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How much of the Macroeconomic Variation in Eastern Europe is Attributable to External Shocks?

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HOW MUCH OF THE MACROECONOMIC VARIATION IN EASTERN EUROPE IS ATTRIBUTABLE TO EXTERNAL SHOCKS?

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Abstract. We decompose by origin the sources of the variation in real aggregate output and aggregate price level in the Czech Republic, Hungary and Poland. We find that a sizable fraction of the variation is attributable to external shocks, especially so for aggregate price level. We show that euroarea interest rate shocks can account for a significant fraction of the external spillover effects. We conclude that theoretical models of advanced transition economies and policy rules for these economies should feature a prominent role for external shocks.

1. Introduction

An important task of open-economy macroeconomics is to quantify how much of the macroeconomic variation in small open economies originates abroad. Evidence on this issue can guide policymakers who must decide how closely to track external developments as well as theorists who want to know whether to feed domestic or external shocks into their models. A related goal is to assess which external shocks matter the most.

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Thus far there is little evidence on these issues for the former centrally-planned, transition economies in Central and Eastern Europe. This paper takes a step toward filling the gap. The paper proceeds in two steps. In the first step, we provide estimates of the fraction of the variation in real aggregate output and aggregate price level in the Czech Republic, Hungary and Poland (henceforth CHP) that can be attributed to external shocks. In the second step, we estimate to what extent interest rate shocks originating in the euroarea (in Germany) can account for the spillovers to CHP from the rest of the world.

Our goal in the first step is to collect stylized facts in the form: $x\%$ of the variance in real aggregate output (aggregate price level) in the Czech Republic (Hungary, Poland) originates abroad. Our model for each of the three transition economies is a vector autoregression (VAR). To measure external shocks we use prices of commodities traded in the world market and key macroeconomic variables in Germany. Germany is by far the largest neighbor and the main trading partner of each of the three transition economies we look at. Within a VAR model for a given transition economy, we test the hypothesis that external variables in the VAR are *Granger causally prior* (henceforth GCP) with respect to domestic variables. It turns out that we find support for the GCP restriction in each of the three transition economies. We then impose the GCP restriction and decompose the sources of the aggregate fluctuations in each transition economy into domestic and external.

Our analysis in the first step is reduced-form in the sense that we do not provide a structural interpretation for shocks driving the VAR dynamics, other than that these shocks are of domestic or external origin. No further structural identification – necessarily arbitrary to some degree – is required to compute the variance decomposition of interest to us. Nevertheless, our results do help choose between two competing theories of aggregate fluctuations in CHP. According to one, external shocks are an important source of these fluctuations. Theoretical models of CHP and policy rules for CHP will be seriously incomplete if they abstract from external developments. According to the other view, external shocks are of minor importance for understanding the macroeconomic dynamics in transition economies.[1]

[1]Perhaps the key factors are e.g. domestic productivity shocks that cause a movement of resources from the state to the private sector.

We find support for the former view. We estimate that external shocks account for about 20-50% of the short-run variance in aggregate price level in CHP between 1992 and 2004. The long-run estimate is about 60-85%. The short-run estimate for real aggregate output is about 15-20%, whereas the long-run estimate is about 25-50%. Thus we find that a sizable fraction of the variation in real aggregate output and aggregate price level in CHP can be attributed to external shocks, especially so for aggregate price level.

The recent literature on business cycles in emerging markets emphasizes the importance of external interest rate shocks (Neumeyer and Perri (2004), Uribe and Yue (2003)).[2] In the second step, we orthogonalize our external shocks in order to investigate to what extent interest rate shocks originating in the euroarea (in Germany) can account for the spillovers to CHP from the rest of the world. We find that euroarea interest rate shocks can account for a sizable fraction of the spillovers. Euroarea interest rate shocks account for more than one-third of the joint effects of external shocks on real aggregate output in CHP. Euroarea interest rate shocks account for about 50% of the joint effects of external shocks on aggregate price level in the Czech Republic and for more than two-thirds in Hungary and Poland.

The estimated effects of euroarea interest rate shocks on real aggregate output and aggregate price level in Germany are consistent with interpreting them as euroarea monetary policy shocks. Remarkably, we find that euroarea interest rate shocks have the same qualitative effects in CHP as in Germany. Real aggregate output and aggregate price level in CHP show the same pattern of gradual decline after a positive euroarea interest rate shock as real aggregate output and aggregate price level in Germany.

Our conclusions are in line with what other authors have found for small open economies in the developed world and for emerging markets outside of Europe. Cushman and Zha (1997) study Canada, using world commodity prices and key macroeconomic variables in the United States to measure external shocks. Cushman and Zha find that a sizable fraction of about 50-75% of the variance in real aggregate output in Canada is attributable to external shocks. Del Negro and Obiols-Hums

[2]This literature focuses on Latin America and interest rate shocks that originate in the United States.

(2001) study Mexico, likewise using world commodity prices and key macroeconomic variables in the United States as measures of external shocks. The estimates of Del Negro and Obiols-Hums attribute to external shocks about 75-85% of the variance in real aggregate output and in aggregate price level in Mexico between 1976 and 1994. Both papers, like us, use VAR models with multiple external variables and the GCP restriction. Uribe and Yue (2003) estimate that U.S. interest rate shocks explain as much as 20% of movements in aggregate activity in emerging markets in Latin America and Asia.

We also make contact with the recent macroeconometric literature on transition economies. In the most closely related paper, Korhonen (2003) estimates bivariate VARs for real aggregate output in a number of transition economies and an index of real output in the euroarea. His variance decomposition suggests that about 10-15% of the variance in real aggregate output in CHP is attributable to external shocks. Our estimate is larger, possibly because we use more than a single variable to measure external shocks. Another difference between Korhonen (2003) and this paper is that we test the GCP restriction and use it to justify the variance decomposition.[3]

A number of recent papers focus on the related question whether macroeconomic fluctuations in Central and Eastern Europe are correlated with the fluctuations in Western Europe. Fidrmuc (2004) reports correlations between detrended aggregate real output in Germany and in a number of transition economies. Fidrmuc and Korhonen (2003) formulate structural VAR models for real aggregate output growth and inflation in several transition economies, identifying aggregate supply and demand shocks (in the spirit of Blanchard and Quah (1989)). Fidrmuc and Korhonen then study the correlation of East European shocks with the euroarea shocks, obtained analogously. Dibooglu and Kutan (2001) similarly use the Blanchard-Quah identification to study the sources of real exchange rate movements in transition economies. The difference between our study and the work of Fidrmuc and Korhonen (2003) and Dibooglu and Kutan (2001) is that these authors rely on the Blanchard-Quah identification that is stronger than the GCP restriction and the orthogonalization that we use. The Blanchard-Quah identification has been criticized in the literature (see Faust

[3]In an early study, Boone and Maurel (1999) find that sizable fraction of the variation in unemployment in transition economies is explained by external factors.

and Leeper (1997) and Cooley and Dwyer (1998)). Furthermore, the GCP restriction is easily testable and our results turn out not depend on the orthogonalization that we use. One can view both approaches as complementary.

Dibooglu and Kutan (2005) and Golinelli and Rovelli (2005) present estimated theoretical models of the same transition economies that we study. Laxton and Pesenti (2003) calibrate an open economy equilibrium model using data from the Czech Republic. None of these papers, however, decomposes the sources of aggregate fluctuations in CHP by origin. Furthermore, the theoretical models that these papers employ naturally make much stronger assumptions than our VAR models. Again, one can view both approaches as complementary.[4]

The paper proceeds as follows. The next section presents the econometric model. Section 3 discusses the results. Section 4 concludes and discusses implications of the results for theoretical modeling and policy. In the first Appendix we list several stylized facts about CHP relevant for the analysis. The second Appendix provides details of the data. In the third Appendix we discuss inference on parameters of the model. We do not give in this paper institutional details on CHP or an account of the history of transition, as this has been done in numerous other papers (see for example Kutan and Brada (2000)).

2. The model

We formulate and test the hypothesis that the Czech Republic, Hungary and Poland are each a small open economy. Textbooks define a small open economy as one that takes exogenously external variables. We conjecture that the textbook model is a plausible description of CHP, countries dependent on conditions abroad yet too small to influence them. In the first Appendix we list several stylized facts

[4]Among other recent macro empirical papers on transition economies, Christoffersen, Slok and Wescott (2001) show that foreign prices and exchange rates are useful for forecasting inflation in Poland. Siklos and Abel (2002) fit a Taylor rule-like monetary policy reaction function for Hungary, finding a feedback from the real exchange rate to policy. Brada and Kutan (2001) and Brada, Kutan and Zhou (2005) study cointegration between macroeconomic variables in transition economies and in Western Europe. Kutan and Yigit (2004) study convergence between transition economies and Western Europe in panel data. Eickmeier and Breitung (2005) study synchronization of Central and East European economies with the euroarea using a dynamic factor model.

relevant for the analysis. First, CHP are relatively open economies, the extent of their trade openness being comparable to Korea and Mexico, for example. Second, CHP conduct two-thirds or more of their trade with member countries of the European Union, of which about half with Germany alone. Third, CHP – even taken together – are small in terms of income relative to Germany.

Consider a macroeconomic variable $z(t)$ in a small open economy. The textbook model suggests a decomposition of the sources of the variation in $z(t)$ into domestic and external, beyond the influence of the small economy. It is convenient to organize the discussion around the following structural model of a linear, stochastic, dynamic form (omitting a constant and other deterministic terms):

$$\sum_{s=0}^{p} A(s)y(t-s) = \varepsilon(t), \tag{2.1}$$

$t = 1, .., T$, where $y(t)$ and $\varepsilon(t)$ are $(M \times 1)$ each, $A(s)$ matrices are $(M \times M)$, $A(0)$ is non-singular, and $\varepsilon(t)$ is Gaussian with zero mean and:

$$E[\varepsilon(t)\varepsilon(t)' \mid y(t-s), s > 0] = I_M.$$

We interpret $\varepsilon(t)$ as the vector of structural disturbances (such as changes in policy, technology and tastes) that generate the data. The model in its general form is familiar from the structural VAR literature (e.g. Leeper, Sims and Zha (1996)).

In our case, the model contains m_1 domestic variables in a small open economy ($y_1(t)$ vector) and m_2 variables external to the small economy ($y_2(t)$ vector), with $m_1 + m_2 = M$. We can partition the model into a domestic and an external block using the notation:

$$y(t) = \begin{bmatrix} y_1(t) \\ y_2(t) \end{bmatrix}, \varepsilon(t) = \begin{bmatrix} \varepsilon_1(t) \\ \varepsilon_2(t) \end{bmatrix}, A(s) = \begin{bmatrix} A_{11}(s) & A_{12}(s) \\ A_{21}(s) & A_{22}(s) \end{bmatrix},$$

for all $s = 0, 1, ..., p$, with $y_i(t)$ and $\varepsilon_i(t)$ each of dimension $(m_i \times 1)$, and $A_{ij}(s)$ of dimension $(m_i \times m_j)$, $i = 1, 2$ and $j = 1, 2$. The small open economy assumption implies the restriction $A_{21}(s) = 0$, for all $s = 0, 1, ..., p$ (see Cushman and Zha (1997) and Zha (1999)). This is the restriction making the $y_2(t)$ vector *block exogenous*, i.e. domestic variables are postulated not to enter the external block equations either contemporaneously or with lags. External variables are a linear combination of external shocks only, whereas domestic variables are generated both by domestic and external disturbances.

Multiplying (2.1) by the inverse of $A(0)$ yields a reduced-form VAR model for $y(t)$ that we can estimate:

$$y(t) = \sum_{s=1}^{p} B(s)y(t-s) + v(t), \tag{2.2}$$

with $B(s)$ of size $(M \times M)$, and $v(t)$ an $(M \times 1)$ Gaussian vector with zero mean and:

$$E[v(t)v(t)' \mid y(t-s), s > 0] = \Omega = A^{-1}(0) \left[A^{-1}(0)\right]'.$$

After partitioning into a domestic and an external block, we obtain:

$$\begin{aligned} y_1(t) &= \sum_{s=1}^{p} B_1(s)y(t-s) + v_1(t) \\ y_2(t) &= \sum_{s=1}^{p} B_2(s)y(t-s) + v_2(t), \end{aligned} \tag{2.3}$$

with $B_1(s) = [B_{11}(s)\ B_{12}(s)]$, $B_2(s) = [B_{21}(s)\ B_{22}(s)]$, $B_{ij}(s)$ of size $(m_i \times m_j)$, $i = 1, 2$, and $j = 1, 2$. Analogously, we partition Ω into four Ω_{ij}'s. It is straightforward to see that the block exogeneity restriction implies a testable restriction on (2.3). Namely, the block exogeneity restriction implies the restriction that the external vector $y_2(t)$ is *Granger causally prior* with respect to the domestic vector $y_1(t)$: formally, $B_{21}(s) = 0$ for $s = 1, ..., p$. Thus fluctuations in the small economy do not help predict fluctuations in external variables.

Below we test the GCP hypothesis for each of the three transition economies we look at. We take failure to reject the GCP hypothesis as evidence for the small open economy model. Upon failing to reject the GCP hypothesis, we decompose by origin the sources of variation in $y_1(t)$ in each of the three countries. The fraction of the variation $y_1(t)$ due to innovations in all elements of $\varepsilon_2(t)$ jointly provides a measure of the extent of external dependence of the small open economy.

We estimate (2.3) *separately* for the Czech Republic, Hungary and Poland. For each economy, the vector of domestic variables $y_1(t)$ includes a measure of real aggregate output and a measure of aggregate price level. We include only two variables for each country since the available data series are short and we save degrees of freedom by estimating a small-scale VAR. We focus on real aggregate output and aggregate price level since these variables are of central importance in macroeconomics, even in textbook models of aggregate demand and supply. Specifically, we use real industrial output and consumer price level. Real industrial output is the only measure of real aggregate output available on monthly basis. Consumer price index is the measure of aggregate price level of interest to a central bank that aims to keep stable prices of goods consumed by households.

The vector of external variables $y_2(t)$ contains two kinds of variables. First, two prices of commodities traded in the world market: an index of export prices of non-fuel primary commodities and the price of crude oil. Changes in commodity prices can be expected to lead to inflationary or deflationary pressures in CHP. Both series of commodity prices are measured in U.S. dollars, but we convert them to euros before estimation.[5] Second, we include three variables meant to summarize macroeconomic conditions in Germany: measures of real aggregate output, aggregate price level and interest rate in Germany. Specifically, we use real industrial output, producer price level and call money interest rate. Producer prices are likely to match the notion of prices of internationally traded goods better than consumer prices, because prices of nontradables account for a greater share of consumer prices than of producer prices. We use call money interest rate to account for the possibility that monetary policy decisions in the euro area (before introduction of the euro, in Germany) can be a source of fluctuations in CHP. Germany is by far the largest neighbor, and also the main trading partner of each of the three transition economies we look at; see the first Appendix for some relevant stylized facts. Thus there are five external variables in total.

All time series are seasonally adjusted, measured at monthly frequency and in logarithm, except that the German interest rate is in percentage points at an annual rate. The sample starts in January 1992 and ends in December 2004. Each equation in the VAR contains a constant term. Details of all the data series are given in the second Appendix. All the data series are plotted in Figure 1.

[5]This allows for the possibility that CHP react to changes in the exchange rate between the U.S. dollar and the euro (the German mark) in addition to reacting to fluctuations in commodity prices.

We use Bayesian inference and employ data in levels. Bayesian inference does not rely on asymptotic results. This feature seems attractive in the context of relatively small samples that we have at our disposal. Furthermore, Bayesian inference allows us to remain agnostic about the presence of unit roots[6] and cointegrating relations. In particular, the Bayesian inference using data in levels allows for the possibility that cointegrating relations are present without imposing them as restrictions. Determining with high confidence whether a cointegrating relation is present would be difficult in our small samples. We prefer not to impose cointegrating restrictions and not to run the risk that imposing such restrictions erroneously may lead to incorrect inference regarding the effects of external shocks.[7] We do not include a linear trend since a VAR in M variables with constant terms and p lags but without a linear trend can account for polynomial trends up to the order Mp (see e.g. Sims and Zha (1998)). The addition of a linear trend would induce collinearity at low frequencies making inference even more uncertain than is unavoidable in the context of our relatively small samples.

We must be careful not to reject erroneously a true GCP restriction, which would lead to incorrect inference regarding the contribution of external shocks. A false rejection can arise if an external factor important for fluctuations in CHP – and *also* for the dynamics of $y_2(t)$ – is omitted from the external block. In this case $y_1(t)$ will be a linear combination of, among others, innovations in the omitted variable, and the estimates will assign spuriously to $y_1(t)$ predictive power for changes in $y_2(t)$. To avoid a false rejection and incorrect inference regarding the contribution of external shocks, it is important to include a sufficient number of external variables in the $y_2(t)$ vector.

[6]The evidence from the ADF tests and the KPSS tests is mixed and inconclusive. The null hypothesis of a unit root can be rejected for five out of eleven time series used in this study. The null hypothesis of trend stationarity cannot be rejected for two time series for which the null of a unit root cannot be rejected as well.

[7]See Sims and Uhlig (1991) for a discussion of Bayesian inference in the possible presence of nonstationarity.

3. The results

3.1. **Model specification.** We begin by examining the fit of the VAR model (2.2) for various values of p without the GCP restriction. We estimate the VAR model for each of the three countries setting alternatively $p = 6, 9, 12$. We evaluate the Laplace approximation to the log marginal likelihood for each estimated model. This method of choosing among models is consistent regardless of whether data are stationary or nonstationary (see Kim (1998)). We find that the specification with $p = 6$ achieves the best fit in each of the three countries. We focus in subsequent analysis on the specification with 6 lags. However, we examine whether our substantive conclusions are robust with respect to the choice of lag length.

We investigate the presence of serial correlation in the residuals from the estimated VAR models. Using the Bayesian approach discussed in Lancaster (2004, chapter 2) and based on the Durbin-Watson statistic, we fail to find evidence of serial correlation. We also investigate stability of the estimated VAR coefficients over time. We find that the VAR models with coefficients restricted to be constant throughout the sample achieve far better fit in terms of marginal likelihood than the VAR models that allow for a break in the middle of the sample.

3.2. **Granger causal priority of external variables.** We consider the null hypothesis that the external vector $y_2(t)$ is block exogenous with respect to the domestic vector $y_1(t)$ in the model (2.1). To test the hull hypothesis, we fit the external block of equations of the VAR model (2.3) with and without the GCP restriction. We evaluate the Laplace approximation to the log marginal likelihood for the VAR model with the GCP restriction and for the VAR model without the GCP restriction. To ensure that our conclusion is robust to lag length, we set alternatively $p = 3, 6, 9, 12$. The results are in Table 1. We find that the model with the GCP restriction achieves far better fit than the model without the restriction, for each of the three countries and for each value of p. Consider the results with $p = 6$, as an example. The log marginal likelihood of the model with the GCP restriction is 1430. The log marginal likelihood of the model without the GCP restriction is 1403 (in the case of the Czech Republic), 1397 (in the case of Hungary) and 1401 (in the case of Poland). Note that this is a log scale, so that differences of 10 or more imply extreme posterior odds ratios in favor of the GCP restriction, while differences of 1 or 2 would mean little.[8]

We also compare, for each of the three countries, the fit of a bivariate VAR model in real aggregate output and aggregate price level to the fit of the model that in addition includes our five external variables, entering contemporaneously and with lags. Using the Laplace approximation to the log marginal likelihood, we find extreme posterior odds ratios in favor of the model with the external regressors in each of the three countries. This suggests that the external variables are important for understanding the dynamics of real aggregate output and aggregate price level in CHP.

3.3. **Inference methodology.** Having found support for the GCP hypothesis, we decompose by origin the sources of the variation in $y_1(t)$ in each of the three countries. Here we follow the work of Cushman and Zha (1997) and Zha (1999) on Bayesian inference in VARs with the block exogeneity restriction; see the third Appendix and Zha (1999) for details. As always in Bayesian analysis, the posterior distribution of unknown parameters of interest is proportional to the product of the likelihood function and the prior. With a flat prior, the posterior for parameters in each block in our VAR model (2.3) is equal to the product of an inverse-Wishart density for the variance-covariance matrix of the error term and a Gaussian conditional density for the equations' coefficients. We take 1000 draws from the posterior distribution, after having imposed the GCP restriction; see the third Appendix for details.[9] For each draw, we compute impulse responses and forecast error variance decomposition of $y_1(t)$. This yields 1000 draws from the posterior distribution of impulse responses and forecast error variance decomposition. Finally, we compute percentiles of the posterior distributions. In each case, we report the median as well as 16th and 84th percentiles of the posterior, i.e. 68% probability bands. The probability bands have the usual Bayesian interpretation that the parameter of interest (e.g. the fraction of

[8]The Schwarz criterion and the Hannan-Quinn criterion also favor the model with the GCP restriction in each of the three countries and for each value of p. The Akaike criterion favors the the GCP restriction in each of the three countries for $p = 6$ and $p = 9$. We already discussed the evidence that the model with $p = 6$ achieves the best fit in each of the three countries.

[9]To ensure robustness to the presence of outliers, we make draws from a multivariate t-distribution with 3 degrees of freedom instead of drawing from a Gaussian pdf.

the variation in $y_1(t)$ attributable to external shocks) is contained within the bands with probability 68%, given the model and the data.[10]

3.4. **Variance decomposition.** In Table 2 we report the median share of the variance in real aggregate output and aggregate price level in CHP attributable to external shocks, at horizons of 6 months (that we refer to as the short-run), 12 months (the medium run) as well as 24 and 48 months (the long-run). External shocks account for about 20-50% of the short-run variance in aggregate price level in CHP. The medium-run estimate is about 40-70%. The long-run estimate is about 60-85%. The short-run estimate for real aggregate output is about 15-20%, the medium-run estimate is about 15-30%, and the long-run estimate is about 25-50%. Thus we find that a sizable fraction of the variation in real aggregate output and aggregate price level in CHP can be attributed to external shocks. This is especially true for aggregate price level. In the last fifteen years, each of the three countries experienced significant disinflation. It is remarkable that most of the variation in aggregate price level during the disinflation emerges as having originated abroad.

In Table 3 we report 68% probability bands for the share of the variance in real aggregate output and aggregate price level in CHP attributable to external shocks. That is to say, we continue looking at the same parameters as in Table 2, but switch from examining medians to examining how uncertain we are about the estimates. As is common in VAR studies, the amount of uncertainty is sizable. However, we are able to make the following statements with 84% probability. External shocks account for at least 35% of the long-run variance in aggregate price level in the Czech Republic and 70% in Hungary and Poland. External shocks account for at least 13% of the long-run variance in real aggregate output in Hungary and 30% in the Czech Republic and Poland.

We fail to find a consistent pattern of differences between the Czech Republic, Hungary and Poland. In particular, we fail to find support for the common view that the smaller countries (the Czech Republic and Hungary) are "more open" – in the sense of being more dependent on external shocks – than the larger Poland.

[10]We employ a flat prior, having decided that the popular non-flat prior due to Sims and Zha (1998) – derived with the United States and other developed economies in mind – may lead to unreliable results when imposed in our study.

Hungary's real aggregate output seems to react somewhat less to external shocks than real aggregate output in the Czech Republic and Poland. Aggregate price level in the Czech Republic seems to react somewhat less to external shocks than aggregate price level in Hungary and Poland.

Tables 2 and 3 show the estimates obtained with 6 lags. The median estimates with 12 lags are typically somewhat higher. For example, the long-run estimate for real aggregate output in Hungary increases to over 35% from 20-25%. This finding is consistent with the notion – also clear from inspecting Tables 2 and 3 – that the effects of external shocks increase over time and thus that dynamic interactions are important.

3.5. **Impulse responses.** We examine to what extent interest rate shocks originating in the euroarea (in Germany) can account for the sizable international spillover effects reported above. We suppose that the matrix $A_{22}(0)$ in the system (2.1) is upper triangular. We order variables in the external vector $y_2(t)$ as follows: call money interest rate, both commodity price series, aggregate real output and aggregate price level.

Figure 2 shows the impulse responses of real aggregate output and aggregate price level in Germany and in CHP to a positive innovation in call money rate in Germany, with $p = 6$. The first column in the figure shows that real aggregate output in Germany starts decreasing on impact and reaches the minimum of roughly -0.4% after about 18-24 months, before starting to recover. Aggregate price level begins decreasing with some delay and stays permanently lower by about 0.2%. The impulse responses suggest that we are justified in interpreting a positive euroarea interest rate shock as an exogenous tightening in euroarea monetary policy (a monetary policy shock).[11]

[11]This amounts to assuming that one-step-ahead surprise changes in the short-term interest rate are monetary policy shocks after we account for any contemporaneous response of the interest rate to both commodity price series, aggregate real output and aggregate price level. Our estimates are unaffected when variables in the vector $y_2(t)$ are reordered. Notice that, due to the small sample size, we must assume that the effects of changes in the short-term interest rate in Germany are the same before and after the introduction of the euro. This is a common assumption in the literature, see e.g. Peersman and Smets (2001) and Eickmeier and Breitung (2005).

The second, third and fourth column in Figure 2 show a remarkable pattern. First, euroarea interest rate shocks have the same effects in the Czech Republic, Hungary and Poland. Real aggregate output and aggregate price level fall by about 1% in each of the three countries. Second, euroarea interest rate shocks have the same qualitative effects in CHP as in Germany. Real aggregate output and aggregate price level in CHP show the same pattern of gradual decline seen in Germany, with the strongest effect about 2 years after the shock.

The estimated euroarea interest rate shocks account for only a modest fraction of the variation in real aggregate output and aggregate price level in Germany. Specifically, the median estimate suggests that euroarea interest rate shocks account for about 7% of the variation in real aggregate output and for about 13% of the variation in aggregate price level in Germany, at the horizon of 24 months.[12] Thus the effects shown in the first column of Figure 2, while statistically significantly different from zero, are modest.

Table 4 reports the median share of the variance in real aggregate output and aggregate price level in CHP attributable to euroarea interest rate shocks. Consider the maximum effect that takes place typically at the horizon of 24 months. Euroarea interest rate shocks account for about 9-16% of the variation in real aggregate output in CHP, somewhat more than in Germany. Aggregate price level in CHP reacts significantly more strongly to euroarea interest rate shocks than does aggregate price level in Germany. Specifically, euroarea interest rate shocks account for about 50-60% of the variation in aggregate price level in Hungary and Poland and for 26% in the Czech Republic.

Comparing Table 4 with Table 2, we conclude that euroarea interest rate shocks can account for a sizable fraction of the external spillover effects into CHP. Euroarea interest rate shocks account for more than one-third of the joint effects of external shocks on real aggregate output in CHP. Euroarea interest rate shocks account for about 50% of the joint effects of external shocks on aggregate price level in the Czech Republic and for more than two-thirds in Hungary and Poland.

[12] This finding matches the findings of the structural VAR literature regarding the effects of monetary policy shocks. See e.g. Leeper, Sims and Zha (1996) for the United States and Kim (1999) for a cross-country comparative study including Germany.

4. Conclusions

We decompose by origin the sources of the variation in real aggregate output and aggregate price level in the first 15 years of transition. We find that a sizable fraction of the variation is attributable to external shocks, especially so for aggregate price level. According to our estimates, external shocks account for about 60-85% of the long-run variance in aggregate price level in the Czech Republic, Hungary and Poland. The same estimate for real aggregate output is about 25-50%. We also find that a sizable fraction of the external effects is accounted for by euroarea interest rate shocks. Remarkably, euroarea interest rate shocks have the same qualitative effects in the Czech Republic, Hungary and Poland as in Germany.

Our results have implications for theoretical modeling and policy. Our results suggest that theoretical models of advanced transition economies should feature a prominent role for external shocks. Modelers face the decisions of what kind of transmission mechanism for external shocks to emphasize. Our results suggest that external interest rates should play an important role in the transmission mechanism.

Our results do not imply that domestic policy has been unimportant in CHP. First, our estimates leave open the possibility that domestic policy *shocks* account for a nontrivial fraction of the macroeconomic variation in CHP, especially in real aggregate output and in the short run. Investigation of the effects of domestic policy shocks in CHP is an important topic for future research. Second, it is an open question whether the macroeconomic dynamics in CHP would have been very different, had the *systematic reaction* of domestic policy to external shocks been much different. Investigation of the systematic reaction of policy in CHP to external shocks – especially euroarea interest rate shocks – emerges as an important topic for future research.

Policymakers have been debating to what extent CHP are ready to join the euroarea. Consider our result that a sizable fraction of the macroeconomic variation in CHP is attributable to external shocks. Consider also our results that euroarea interest rate shocks are important and that they generate qualitatively similar responses in CHP as in Germany. These results can be seen as an indication that the costs of adopting the euro are unlikely to be large for CHP. On the other hand, consider our result that the effects of euroarea interest rate shocks are greater in CHP than in

Germany. Consider also our result that a nontrivial fraction of the macroeconomic variation in CHP is attributable to domestic shocks, especially for real aggregate output. These results can be seen as suggesting caution in adopting the euro. Assessing the costs and benefits of adopting the euro remains an important topic for future research.

Appendix A. Some stylized facts about the Czech Republic, Hungary and Poland

Trade openness (half the sum of exports and imports of goods over GDP, in 1996): the Czech Republic 44%, Hungary 39%, Poland 26% – compare with Korea 29% and Mexico 28% (OECD, 2000).

The direction of exports of the Czech Republic, in 2000: 69% to the E.U., 40% to Germany; imports: 62% from the E.U., 32% from Germany (IMF 2002a).

The direction of exports of Hungary, in 2001: 74% to the E.U., 37% to Germany; imports: 58% from the E.U., 26% from Germany (IMF 2002b).

The direction of exports of Poland, in 2001: 69% to the E.U.; imports: 61% from the E.U. (IMF 2002c).

The combined GDP of CHP (at current prices and exchange rates in 1996) equals 9% of Germany's GDP (OECD, 2000).

Appendix B. The data

All data was downloaded from Datastream in June 2005. The following time series were used.

For commodity prices: index of export prices of non-fuel primary commodities WDI76NFDF, price of crude petroleum WDI76AADF. Both commodity price series, originally measured in U.S. dollars, were converted into euros prior to estimation. The following data series were used in the conversion: exchange rate between U.S. dollar and euro EMI..AG., exchange rate between German mark and U.S. dollar BDI..DE., exchange rate between German mark and euro BDI..EA..

For Germany: index of real industrial production BDINPRODG, producer price index BDI63...F, call money interest rate monthly average BDSU0101.

For the Czech Republic: index of real industrial production CZIPTOT.H, consumer price index CZP1PITAF.

For Hungary: index of real industrial production HNI66...F, consumer price index HNI64...F.

For Poland: index of real industrial production POIPTOT.H, consumer price index POI64...F.

The data on real output and price level in the Czech Republic, Hungary and Poland and the data on price level in Germany, downloaded in seasonally unadjusted form, were deseasonalized prior to estimation using the X11 multiplicative command in RATS.

Appendix C. Inference methodology

We describe the posterior probability density functions from which we make draws. The joint pdf of $y(t)$, conditional on the data until $t-1$, is the product of the marginal density of $y_2(t)$ and the conditional density of $y_1(t)$ given $y_2(t)$:

$$\begin{aligned} p\left[y(t) \mid y(t-s), s>0\right] &= p\left[y_2(t) \mid y(t-s), s>0\right] \\ &\quad \times p\left[y_1(t) \mid y_2(t), y(t-s), s>0\right], \end{aligned}$$

where:

$$p\left[y_2(t) \mid y(t-s), s>0\right] = (2\pi)^{-m_2/2} \left|\Omega_{22}\right|^{-1/2}$$
$$\times \exp\left\{(-1/2)\left[y_2(t) - \sum_{s=1}^{p} B_2(s) y(t-s)\right]' \Omega_{22}^{-1} \left[y_2(t) - \sum_{s=1}^{p} B_2(s) y(t-s)\right]\right\},$$

and

$$p\left[y_1(t) \mid y_2(t), y(t-s), s>0\right] = (2\pi)^{-m_1/2} \left|\Sigma\right|^{-1/2}$$
$$\times \exp\left\{(-1/2)\left[y_1(t) - \mu_1(t)\right]' \Sigma^{-1} \left[y_1(t) - \mu_1(t)\right]\right\},$$

The terms in the conditional density of $y_1(t)$ are defined as follows:

$$\Sigma = \Omega_{11} - \Omega_{12}\Omega_{22}^{-1}\Omega_{21},$$

$$\mu_1(t) = D_0 y_2(t) + \sum_{s=1}^{p} D_1(s)\, y(t-s),$$

$$D_0 = \Omega_{12}\Omega_{22}^{-1},$$

$$D_1(s) = B_1(s) - \Omega_{12}\Omega_{22}^{-1} B_2(s).$$

Observe that Σ has size $(m_1 \times m_1)$, $\mu_1(t)$ is $(m_1 \times 1)$, D_0 is $(m_1 \times m_2)$ and $D_1(s)$ is $(m_1 \times M)$.

Using the expressions above, we can write an alternative expression for the VAR (2.3):

$$
\begin{aligned}
y_1(t) &= D_0 y_2(t) + \sum_{s=1}^{p} D_1(s)\, y(t-s) + e_1(t) \\
y_2(t) &= \sum_{s=1}^{p} B_2(s) y(t-s) + v_2(t),
\end{aligned}
\tag{C.1}
$$

where Σ is the variance-covariance matrix of $e_1(t)$. Parameters of (C.1) are linked to those of (2.3) via the one-to-one mappings given above.

We can write the system (C.1) in matrix notation as follows:

$$
\begin{aligned}
Y_1 &= X_1 D + E_1 \\
Y_2 &= X_2 \mathfrak{B}_2 + V_1,
\end{aligned}
\tag{C.2}
$$

where Y_1 and E_1 are of dimension $(T \times m_1)$, Y_2 and V_1 have size $(T \times m_2)$, X_1 is $(T \times (Mp + m_2 + J))$, where J is the number of non-stochastic regressors, X_2 is $(T \times (m_2 p + J))$ (assuming the GCP restriction), and matrices of coefficients D and $\mathfrak{B}_2$ have dimensions $((Mp + m_2 + J) \times m_1)$ and $((m_2 p + J) \times m_2)$, respectively.

Let us define:

$$\widehat{D} = \left(X_1' X_1\right)^{-1} X_1' Y_1,$$

and

$$\widehat{\mathfrak{B}}_2 = \left(X_2' X_2\right)^{-1} X_2' Y_2.$$

We make draws of Ω_{22}^{-1} from a Wishart distribution with parameter:

$$\left[\left(Y_2 - X_2 \widehat{\mathfrak{B}}_2\right)' \left(Y_2 - X_2 \widehat{\mathfrak{B}}_2\right)\right]^{-1},$$

and degrees of freedom $T - (m_2 p + J)$ (i.e. T less the number of regressors in each equation), and afterwards draws of $\mathfrak{B}_2$ from its conditional distribution:

$$N\left(vec\left(\widehat{\mathfrak{B}}_2\right), \Omega_{22} \otimes \left(X_2' X_2\right)^{-1}\right).$$

We make draws of Σ^{-1} from a Wishart distribution with parameter:

$$\left[\left(Y_1 - X_1 \widehat{D}\right)' \left(Y_1 - X_1 \widehat{D}\right)\right]^{-1},$$

and degrees of freedom $T-(Mp+m_2+J)$, and afterwards draws of D from its conditional distribution:

$$N\left(vec\left(\widehat{D}\right),\Sigma\otimes\left(X_1'X_1\right)^{-1}\right).$$

For each draw of Ω_{22}^{-1}, we find its upper triangular Choleski square root $A_{22}(0)$ satisfying $\Omega_{22}=A_{22}^{-1}(0)\left[A_{22}^{-1}(0)\right]'$. For each draw of Σ^{-1}, we find its upper triangular Choleski square root $A_{11}(0)$ satisfying $\Sigma=A_{11}^{-1}(0)\left[A_{11}^{-1}(0)\right]'$. We obtain $A_{12}(0)$ by noting that $A_{12}(0)=-A_{11}(0)D_0$.

Each complete set of draws described above constitutes a single draw of the coefficients of the VAR model (2.3) with the GCP restriction and single a draw of the matrix $A(0)$. We use 1000 such draws to compute medians and probability bands.

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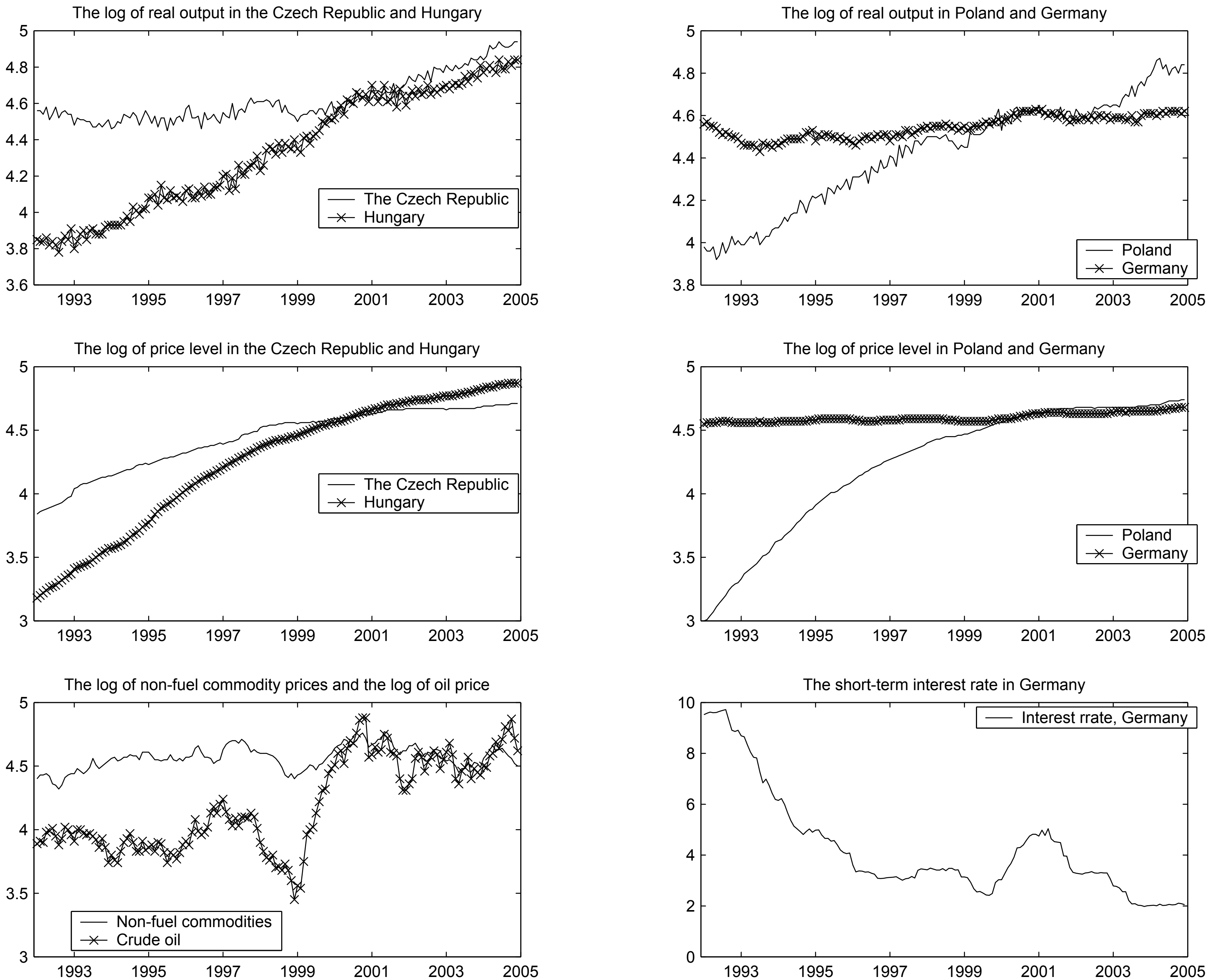
Figure 1: The data used in the estimation.



See Section 2 and the second Appendix for details of the data and for the data sources.

Table 1: Results of testing the Granger causal priority restrictions

	Lag length	**Model with the GCP restriction**	**Model without the GCP restriction for:** Czech Republic	Hungary	Poland
Log marginal likelihood	*3*	1543	1527	1528	1525
	6	1430	1403	1397	1401
	9	1306	1260	1257	1267
	12	1206	1153	1142	1165

Note: Each entry in the table reports the log marginal likelihood for a given model. See Section 3.2 for a discussion.

Table 2: Variance decomposition (medians)
The contribution of **external shocks jointly** to forecast error variances

	Horizon (months)	**Czech Republic**	**Hungary**	**Poland**
Real output	*6*	16.3%	14.0%	20.6%
	12	21.5%	15.1%	29.3%
	24	34.9%	21.6%	38.4%
	48	49.5%	24.7%	49.6%
Price level	*6*	28.5%	22.3%	52.9%
	12	39.3%	51.0%	72.5%
	24	54.4%	75.7%	81.4%
	48	59.2%	87.6%	82.2%

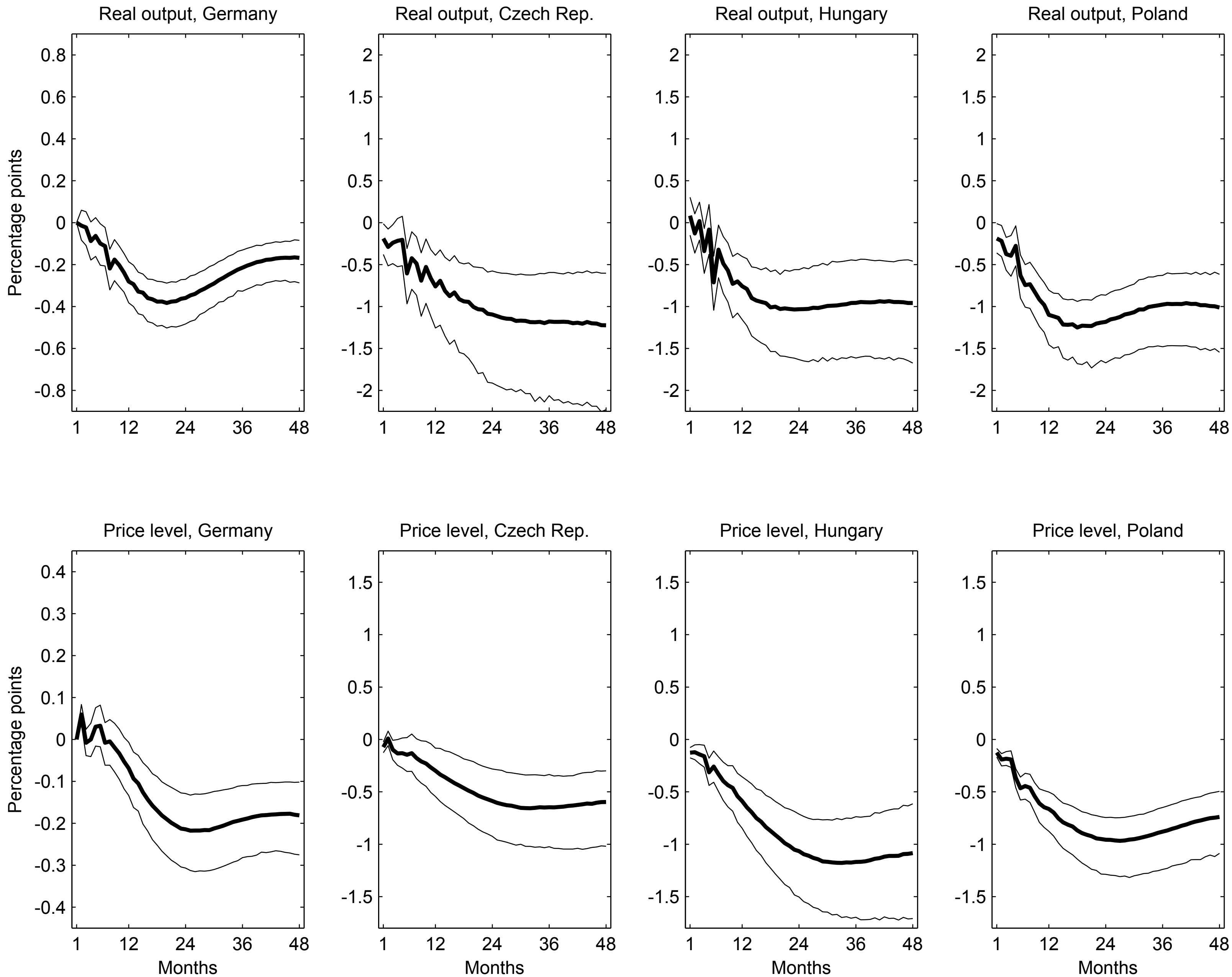
Note: Medians reported in each entry in the table.

Table 3: Variance decomposition (68% probability bands)
The contribution of **external shocks jointly** to forecast error variances

	Horizon (months)	**Czech Republic**	**Hungary**	**Poland**
Real output	*6*	11% - 26%	9% - 23%	15% - 29%
	12	15% - 35%	9% - 27%	21% - 39%
	24	25% - 53%	13% - 39%	29% - 53%
	48	32% - 72%	13% - 51%	35% - 69%
Price level	*6*	17% - 45%	13% - 36%	44% - 62%
	12	22% - 60%	35% - 66%	63% - 81%
	24	37% - 75%	62% - 87%	72% - 89%
	48	39% - 81%	76% - 94%	72% - 92%

Note: 68% Bayesian probability bands reported in each entry in the table.

Figure 2: Estimated impulse responses to a positive innovation in the short-term interest rate in Germany.



Each chart displays the impulse response (median in bold, with 68 percent probability bands), in percentage points and over 48 months, to a positive innovation in the short-term interest rate in Germany (one stardard deviation in size).

Table 4: Variance decomposition (medians)
The contribution of **euroarea interest rate shocks** to forecast error variances

	Horizon (months)	**Czech Republic**	**Hungary**	**Poland**
Real output	*6*	4.6%	4.5%	6.4%
	12	6.2%	5.2%	14.7%
	24	10.7%	8.9%	15.7%
	48	10.9%	7.8%	9.8%
Price level	*6*	2.8%	8.7%	27.9%
	12	9.1%	27.8%	36.6%
	24	26.0%	53.2%	59.8%
	48	25.7%	34.7%	35.5%

Note: Medians reported in each entry in the table.

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