambda_{2i},\ i = 1, 2, 3, 4$, being their individual variances. Note also that, since $Var(\varepsilon_{ti}|S_t = 1, \theta) = 1$ and $Var(\varepsilon_{ti}|S_t = 2, \theta) = \lambda_{2i}$, each λ_{2i} may be regarded as a relative change of ε_{ti}'s variance in the second state with respect to its (unit) value in the first regime (Lanne et al., 2010, Dąbrowski et al., 2018).

As indicated by Herwartz and Lütkepohl (2014), even in the apparent presence of heteroscedasticity it may still be the case that not all shocks can be successfully isolated. Then, introducing some additional, conventional constraints such as zero or sign restrictions (typically imposed and indispensable in homoscedastic SVARs) may prove useful. Sign restrictions are implemented by truncating the domains of the models' parameters, which requires an additional acceptance-rejection step

within the Gibbs sampler. Let us note that facilitating statistical identification (via Markovian heteroscedasticity) with ancillary sign restrictions also help us to avoid well-known problems occurring in models identified solely with the sign constraints (see Baumeister and Hamilton, 2015, Baumeister and Hamilton, 2018).
Admittedly, the above-presented identification scheme can be employed only for such data sets which do support Markovian breaks in volatility. Since this is not the case for all CEE countries under study, we also employ the conventional sign restrictions approach (see, e.g., Kilian and Lütkepohl, 2017, Arias et al., 2018) within a homoscedastic VAR(2) framework:

$$y_t = A_1 y_{t-1} + A_2 y_{t-2} + \mu + u_t, \quad u_t \sim iiN(0, \Sigma), \quad t = 1, 2, \ldots, T. \tag{9}$$

The prior for the matrix A remains the same as in the MSH-VAR(2) model considered above: $A' = \begin{pmatrix} A_1 & A_2 & \mu \end{pmatrix}$: $A \sim mN(0, I_4, \underline{\Omega})$, with $\underline{\Omega} = diag(0.2I_4,\ 0.1I_4,\ 0.2)$. For the covariance matrix Σ we impose the same prior as for Σ_1, i.e. $\Sigma \sim iW(2S_\Sigma, n+2)$, with S_Σ obtained via a training sample, similarly as in model (7).
The sign restrictions are depicted in Table 1. Note that the sign patterns of the shocks differ from each other, so the model is fully identified. However, it should be emphasised that by imposing only sign restrictions the considered model is not point, but set identified. For this reason, we are able only to bound the parameters of interest, so we employ the method proposed by Inoue and Kilian (2013) (see also Kilian and Lütkepohl, 2017, pp. 445–447, Inoue and Kilian, 2019) to summarise the obtained results. The point estimate is obtained within the modal (most likely) model and the uncertainty is presented by displaying a set of trajectories obtained in 68% highest posterior density (HPD) credible sets of models.
We close this subsection with a short reflection on the very specification of the volatility process. Obviously, heteroscedasticity can be modelled with a wide range of approaches (see Lütkepohl and Netšunajev, 2017, for a recent review). Arguably, for some economies a two-state Markov-chain may emerge too limited for capturing changes in volatility and correlation over time, and more flexible processes should be considered instead, such as multivariate GARCH (MGARCH), stochastic volatility (MSV) or hybrid MSV-MGARCH structures (see Osiewalski, 2009; Osiewalski and Pajor, 2009). In a different context of modelling long-term relationships (within VEC framework) and probabilistic forecasting of macroeconomic time series, such extensions (and the hybrid structures, in particular) proved far superior to homoscedastic VAR/VEC models (see Pajor and Wróblewska, 2017; Wróblewska and Pajor, 2019). Moreover, results presented by Kwiatkowski (2020) indicate that although introducing Markovian heteroscedasticity into SVAR/SVEC model greatly improves the out-of-sample performance, it still gives way to an outrunning VEC-MSV-MGARCH specification. Utilizing MGARCH or MSV processes in structural shocks identification has already been considered in literature (see, e.g., Bertsche and Braun, 2020, Lewis, 2019, Lütkepohl and Netšunajev, 2017 and references therein). Extending our current methodology to the VAR-MSV-MGARCH

(instead of MSH-VAR) models appears a valid and promising avenue for further research.

4.2 Data

We use data for four Central and Eastern European countries: the Czech Republic, Hungary, Poland and Slovakia, whereas the euro area (19 countries) represents a foreign economy. In order to construct four key variables implied by the theoretical framework adopted, we use quarterly data for the real GDP and three-month money market interest rate and monthly data on the average euro (ECU) exchange rate (the price of domestic currency in euros) and a harmonised index of consumer prices (HICP). Monthly data are used to obtain quarterly averages. All data span, in principle, 1995q1–2018q4 period. The details and data sources are depicted in Table 6 in Appendix A.

Four variables used in the empirical part include the relative output gap, the deviation of the real exchange rate from its normal level (the exchange rate gap), the relative inflation gap and the interest rate gap differential. These key variables are obtained in two steps. First, gaps are constructed for each of the five countries as cyclical components of the relevant time series. We employ the method put forward by Hamilton (2018). Thus, taking into account that our data are quarterly gaps are obtained as residuals from a set of OLS regressions:

$$y_{t+8} = \beta_0 + \beta_1 y_t + \beta_2 y_{t-1} + \beta_3 y_{t-2} + \beta_4 y_{t-3} + \nu_{t+8}, \tag{10}$$

where y_t is the level of relevant variable and residual $\hat{\nu}$ is the transient component (gap). The method allows on the use of not seasonally data as any cycles with a frequency of one year are wiped out. This is important since we do not need to deseasonalize data risking the disturbance of their properties.

Second, the difference between the country-specific gap and the gap for the euro area for a given variable is calculated. This step is, of course, not needed for the real exchange deviation as the exchange rate is by definition a relative variable. For details see Table 6 in Appendix A.

Descriptive statistics of variables used in the empirical part are reported in Table 7 in Appendix A.

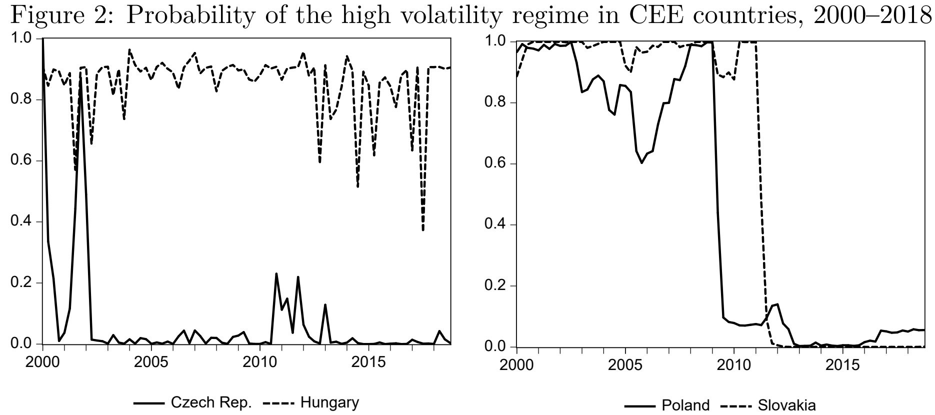
5 Empirical results

5.1 Models estimation results

The VAR model that allows Markovian breaks in the conditional covariance matrix is estimated for each CEE country. The advantage of such a model is, as explained in Section 4, that it does not require variances of shocks to be constant over time. Moreover, it enables us to estimate the timing of the volatility changes: each state is

assigned a posterior probability at each time t.
The posterior probabilities of the first state, i.e. the one with a higher (conditional) variance of shocks, are illustrated in Figure 2. Two observations can be made. First, there is little evidence that lends support to Markov switching in error covariances in the Czech Republic and Hungary. In the former country, the first state probability is greater than 0.5 in only two quarters in the early 2000s and stays very close to zero in almost all remaining periods. In the latter country, the probability fluctuates between 0.7 and 1 for more than 90% of the time and is lower than 0.5 in just one quarter. Second, switching between regimes is more important in Poland and Slovakia. In both countries, the switch from the high volatility regime to the low volatility regime can be observed in 2009 and 2011, respectively. Following these observations, we decided to retain to the MSH-VAR models for Poland and Slovakia and use time-invariant error covariance models for the Czech Republic and Hungary.

Figure 2: Probability of the high volatility regime in CEE countries, 2000–2018



As mentioned in Section 4, the structural shocks within the MSH-VAR framework (model 7) can be statistically identified if their variances in the second state are all distinct. Therefore, in Table 2 we inspect the posterior quantiles for λ_{2i}'s, obtained in the MSH-VAR models estimated for two out of four countries under consideration (i.e. for Poland and Slovakia, for the economies of which the VAR models with Markovian breaks in conditional covariance matrices emerged empirically relevant).
It can be noticed that the 90%-credible intervals of the second-state structural shocks' variances overlap in both cases: for two pairs $\lambda_{22},\ \lambda_{23}$ and $\lambda_{23},\ \lambda_{24}$ for Poland, and for one pair $\lambda_{21},\ \lambda_{22}$ for Slovakia. Therefore, the conditions for the local identification appear not fulfilled. On the other hand, the models are partially identified: in the model for Poland the first shock is fully identified, whereas in the model for Slovakia – the third and the fourth one.
It should be also emphasised that the approach employed in this work enables a statistical identification (and thereby, distinction) of the shocks, but they still may

Table 2: Posterior quantiles of the structural shocks' variances in the second state

Poland					
probability	0.05	0.16	0.5	0.84	0.95
λ_{21}	0.798	1.081	1.718	2.761	3.797
λ_{22}	0.203	0.251	0.363	0.565	0.776
λ_{23}	0.091	0.112	0.152	0.205	0.249
λ_{24}	0.036	0.046	0.064	0.089	0.108
Slovakia					
probability	0.05	0.16	0.5	0.84	0.95
λ_{21}	0.213	0.257	0.353	0.495	0.629
λ_{22}	0.111	0.133	0.178	0.238	0.287
λ_{23}	0.034	0.042	0.058	0.080	0.099
λ_{24}	0.011	0.013	0.019	0.026	0.032

not have any economic meaning. Therefore, before introducing additional restrictions, we first inspect the results obtained in these partially identified models (such as IRFs, FEVDs and historical decompositions) and attempt to name the shocks, what should permit us to establish reasonable additional sign restrictions (to facilitate and complete the identification), posited by the economic model, for the elements of the matrix B (i.e. instantaneous effects of the shocks upon the endogenous variables).

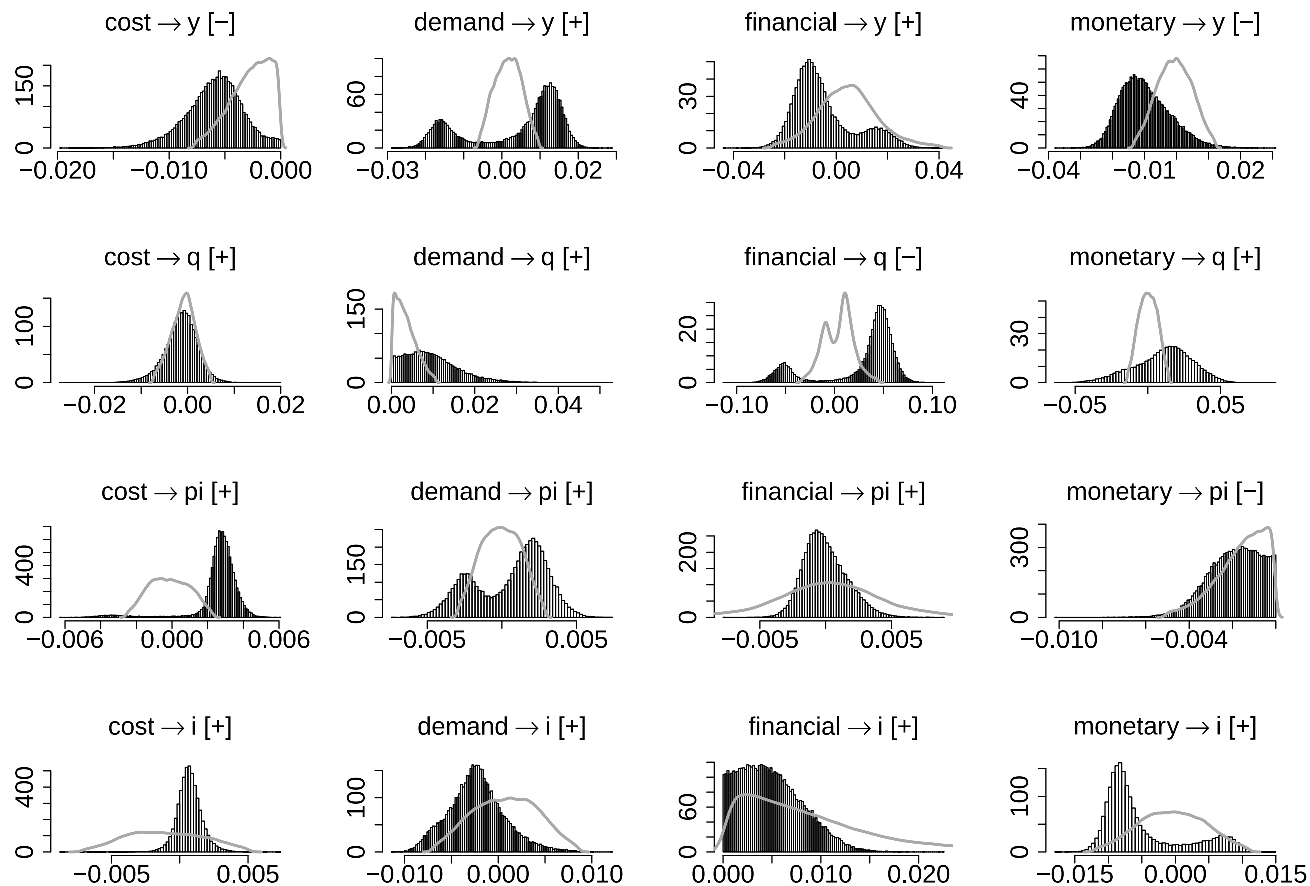
The posterior distributions of some elements of matrix B (in the models without additional sign restrictions) feature pronounced bimodality, which corroborates only partial identification of shocks (see Figures 3 and 4). Most importantly, however, in most of these bimodal cases, the prominent mode is supported by the economic model.

In models for both economies, we set the following order of shocks: cost, demand, monetary and financial. The sets of the imposed restrictions are displayed in Table 1. Note that in the tables and figures with empirical results, the alphabetical order of shocks is applied.

Combining purely statistical with sign identification restrictions leads us to a different Bayesian model, further referred to as the second MSH-VAR model. It appears that introducing these additional sign constraints removed the bimodalities almost completely (see Figures 5 and 6).

In what follows we present the results obtained from the VAR models in which sign restrictions were imposed on the structural impact matrix multiplier matrix in order to identify structural shocks. As explained above, these restrictions are needed in the MSH-VAR models for Poland and Slovakia because structural shocks can only be partially identified via heteroscedasticity (not all variances of shocks in the second state are distinct). For the Czech Republic and Hungary in turn, the regime switching heteroscedasticity seem to be redundant as one of the states dominates the other. The sign restrictions are, therefore, the only restrictions used to uncover structural shocks in these countries.

Figure 3: Marginal posteriors (histograms) and priors (lines) of the elements of the instantaneous reactions matrix (B) obtained within the MSH-VAR(2) model without sign restrictions built for Poland. The reactions' directions predicted by the economic theory are presented in brackets



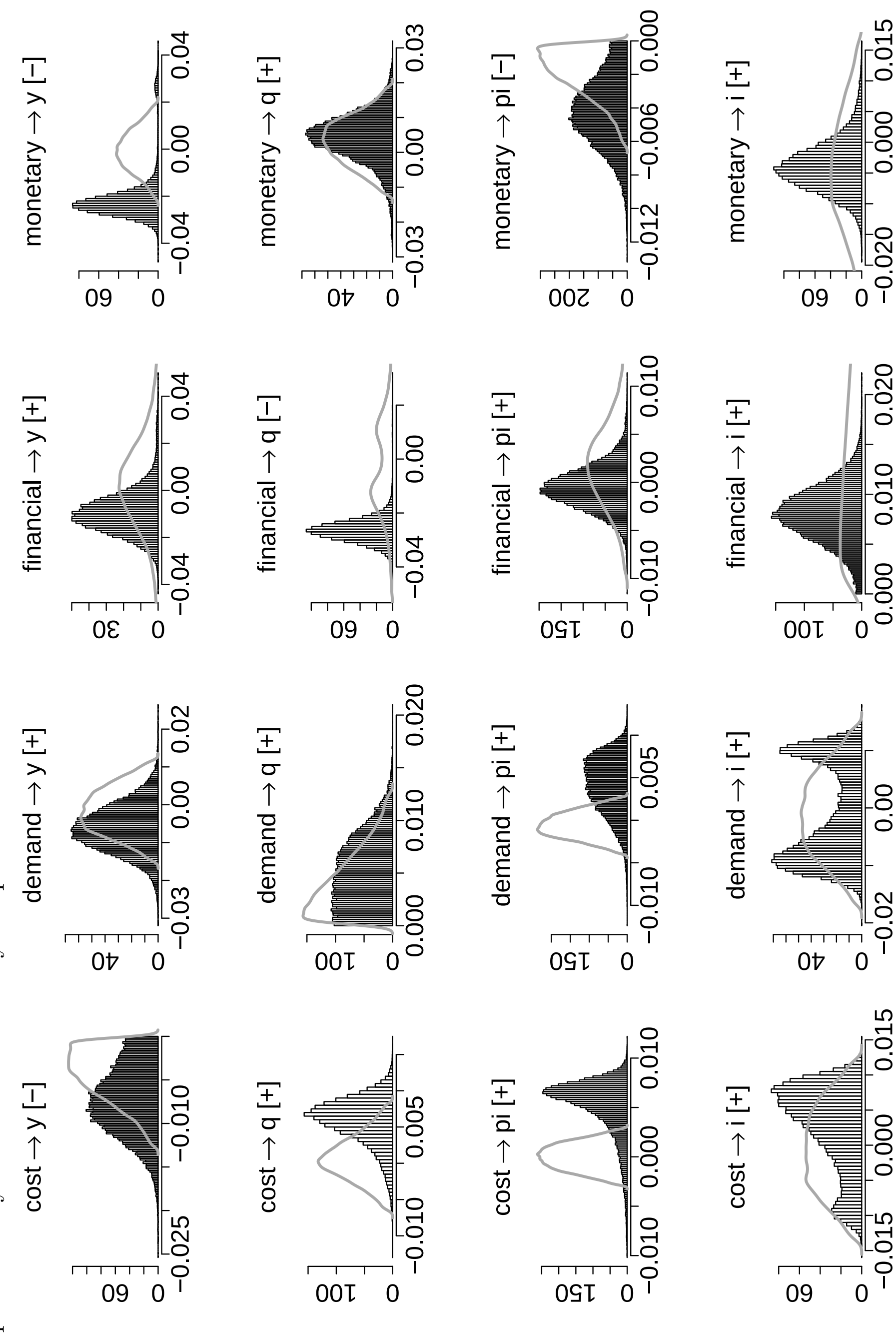


Figure 4: Marginal posteriors (histograms) and priors (lines) of the elements of the instantaneous reactions matrix (B) obtained within the MSH-VAR(2) model without sign restrictions built for Slovakia. The reactions' directions predicted by the economic theory are presented in brackets

Figure 5: Marginal posteriors (histograms) and priors (lines) of the elements of the instantaneous reactions matrix (B) obtained within the MSH-VAR(2) model combining statistical and sign restrictions built for Poland. The imposed sign are presented in brackets

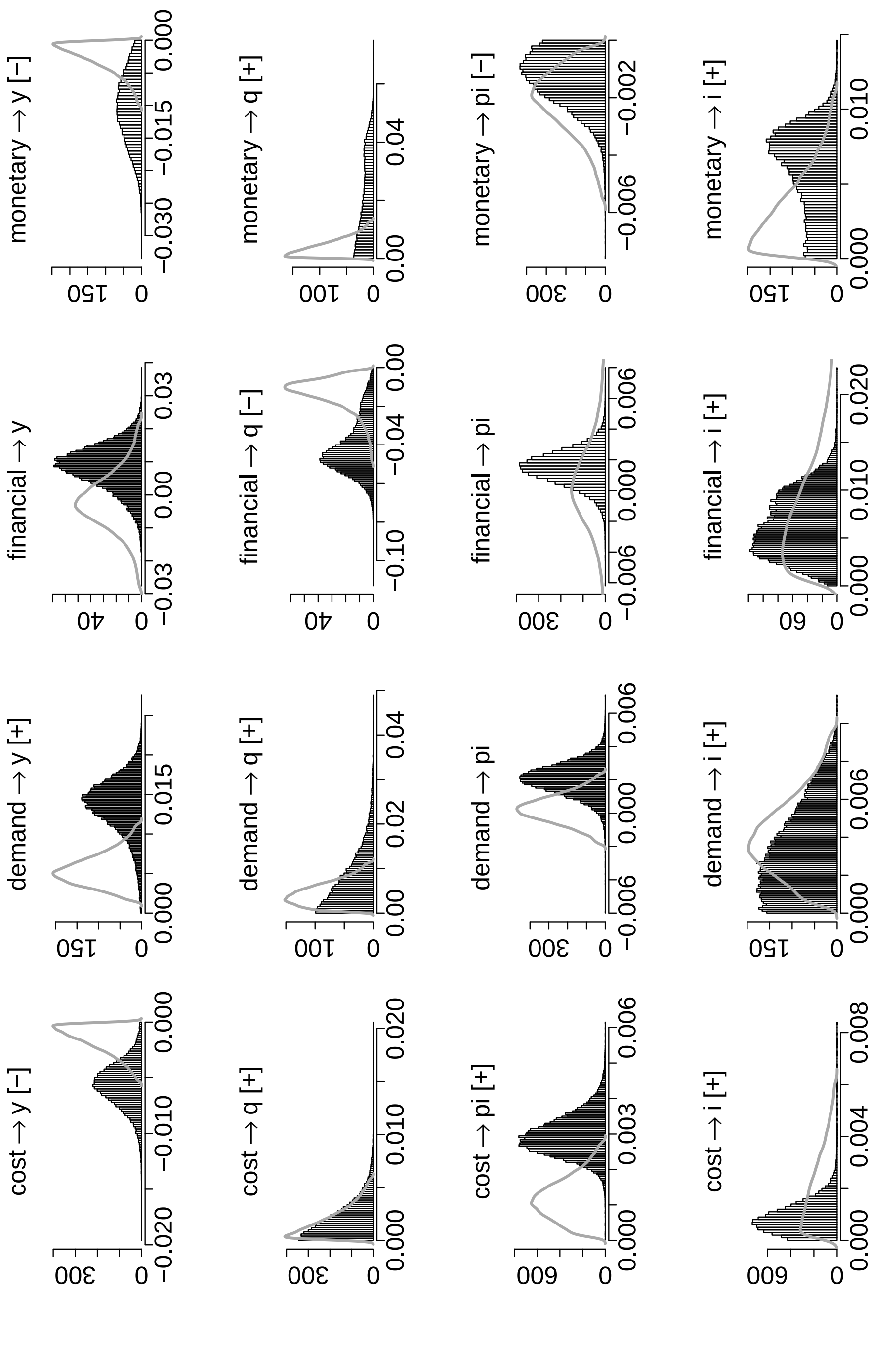


Figure 6: Marginal posteriors (histograms) and priors (lines) of the elements of the instantaneous reactions matrix (B) obtained within the MSH-VAR(2) model combining statistical and sign restrictions built for Slovakia. The imposed sign are presented in brackets

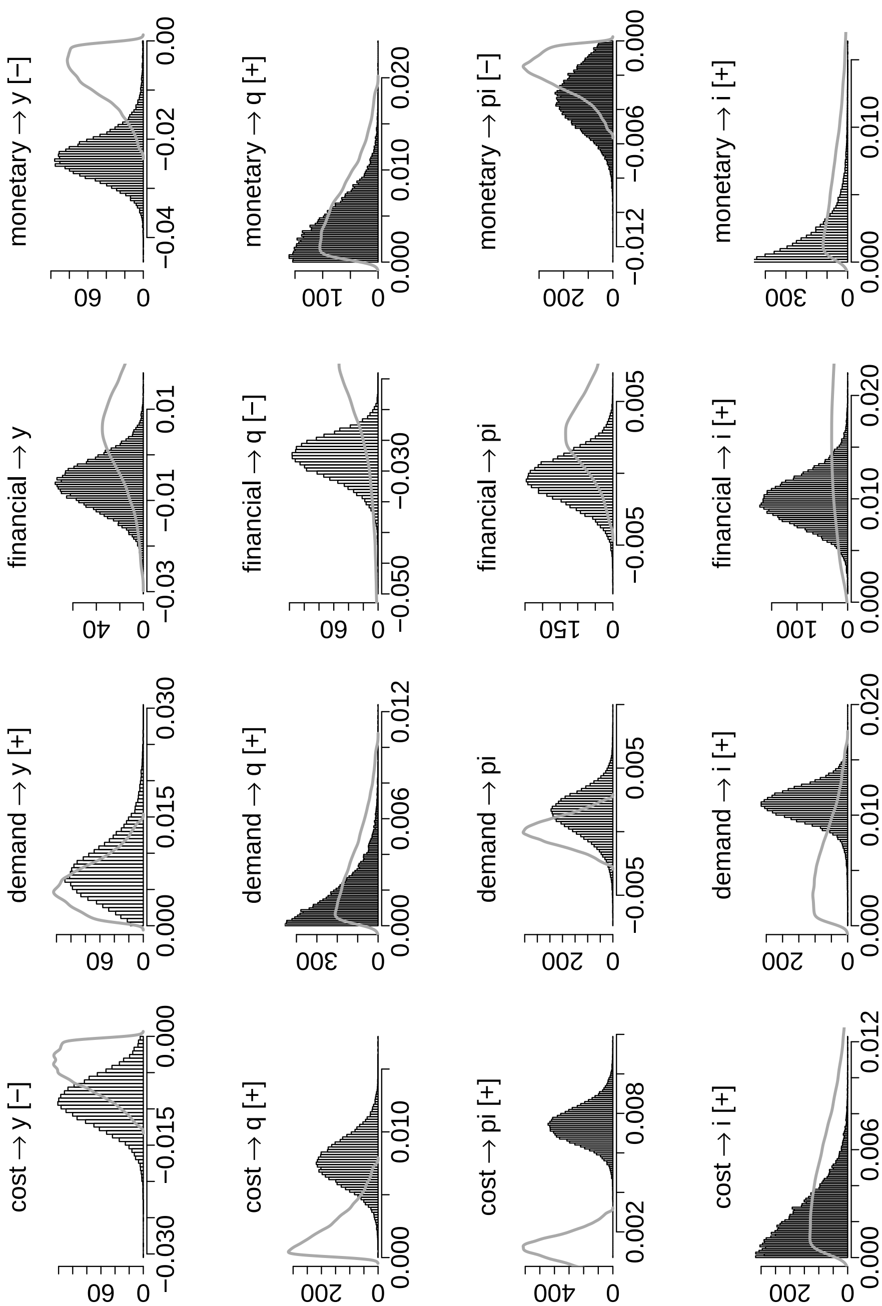
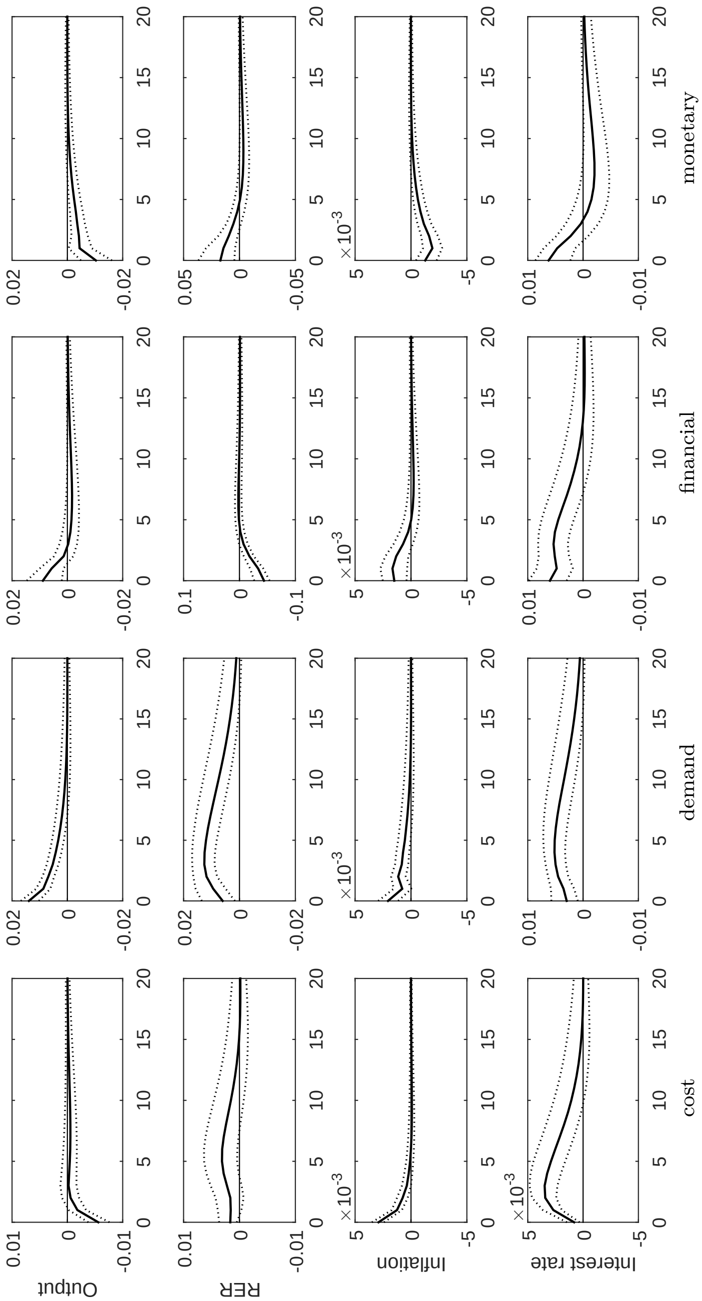


Figure 7: Impulse response functions for Poland



Note: Median (solid line) and 16th and 84th quantiles (dashed lines).

The assessment of the importance of cost, demand, financial and monetary shocks in driving the exchange rate in CEE countries is carried out with the use of conventional tools of structural vector autoregressive analysis that include structural impulse response functions, forecast error variance and historical decompositions.

5.2 Analysis of impulse response functions

The impulse response functions (IRFs) to structural shocks over a five-year period estimated for Poland are illustrated in Figure 7. Solid lines correspond to the posterior median reactions and dashed lines represent the 16th and 84th quantiles. All reactions are consistent with the model of a small open economy, what is, to a certain extent, the outcome of sign restrictions imposed. It is worthwhile pointing out that the restrictions are on the instantaneous reactions only, so trajectories of the IRFs are not determined a priori. All variables change in response to all shocks and then gradually return towards their long-run paths.

A positive cost shock raises the inflation gap and thus induces both a real appreciation of the domestic currency and an increase in the interest rate. These changes, in turn, contribute to a decrease in the output gap. The shock is only temporary, so all variables return to their long-run levels. Since the path of the inflation gap is non-negative, the relative price level goes up both in the short and long run. Thus, closing the gap in the real exchange rate requires some nominal exchange rate adjustment as the domestic currency needs to depreciate in nominal terms.

In the wake of a positive demand shock the output and inflation gaps increase, but their responses are diminished by a real appreciation of the domestic currency. Interestingly, the output gap is rather long-lived, especially in comparison with responses to other shocks, and the real exchange rate gap changes contribute to the shock absorption.

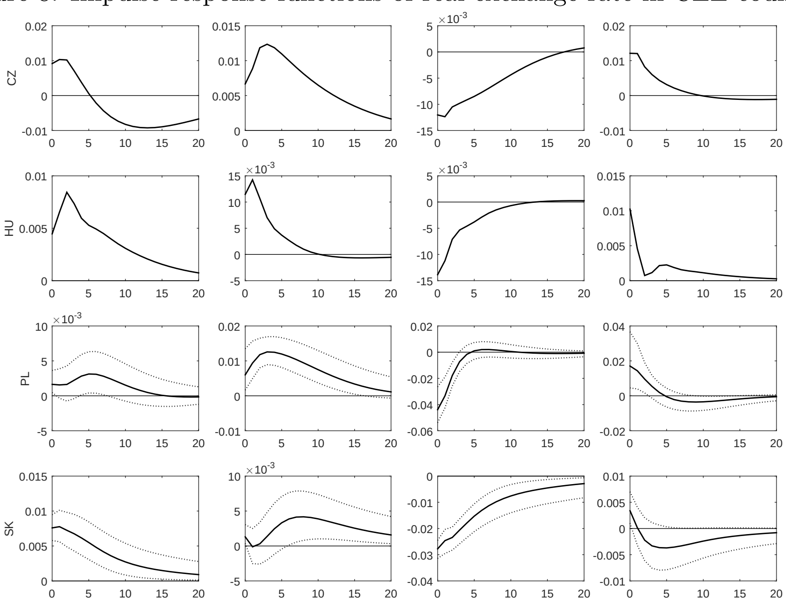
The output and inflation gaps react negatively to a positive monetary shock. The exchange rate changes in line with the interest rate. This reaction can be considered a manifestation of the exchange rate channel of monetary transmission.

A positive financial shock results in a depreciation of the domestic currency and induces the central bank to raise the interest rate. The responses of output and inflation gaps are both positive, although rather short-lived. Thus, the exchange rate changes do not seem to propagate financial shocks making them more costly to a real economy.

The impulse response functions for the Czech Republic, Hungary and Slovakia are also in line with the small open economy model. They are similar to those estimated for Poland, although there are some minor differences. The complete set of estimated IRFs is reported in Figures 14, 15, 16 in Appendix A. Output gap reactions to cost shocks in the Czech Republic and Hungary are longer-lived than in Poland. There are also some differences in responses whose direction is ambiguous in the theoretical framework: output gap response to a financial shock in the Czech Republic, Hungary and Slovakia and inflation gap reaction to a financial shock in Slovakia are all negative.

Looking closer at the real exchange rate reactions to shocks that are depicted in Figure 8, one can observe that they are similar across CEE economies. There are, however, two differences and both are in Slovakia. First, the on-impact response to a monetary shock is in accordance with the model, but at longer horizons, the currency is slightly below its long-term level. This is related to a relatively fast decrease in the interest rate gap towards zero and negative response of the inflation gap (see Figure 16 in Appendix A). The latter results in an opening of the negative real exchange rate gap which fits well the case of constrained nominal exchange rate flexibility. Second, the response to a demand shock is almost nil. This implies that the real exchange rate changes do not moderate the impact of the shock on the real economy. Both these findings are consistent with policy oriented at the exchange rate stability that has characterised Slovakia, especially after joining the European Union.

Figure 8: Impulse response functions of real exchange rate in CEE countries



Note: For the Czech Republic and Hungary the modal. For Poland and Slovakia median (solid line) and 16th and 84th quantiles (dashed lines).

5.3 Importance of structural shocks

The importance of shocks contribution can be assessed with the forecast error variance decomposition. In Table 3 results for Poland over forecast horizons of one and twenty quarters are reported, i.e. the short and long run, respectively (full results are available upon request). The results depend on the volatility regime. During the normal times (state 2) the output gap is driven mainly by real shocks (their contribution is about 80%) and the real exchange rate by financial shocks in the short run (58.9%) and real, especially demand, shocks in the long run (59.2%). Variability of the inflation gap is dominated by cost shocks (more than 75%) and the interest rate gap by monetary shocks in the short run and real, particularly cost, shocks in the long run. In turbulent times a contribution of nominal shocks is much larger than in normal times, whereas that of real, especially cost, shocks decreases substantially.

Table 3: Proportions of forecast error variance: estimates for Poland

Forecast horizon	Shocks in turbulent times				Shocks in normal times			
	cost	dem'd	fin'l	mon'y	cost	dem'd	fin'l	mon'y
Relative output gap								
1	8.0	41.8	22.0	28.2	41.0	39.2	6.1	13.7
20	6.6	43.0	24.5	25.9	36.8	42.9	7.0	13.4
Real exchange rate deviation								
1	0.3	4.2	70.5	25.0	4.7	9.6	58.9	26.9
20	2.7	23.8	51.2	22.2	21.8	37.4	24.9	15.9
Relative inflation gap								
1	46.6	24.5	15.9	13.1	87.4	8.9	1.5	2.2
20	24.5	23.1	25.2	27.2	75.1	13.5	4.1	7.4
Interest rate gap differential								
1	1.2	17.3	41.9	39.6	11.6	25.3	27.8	35.3
20	11.2	35.0	33.4	20.4	50.2	31.2	9.8	8.9

Two more important observations should be made concerning the sources of the real exchange rate variability. First, even though financial shocks are the most important source, the contribution of the two real shocks is non-negligible in normal times: it ranges from about 15% in the short run to almost 60% in the long run. Bearing in mind that these shocks account for 80% of output variability, one can argue that the exchange rate changes are not detached from changes in a real economy.

Second, in turbulent times, financial shocks become more prevalent. The real, in particular demand, shocks remain, however, important sources of both exchange rate variability and output variability (with the contribution at 25% and 50%, respectively). It does not seem that the exchange rate variability is unduly costly to an economy since output gap response to a financial shock, as measured by the IRF, is rather short-lived and abates in about two quarters.

The estimates of forecast error variance decompositions for Slovakia reported in Table 4 can be interpreted in a very similar way as those obtained for Poland. In normal times the output gap is driven by real shocks, though the contribution of monetary shocks is also large, even in the long run (almost 25%). The real exchange rate is driven by financial and real shocks, but it is cost shocks, not demand, that are more important. The inflation gap is almost exclusively influenced by cost shocks, whereas the interest rate gap by demand shocks. The latter shocks are more important than in Poland, where cost shocks' contribution is larger. Like in Poland, nominal shocks gain in importance in turbulent times and real, especially cost, shocks' contribution to variability shrinks.

For the Czech Republic and Hungary, the time-invariant error covariance VAR models are employed, so forecast error variance decompositions are not state-dependent. Results are reported in Table 5. There are some differences between the results obtained for the Czech Republic and those for Poland. The output gap is driven by real shocks to a greater extent as the relevant contribution is about 95% (and not 80%). Financial shocks are less important for the real exchange rate in the short run (35%) than in Poland and the contribution of real shocks is twice as large as in Poland (31% vs. 15%). The differences, however, wane in the long run. There is less pronounced dominance of cost shocks in variability of the inflation gap than has been observed in other CEE countries since demand and monetary shocks are more important. Finally, the interest rate gap is mainly driven by financial shocks both in the short and long term and the contribution of real shocks is rather small.

Table 4: Proportions of forecast error variance: estimates for Slovakia

Forecast horizon	Shocks in turbulent times				Shocks in normal times			
	cost	dem'd	fin'l	mon'y	cost	dem'd	fin'l	mon'y
Relative output gap								
1	12.0	10.0	8.7	69.3	37.8	14.2	1.8	46.2
20	8.3	27.8	27.2	36.6	28.2	42.1	6.0	23.6
Real exchange rate deviation								
1	7.3	0.6	89.1	2.9	53.6	2.1	40.1	4.1
20	9.6	7.5	76.4	6.5	49.6	17.5	26.2	6.8
Relative inflation gap								
1	72.6	6.0	3.6	17.8	91.3	3.8	0.3	4.6
20	49.1	10.5	22.7	17.7	83.0	8.8	2.3	5.9
Interest rate gap differential								
1	3.0	52.5	41.6	2.9	9.8	79.7	8.3	2.3
20	4.8	48.0	41.9	5.2	15.8	72.2	8.5	3.5

The case of Hungary is not too different from that of Poland. Thus, the output gap is mainly determined by real shocks, although unlike in other CEE countries the demand shocks are relatively less important than cost shocks (12% and 70% in the long run, respectively). The real exchange rate is driven by financial shocks in the short run (43%) and real shocks in the long run (60%). Like in Poland, cost shocks are behind inflation gap variability, but the contribution of monetary shocks is also important (about 30%). Like in the Czech Republic, the interest rate gap is driven by financial shocks in the short run (57%), but in the long run their importance is

Table 5: Proportions of forecast error variance: estimates for the Czech Republic and Hungary

Forecast horizon	Shocks in the Czech Republic				Shocks in Hungary			
	cost	dem'd	fin'l	mon'y	cost	dem'd	fin'l	mon'y
Relative output gap								
1	25.5	70.4	3.8	0.3	45.6	27.4	14.3	12.7
20	69.7	26.0	3.2	1.1	69.4	12.4	6.7	11.5
Real exchange rate deviation								
1	20.0	10.6	34.5	34.9	4.4	29.3	42.9	23.4
20	33.5	31.0	23.6	11.9	23.9	36.4	29.7	10.0
Relative inflation gap								
1	44.1	19.6	4.2	32.1	62.8	6.7	0.5	30.0
20	39.5	20.9	10.5	29.1	55.8	9.1	3.6	31.5
Interest rate gap differential								
1	7.9	15.1	65.6	11.4	13.1	29.4	56.7	0.8
20	4.9	23.7	67.7	3.7	53.6	26.3	16.0	4.1

Note: Estimates from the most likely (modal) model.

displaced by that of cost shocks (54% vs 16%) which makes the contributions more similar to those observed in Poland.

5.4 Sources of exchange rate fluctuations: dynamic analysis

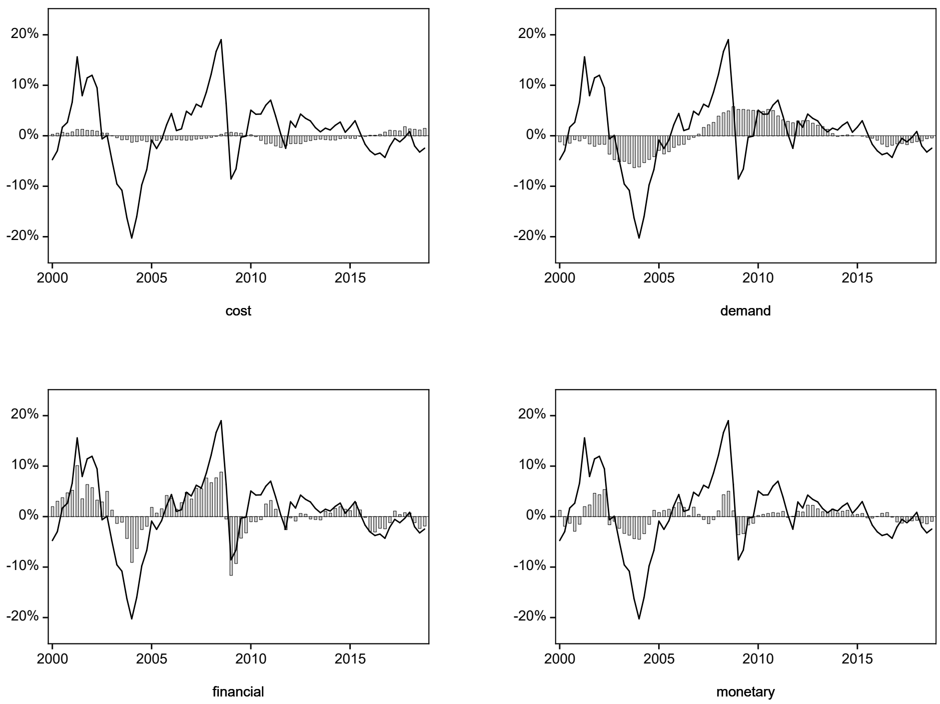
The final part of results presents historical decompositions of the real exchange rate. Unlike forecast error variance decomposition that describes the importance of shocks in the whole period (or across the states), the historical decompositions make it possible to get insight into the shocks' importance in particular subperiods.

In Section 2 two research questions are raised with regard to the real exchange rate developments in CEE countries. The first question is whether sources of exchange rate variability are similar across time, in particular in the pre-GFC and crisis periods. The second question is whether exchange rates in CEE countries are driven by similar shocks.

In Figure 9 the real exchange rate deviation from the long-term level (a cyclical component) in Poland is illustrated with a solid line. It ranged from −20% to 20% till the GFC and then fluctuated much closer to zero (±5%). Bars correspond to contributions of particular shocks to the overall deviation in a given quarter. In general, the picture is consistent with the results for the forecast error variance decomposition: financial and demand shocks account for the bulk of exchange rate variability.

Historical decompositions of the real exchange rate cyclical component in other CEE countries are reported in Figures 17, 18 and 19 in Appendix A. Basically, the deviation from the long-term level was similar in these countries to that in Poland in this sense that till the GFC it was rather large (from −10% to 15%), although not as large as in Poland, and then decreased (from −10% to 5%), but to a lesser extent than in Poland. The general picture is again in line with the forecast error variance decomposition results: financial and cost shocks account for the exchange rate variability. Unlike in Poland, the latter shocks are more important in the Czech Republic and Slovakia than demand shocks.

Figure 9: Historical decompositions of the real exchange rate in Poland



Note: Overall deviation from the long-term level depicted with a solid line.

In order to get more insights on the sources of exchange rate variability, we examined the importance of shocks across six subperiods that span 2005q1–2018q4 period. The first subperiod, 2005q1–2007q2, can be considered representative of times of real appreciation that started in the early 1990s. It is the subperiod in which cyclical component oscillated in a relatively narrow and symmetric corridor of 10 p.p. around zero. The only exception was Slovakia where the cyclical component was above zero, but remained in an even narrower corridor.

The second subperiod, 2007q3–2008q3, directly precedes the outbreak of the GFC. All CEE currencies sharply appreciated and the cyclical component of the real exchange rate increased by 7 to 13 p.p. Wide gyrations in the exchange rates took place in the third subperiod, 2008q4–2010q2, that covers the most turbulent time of the GFC. The cyclical component decreased by about 15 p.p. in all CEE countries more than reversing its rise in the pre-crisis period.

The fourth subperiod, 2010q3–2012q2, is marked by the sovereign debt crises and increased uncertainty in the euro area. The cyclical component ranged from −3% to 7%. In Slovakia, at that time a full euro area member state, the component remained in a negative part of that range.

The next subperiod starts with ECB President Mario Draghi's 'whatever it takes' speech in July 2012 and ends before the launch of the asset purchase programme by the ECB in January 2015. (For more on the shift in the monetary policy of the ECB see, e.g., Alcaraz et al. (2019), Praet (2018).) The cyclical component decreased by 1–4 p.p. in CEE countries but the Czech Republic. In the latter it decreased by more than 10 p.p. which was related to the foreign exchange market interventions of the Czech National Bank oriented at the depreciation of its currency. In fact, between November 2013 and April 2017 the CNB set the floor for the exchange rate at 27 korunas per euro. (For more on the CNB policy at that time see, e.g., Caselli (2017), Alichi et al. (2015).)

The last subperiod, 2015q1–2018q4, is the one in which the cyclical component of the real exchange rate remained or turned negative on average in all CEE countries. It can also be observed that it gravitated towards zero with time, except for the Czech Republic where it overshot zero by 5 p.p. in 2018q4.

For each subperiod we calculated the (quarterly) average contribution of a shock to the (quarterly) deviation of the real exchange rate from its long-term level. The results of this exercise are reported in Figures 10, 11, 12 and 13. In the first subperiod, 2005q1–2007q2, the Czech and Hungarian currencies were at their long-term levels, the Polish currency was slightly overvalued and the Slovak currency by more than 7% (terms over- and undervaluation are used to inform about the sign of the cyclical component of the real exchange rate). Interestingly, financial shocks had a positive impact on the real exchange rate gap ranging from 1.8 p.p. to 4.8 p.p. and only in the Czech Republic it was slightly negative. The impact of other shocks was diversified: it was negative for all shocks in Hungary, mixed in the Czech Republic and Poland and positive in Slovakia.

In the run-up to the GFC, i.e. in the second subperiod, 2007q3–2008q3, all CEE currencies became overvalued. In Poland and Slovakia, the average quarterly overvaluation was more than 12%, whereas in the two other countries it was below 8%. Financial shocks accounted for more than half of the deviation in all countries but the Czech Republic. In the latter cost shocks were more important, although the contribution of financial shocks was also important.

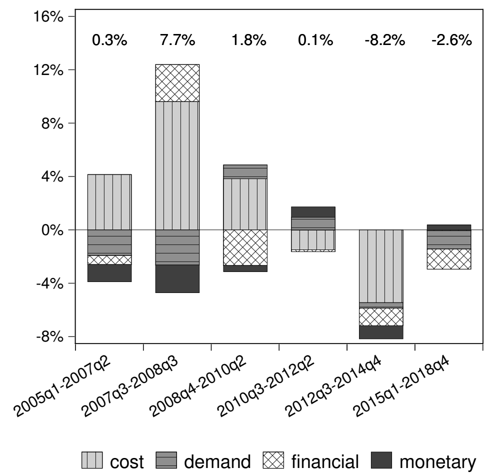
Overvaluation of CEE currencies was eliminated in the third subperiod, 2008q4–2010q2. As observed above, the long-term levels were overshot in the most turbulent phase of the GFC and only in the following quarters the resulting undervaluation was eliminated. The only exception was Slovakia where undershooting rather than overshooting could be observed and the average quarterly overvaluation remained relatively high at almost 10%. The likely reason was that the contribution of financial shock did not turn negative like in other CEE countries. Interestingly, real shocks had a positive impact on the value of all CEE currencies offsetting to a certain extent real deprecation triggered off by adverse financial shocks.

A dissimilarity in the path of the exchange rate gap in Slovakia observed in this subperiod can be related to the euro adoption. At the end of 2008, the gap was large and positive. This is in line with the finding that Slovakia entered the euro area with an exchange rate that turned out to be overvalued (Fidrmuc and Wörgötter, 2013). As the nominal exchange rate was irrevocably fixed the gap had to be closed by price adjustment and this was sluggish. For example, Lukacsy (2009) finds that prices in Slovakia are 'relatively rigid' and the 'overall price spell duration' is 15 months. Kupczyk (2018) argues that increases in the productivity-enhancing investments and reduction of employment in low productivity industries contributed to internal devaluation. This relatively lengthy adjustment also explains why the switch between regimes in Slovakia was found to be in 2011 rather than immediately upon the euro adoption (see Figure 2). The volatility of the real exchange rate remained high as the adjustment process was still under way for several quarters. Interestingly, in Poland the switch to the low volatility regime occurred earlier because under the flexible exchange rate the gap could be closed much faster.

In the fourth subperiod, 2010q3–2012q2, the real exchange rate gaps were surprisingly close to zero given uncertainty stemming from the sovereign debt crises in the euro area. The average quarterly exchange rate gap was less than 1% and only for the Polish zloty it was about 3%. This could be explained with the finding that financial and real shocks subsided in all CEE countries. In Slovakia, a negative contribution of financial shocks was more than compensated by real shocks. The latter shocks, especially demand ones, accounted for the relatively higher overvaluation in Poland.

In the fifth subperiod, 2012q3–2014q4, when central banks brought policy interest rates down to very low levels, exchange rate gaps decreased and in the Czech Republic, it moved sharply into a negative range. The large negative gap in the Czech Republic was quite likely related to the shift in the monetary policy strategy of the CNB that decided to defend koruna against appreciation. The contribution of monetary shocks, however, was quite small in comparison to those of cost and financial shocks. The relatively small importance of monetary shocks in this subperiod is a bit puzzling. The likely explanation could be that the monetary policy of the CNB was well explained to financial market participants and credible. Thus, the targeted negative real exchange gap was driven by financial shocks. The exchange rate gap was in the negative territory but close to zero in the last subperiod, 2015q1–2018q4. Almost all shocks contributed to small undervaluation in CEE countries. The largest change was observed in the Czech Republic where the average quarterly gap increased to −2.6%

Figure 10: Contribution of shocks to the real exchange rate fluctuations in the Czech Republic



Note to Figures 10–13: Quarterly average contributions. Numbers above the bars correspond to the overall quarterly average deviation from the long-term level.

Figure 11: Contribution of shocks to the real exchange rate fluctuations in Hungary

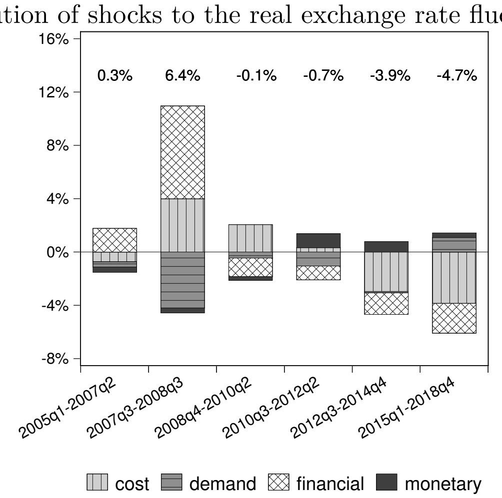


Figure 12: Contribution of shocks to the real exchange rate fluctuations in Poland

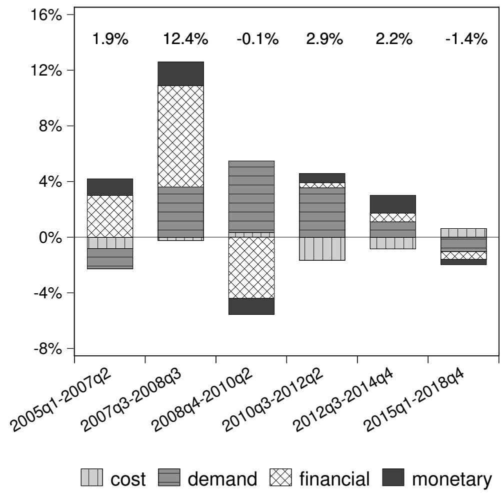
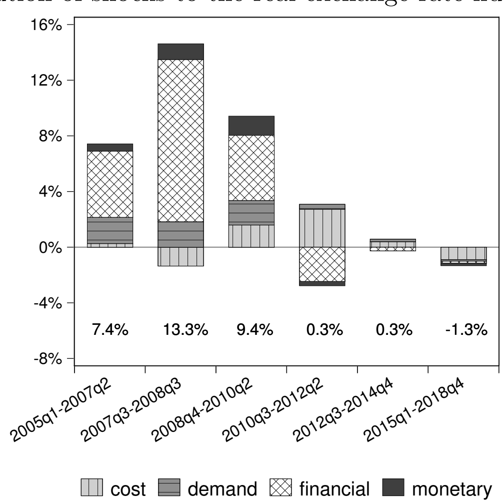


Figure 13: Contribution of shocks to the real exchange rate fluctuations in Slovakia



from −8.2% observed in the previous subperiod. Only in Hungary the exchange rate was slightly pushed away from its long-term level by cost and financial shocks.

6 Conclusions

In this paper the sources of real exchange rate variability in four CEE countries are examined. According to a stylised view, CEE currencies appreciated in real terms in the economic transition period in the 1990s and 2000s. Such a clear appreciation trend was interrupted by the global financial crisis (GFC) and then in the 2010s, it flattened out (if not transformed into a weak deprecation). Wide gyrations of CEE exchange rates in the late 2000s consisted of strong appreciation in the run-up to the crisis and even stronger deprecation during the global financial turmoils. We raise two research questions concerning the importance of shocks behind real exchange rate variability: (1) Whether sources of this variability are similar across time, especially in the pre-crisis and crisis periods; and (2) Whether similar shocks drive the exchange rates across CEE countries.

Our analysis is motivated by the gap between the 'consensus' view that nominal shocks are the main driver of the real exchange rate and empirical evidence that does not lend much support to this idea. In order to bridge the gap, we employ the modelling strategy that distinguishes financial shocks and allows on shifts in shocks' volatility. These features enable us to demonstrate that the relative importance of real and nominal shocks is not inconsistent with the 'consensus', although the gap cannot be explained completely. Our main findings in support of this claim can be summarised in four points. First, in line with existing literature real shocks, i.e. cost and demand, are found to be important drivers of real exchange volatility in the whole period whereas the contribution of monetary shocks is negligible. The short-run volatility, however, is accounted for by financial shocks, which lends support to the 'consensus'. The important role of financial shocks can be observed in particular in the run-up to and during the GFC. In Poland and Slovakia, for which the MSH-VAR models prove useful, these shocks became a single major source of exchange rate variability in the high-volatility regime.

Second, financial shocks amplify the real exchange rate gap caused by the real shocks, that is the sign of contribution of financial shocks is rarely different from that of a stronger of real shocks. This can be observed in particular in the run-up to the GFC. In the crisis, such an amplification vanishes in all CEE countries except Slovakia where this effect is not observed until the euro area debt crisis. This finding is in line with intuition: in normal times financial shocks play a background role but in financial crises they become a dominant force behind the exchange rate volatility.

Third, the real exchange rate gaps are quite similar across CEE countries. The most similar in this regard, as well as with respect to the importance of shocks, are the Czech Republic and Hungary. The important difference between them occurred when the CNB maintained the floor for the exchange rate (2013–2017). The gap in

Poland turns out not too different from that in the Czech Republic and Hungary. The contribution of financial shocks, however, is smaller in the post-crisis period and, unlike in the Czech Republic and Hungary, demand shocks are more important than cost ones.

Four, Slovakia is the most different from other CEE countries in terms of real exchange rate developments. Starting in 2009 the real exchange rate remained very stable which was related to irrevocable fixing of the nominal exchange rate and the euro adoption. This was the reason behind a relatively large overvaluation in the initial period of euro area membership and a lengthy adjustment of the real exchange rate.

Our analysis enables us to put forward some policy implications as well. Given that all countries but Slovakia pursed relatively flexible exchange rate regimes, our results imply that exchange rate management is not effective in curbing real exchange rate gaps. When the Národná Banka Slovenska, the central bank of Slovakia, kept its currency in the ERM II, gaps were not smaller than in other CEE countries. The case of one-side exchange rate band adopted by the CNB in 2013–2017 also lends support to this claim. Moreover, even though financial shocks are dominant drivers of exchange rate variability, especially in the short term and in turbulent times, the contribution of real shocks is significant and they become the dominant source of variability over the long run. Thus, policy actions that contain exchange rate fluctuations entail not only benefits but also costs as the adjustment to real shocks can become more sluggish due to price stickiness and/or prevalence of deflationary processes. The net balance of cost and benefits remains an interesting field for further research, all the more that national currencies are still in use in the Czech Republic, Hungary and Poland and prospects of the euro adoption are unclear.

Finally, we see some room for improvement also in terms of the methodology underlying our study. Arguably, for some economies it may be the case that capturing adequately the dynamics of conditional covariances may require resorting to more elaborate and flexible volatility processes than a two-state Markov chain, such as (multivariate) GARCH, SV or their hybrid structures. Extending our current methodology to the VAR-MSV-MGARCH models, particularly for the Czech Republic and Hungary (for which no clear regime switches emerged in our analysis), appears a valid and promising avenue for further research, not only as regards statistical identification of shocks, but also bridging the gap between the 'consensus' and empirical evidence.

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A Supplementary material

Below we provide additional information in Tables and Figures that is summarized in the main text.

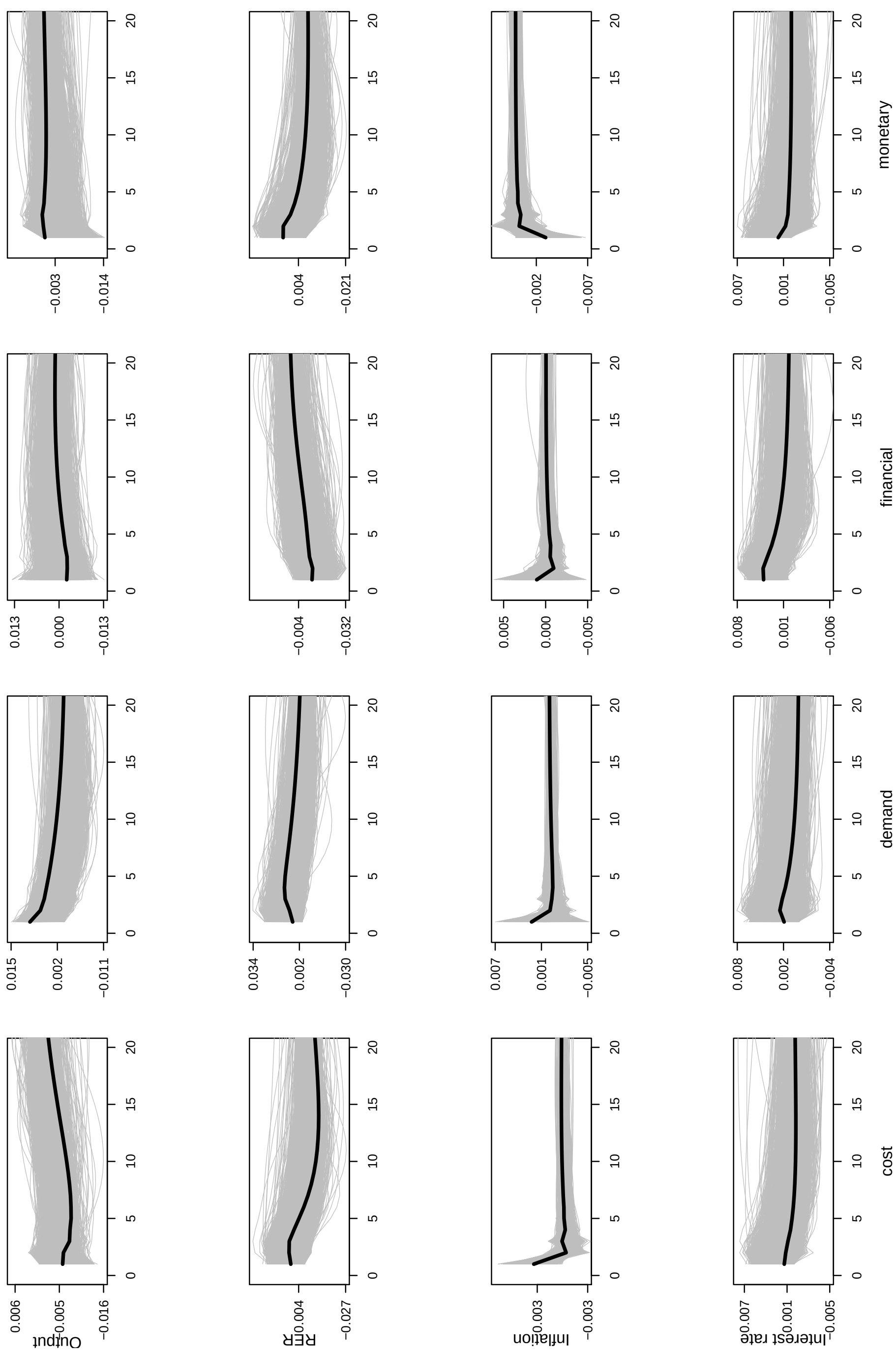
Table 6: Data description

Variable	Description	Source
Real GDP	Gross domestic product at market prices, chain linked volumes, index 2005 = 100, not seasonally adjusted data; time span: 1995q1–2018q4, for the Czech Rep. 1996q1–2018q4	Eurostat
Nominal exchange rate	Average nominal exchange rate, index 2005 = 100; an increase is an appreciation of domestic currency against the euro (ECU); time span: 1995m01–2018m12	based on Eurostat data
Price level	Harmonised index of consumer prices (All-items HICP), index 2005 = 100, not seasonally adjusted data; time span: 1996m01–2018m12	Eurostat
Nominal interest rate	Three-month money market nominal interest rate; time span: 1995q1–2018q4, for Slovakia 1995q3–2018q4; average of four adjacent quarters used for missing values for Hungary (2004q3) and Slovakia (1997q3)	Eurostat (1995q1–2015q3) and OECD (2015q4–2018q4)
Relative output gap	A difference between domestic and euro area (log) real GDP gaps (Hamilton method)	Authors' calculations
Real exchange rate deviation	The (log of the) real exchange rate calculated as the nominal exchange rate corrected for price ratio; average quarterly data obtained from monthly data and used to construct a transient component (deviation) (Hamilton method)	Authors' calculations
Relative inflation gap	A difference between domestic and euro area quarterly inflation rate gaps (Hamilton method); inflation rate obtained from the (log of the) HICP index (quarter-on-quarter rates)	Authors' calculations
Interest rate gap differential	A difference between domestic and euro area (log) interest rate gaps (Hamilton method)	Authors' calculations

Table 7: Descriptive statistics, 2000–2018

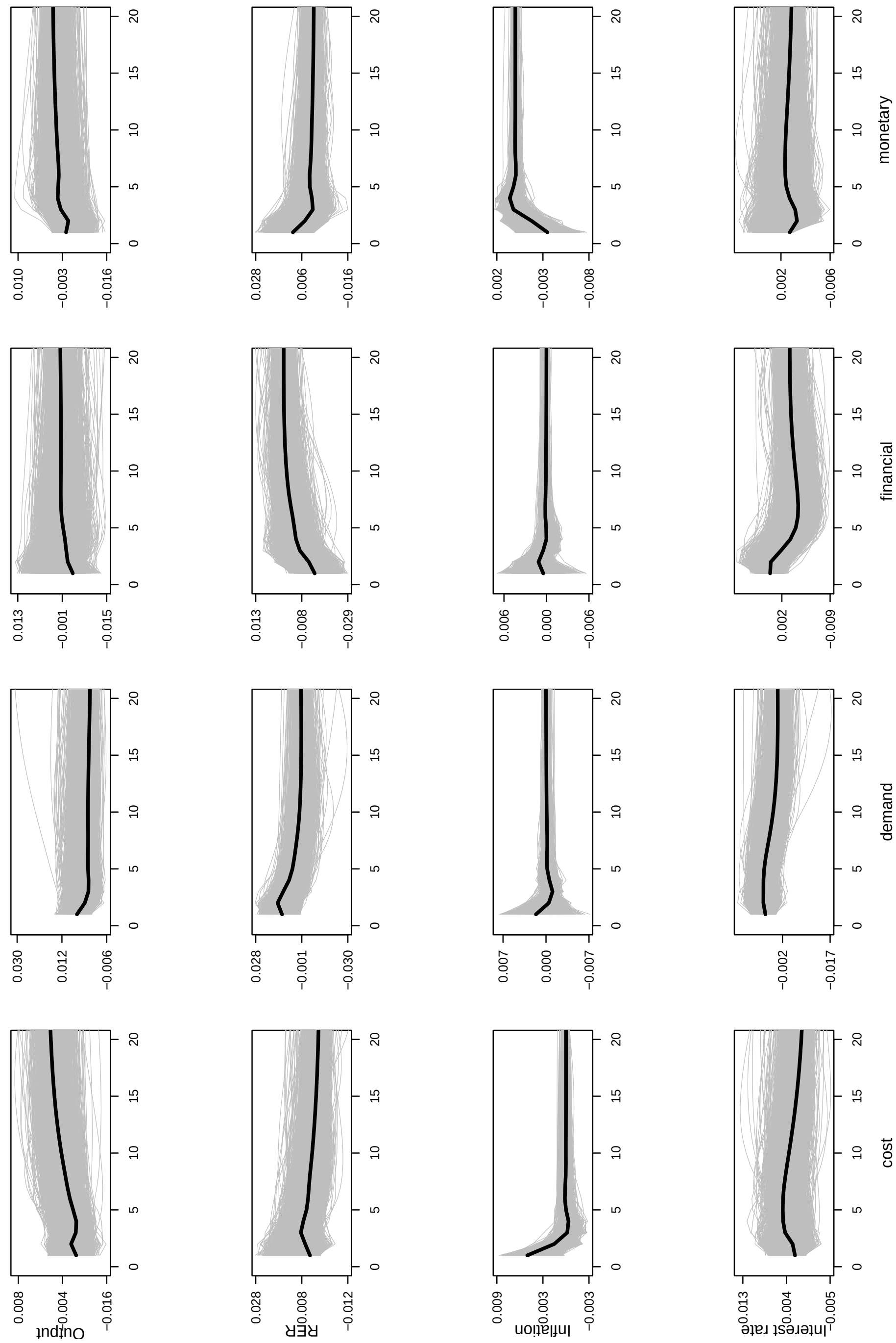
Statistic	Relative output gap	Real exchange rate dev.	Relative inflation gap	Interest rate gap diff.	Relative output gap	Real exchange rate dev.	Relative inflation gap	Interest rate gap diff.
	The Czech Republic				*Hungary*			
Mean	0.0033	0.0025	0.0000	−0.0036	0.0006	0.0052	−0.0001	0.0002
Median	−0.0006	0.0036	−0.0001	−0.0015	0.0011	0.0046	−0.0001	−0.0075
Maximum	0.0646	0.1592	0.0261	0.0175	0.0449	0.1562	0.0209	0.0582
Minimum	−0.0460	−0.1075	−0.0116	−0.0507	−0.0576	−0.0710	−0.0133	−0.0359
Std. dev.	0.0255	0.0587	0.0061	0.0127	0.0290	0.0462	0.0074	0.0234
Skewness	0.4599	0.1060	1.1082	−1.8412	−0.1731	0.5328	0.4443	0.7386
Kurtosis	2.4339	2.8424	6.3273	7.7432	1.8086	3.1542	3.0819	2.7116
Observations	76	76	76	76	76	76	76	76
	Poland				*Slovakia*			
Mean	−0.0022	0.0091	−0.0003	−0.0018	0.0018	0.0087	−0.0013	−0.0036
Median	−0.0061	0.0106	−0.0012	0.0028	−0.0042	−0.0109	−0.0030	−0.0025
Maximum	0.0572	0.1901	0.0150	0.0326	0.1050	0.1628	0.0353	0.0375
Minimum	−0.0634	−0.2027	−0.0168	−0.0613	−0.1286	−0.0678	−0.0216	−0.0869
Std. dev.	0.0244	0.0683	0.0059	0.0210	0.0441	0.0512	0.0093	0.0175
Skewness	0.3783	−0.2758	0.2564	−1.0205	−0.3305	0.8862	1.9823	−1.5277
Kurtosis	3.0218	4.4240	3.1383	3.9049	3.7114	2.9034	8.9120	9.5624
Observations	76	76	76	76	76	76	76	76

Figure 14: Impulse response functions for the Czech Republic



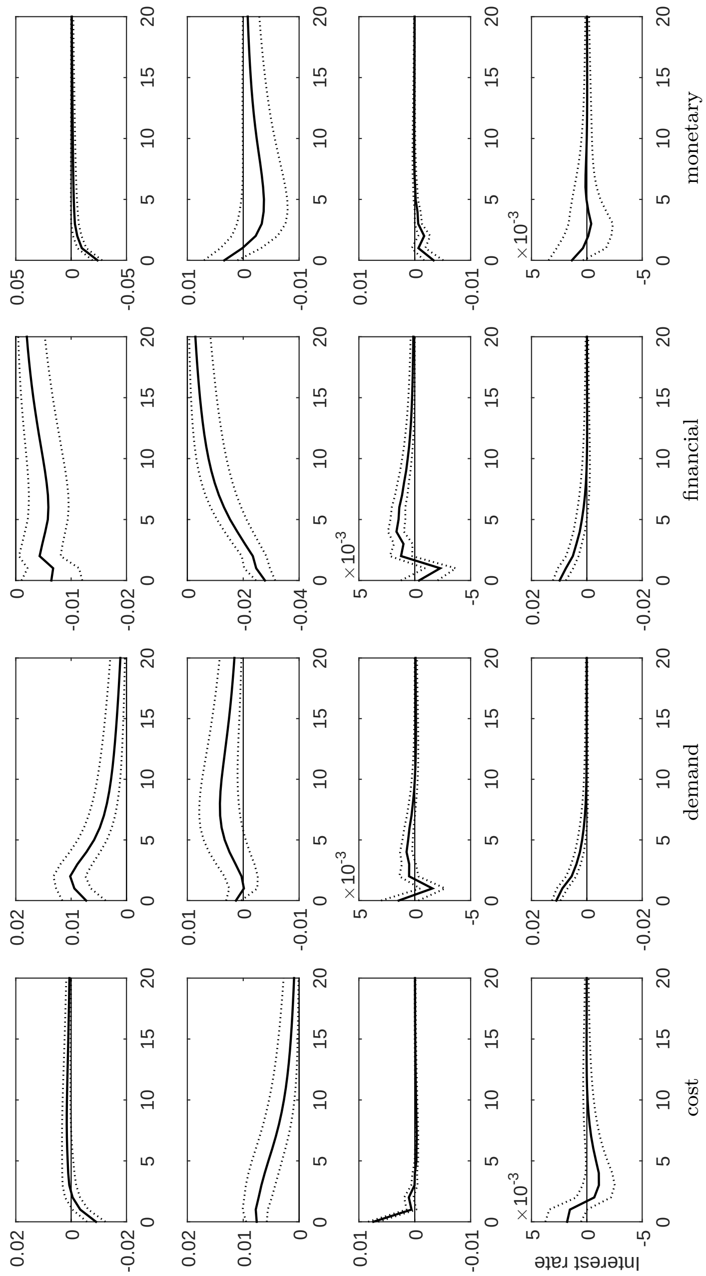
Note: Modal response (thick black line) and 68% credible sets.

Figure 15: Impulse response functions for Hungary



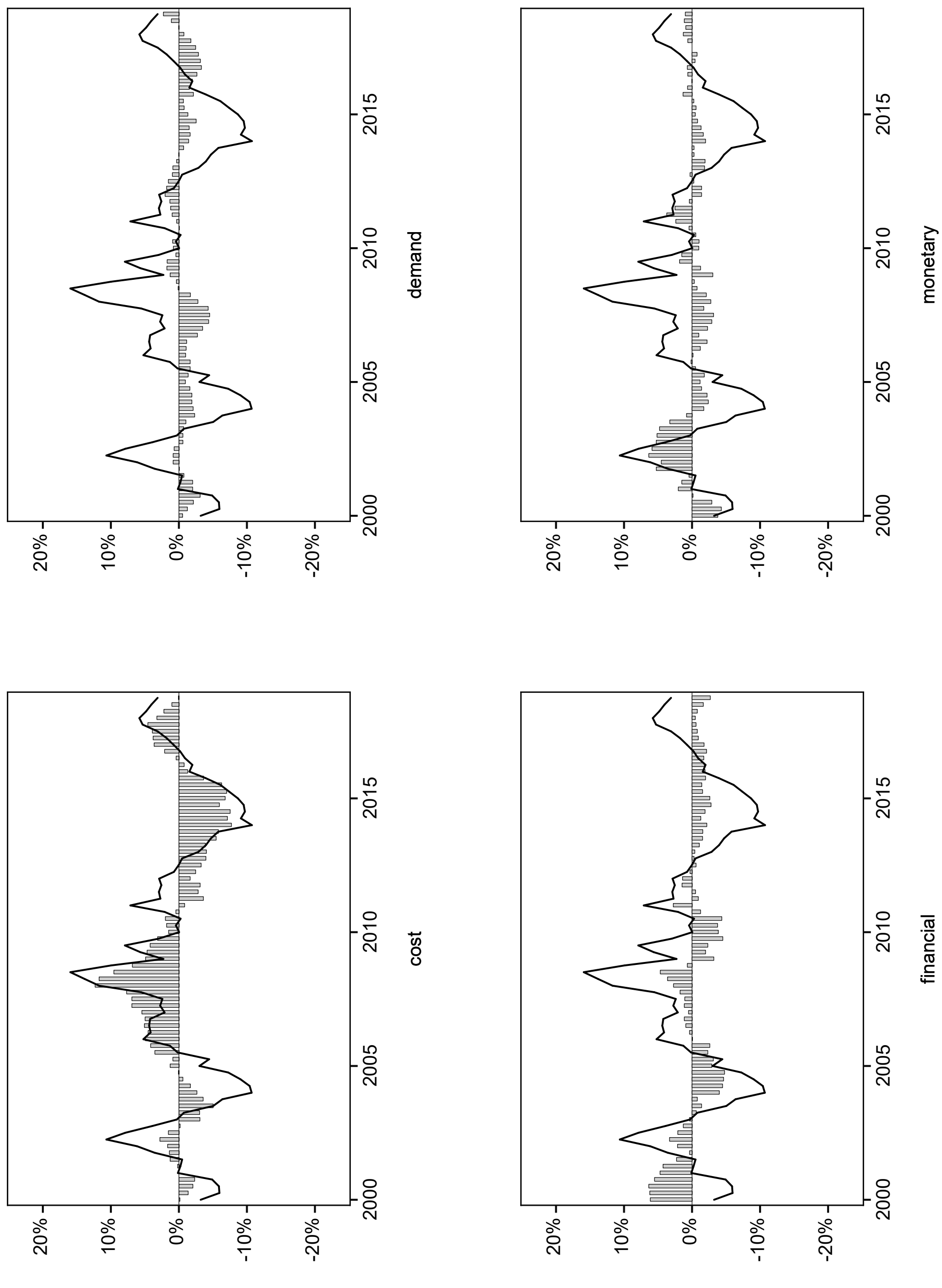
Note: Modal response (thick black line) and 68% credible sets.

Figure 16: Impulse response functions for Slovakia



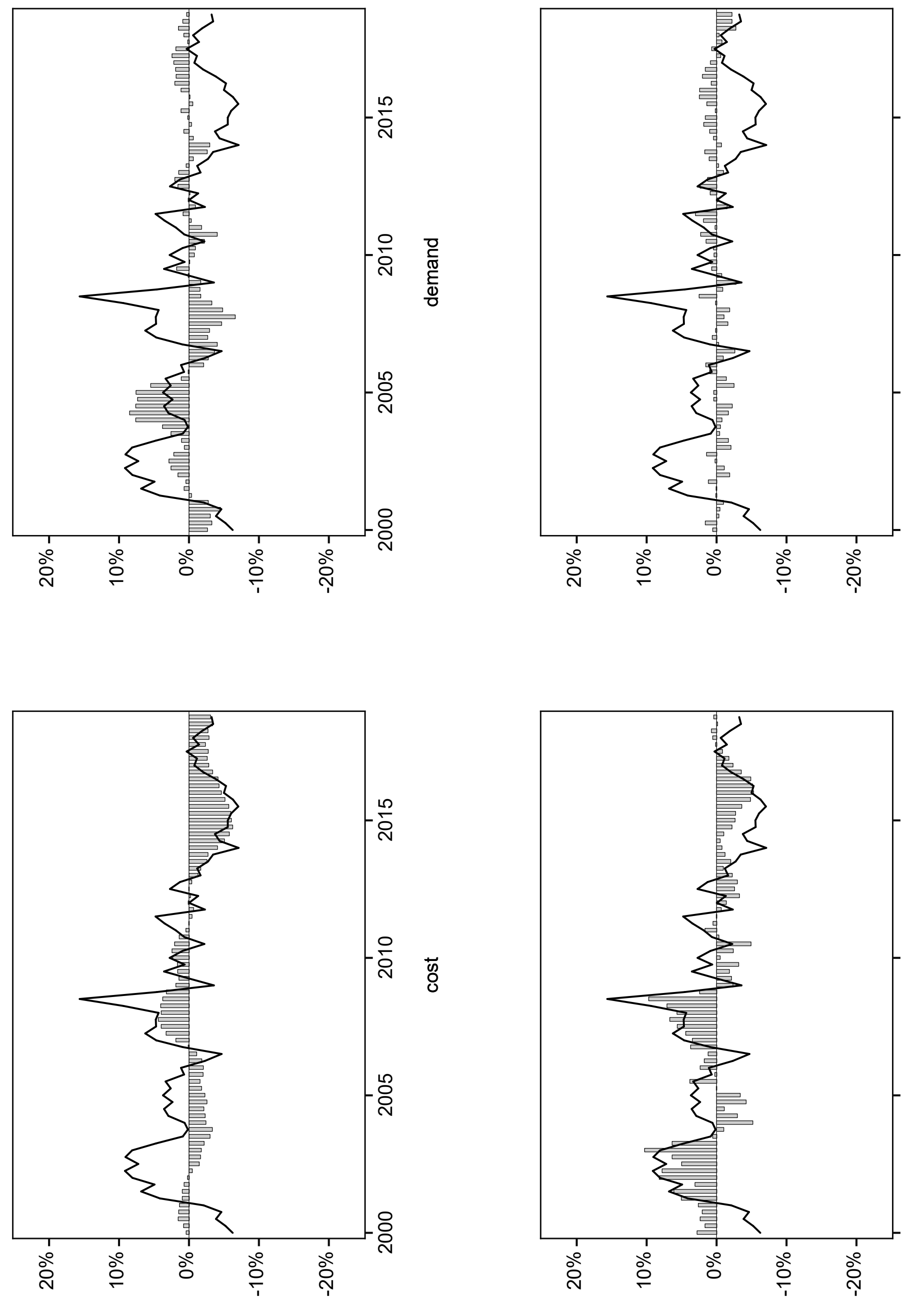
Note: Median (solid line) and 16th and 84th quantiles (dashed lines).

Figure 17: Historical decompositions of the real exchange rate in the Czech Republic



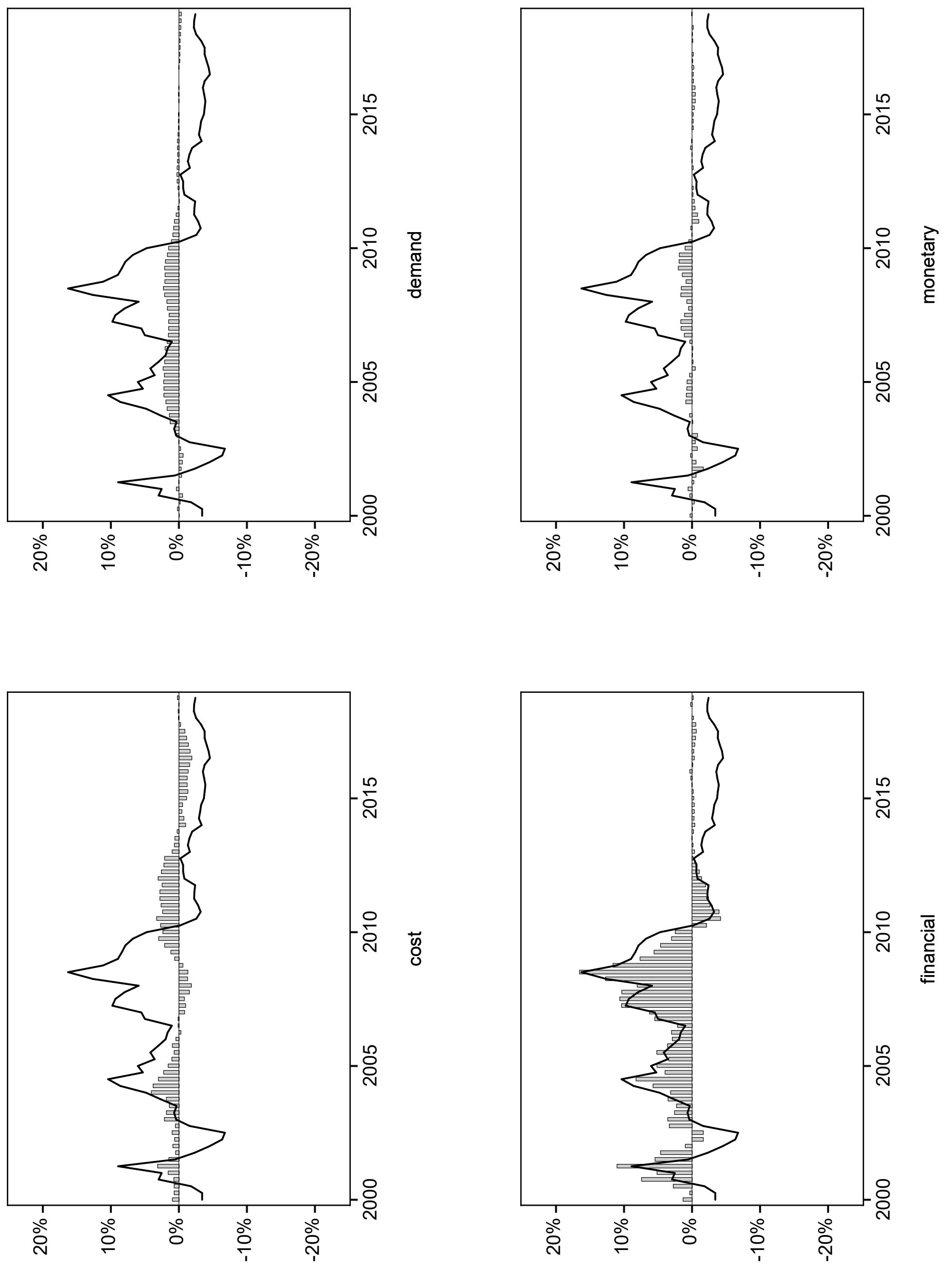
Note: Overall deviation from the long-term level depicted with a solid line.

Figure 18: Historical decompositions of the real exchange rate in Hungary



Note: Overall deviation from the long-term level depicted with a solid line.

Figure 19: Historical decompositions of the real exchange rate in Slovakia



Note: Overall deviation from the long-term level depicted with a solid line.