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Call for entries: Olga Radzyner Award 2016

In 2000, the Oesterreichische Nationalbank (OeNB) established an award to commemorate Olga Radzyner, former Head of the OeNB's Foreign Research Division, who pioneered the OeNB's CESEE-related research activities. The award is bestowed on young economists for excellent research on topics of European economic integration and is conferred annually. In 2016, four applicants are eligible to receive a single payment of EUR 3,000 each from an annual total of EUR 12,000.

Submitted papers should cover European economic integration issues and be in English or German. They should not exceed 30 pages and should preferably be in the form of a working paper or scientific article. Authors shall submit their work before their 35th birthday and shall be citizens of any of the following countries: Albania, Belarus, Bosnia and Herzegovina, Bulgaria, Croatia, the Czech Republic, Estonia, FYR Macedonia, Hungary, Kosovo, Latvia, Lithuania, Moldova, Montenegro, Poland, Romania, Russia, Serbia, Slovakia, Slovenia or Ukraine. Previous winners of the Olga Radzyner Award, ESCB central bank employees as well as current and former OeNB staff are not eligible. In case of co-authored work, each of the co-authors has to fulfill all the entry criteria.

Authors shall send their submissions either by electronic mail to *eva.gehringer-wasserbauer@oenb.at* or by postal mail – with the envelope marked "Olga Radzyner Award 2016" – to the Oesterreichische Nationalbank, Foreign Research Division, POB 61, 1011 Vienna, Austria. Entries for the 2016 award should arrive by September 16, 2016, at the latest. Together with their submissions, applicants shall provide copies of their birth or citizenship certificates and a brief CV.

For detailed information, please visit the OeNB's website at *www.oenb.at/en/About-Us/Research-Promotion/Grants/Olga-Radzyner-Award.html* or contact Ms. Eva Gehringer-Wasserbauer in the OeNB's Foreign Research Division (write to *eva.gehringer-wasserbauer@oenb.at* or phone +43-1-40420-5226).

Call for applications: Visiting Research Program

The Oesterreichische Nationalbank (OeNB) invites applications from external researchers (EU or Swiss nationals) for participation in a Visiting Research Program established by the OeNB's Economic Analysis and Research Department. The purpose of this program is to enhance cooperation with members of academic and research institutions (preferably postdoc) who work in the fields of macroeconomics, international economics or financial economics and/or pursue a regional focus on Central, Eastern and Southeastern Europe.

The OeNB offers a stimulating and professional research environment in close proximity to the policymaking process. Visiting researchers are expected to collaborate with the OeNB's research staff on a prespecified topic and to participate actively in the department's internal seminars and other research activities. They will be provided with accommodation on demand and will, as a rule, have access to the department's computer resources. Their research output may be published in one of the department's publication outlets or as an OeNB Working Paper. Research visits should ideally last between three and six months, but timing is flexible.

Applications (in English) should include

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- a research proposal that motivates and clearly describes the envisaged research project,
- an indication of the period envisaged for the research visit, and
- information on previous scientific work.

Applications for 2016 should be e-mailed to *eva.gehringer-wasserbauer@oenb.at* by May 1, 2016.

Applicants will be notified of the jury's decision by mid-June. The following round of applications will close on November 1, 2016.

Studies

Modeling the evolution of monetary policy rules in CESEE

Martin Feldkircher,
Florian Huber,
Isabella Moder[1]

In this paper we use a novel econometric approach to estimate time-varying monetary policy rules for four inflation-targeting economies in Central, Eastern and Southeastern Europe (CESEE). Our results indicate that monetary policy in the Czech Republic, Hungary, Poland and Romania is strongly anchored to inflation stabilization, which implies that these economies follow a comparatively strict version of inflation targeting. By contrast, there is less evidence for output stabilization playing an important role in the conduct of monetary policy. Other factors that are of relevance in the monetary policy reaction function include the short-term interest rate in the euro area and – depending on the country under consideration – a measure of exchange rate movements. We find that the coefficients on domestic inflation expectations and euro area interest rates have declined since the mid-2000s, but that they still play an important role in central banks' reaction functions. This decline in the size of estimated coefficients may mirror an international environment characterized by loose and unconventional monetary policies that cannot be appropriately captured by euro area interest rates. It may also reflect contained global and domestic price growth.

JEL classification: E58, E52, C11, C22
Keywords: time-varying parameter model, monetary policy, Taylor rule

Since Taylor's (1993) seminal work on U.S. monetary policy, there has been a vastly growing literature that models central banks' reaction functions. These feedback rules often include a measure of inflation, real activity and other, more specific indicators of the economic environment (e.g. exchange rates, foreign interest rates and financial market conditions). Recent empirical estimates of the monetary policy reaction function take into account the fact that the conduct of monetary policy changes over time. Monetary policy may adapt due to changes in the economic environment or other, more abrupt events such as a switch from exchange rate targeting to inflation targeting. This was the case for most CESEE countries during the early 1990s. Another reason why policy rules change over time may be changes in the composition of monetary policy committees, which has recently been demonstrated by Jung and Kiss (2012). More generally, Orphanides (2004) has shown that most central banks' preferences have changed since the 1970s.

In this paper, we aim to close a gap in the literature by estimating a time-varying monetary policy rule for four CESEE economies which are currently pursuing inflation targeting, namely the Czech Republic, Hungary, Poland and Romania. Research on monetary policy rules in CESEE economies is scant, and the number of studies that estimate feedback rules that adapt to changes in the underlying macroeconomic conditions is even more limited. Also, existing research only covers the time span up to 2009, therefore we aim to shed light on how traditional monetary policy rules fit the most recent past, which has been characterized by major central banks, including the European Central Bank (ECB), pursuing loose monetary policies and low inflation in the euro area and the countries under consideration. To this end, we use a novel econometric approach in the spirit of Nakajima and West (2013) that allows estimating time-varying monetary policy

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rules for a rich dataset covering additional control variables identified in the existing literature. This approach does not only feature time-varying coefficients but a time-varying variable selection in a coherent statistical fashion. This means that we can find out whether interest rates were adjusted to particular domestic macroeconomic developments in one period, while this was not the case in another period.

Our results can be summarized as follows: First, for the period from mid-2000 to early 2015 we find that all CESEE economies under consideration respond strongly to changes in (expected) domestic inflation, while there is almost no evidence of policy rates adjusting to deviations in real activity. This result is in line with existing literature, which suggests that the CESEE countries under consideration seem to follow a comparatively strict version of inflation targeting. Second, we show that short-term interest rates in the euro area play an important role in domestic interest rate setting in the countries covered; however, the significance of this effect has declined recently. This may be explained by the fact that the three-month EURIBOR is not an optimal proxy to reflect the unconventional monetary policies that were adopted after the outbreak of the crisis. Third, we find that in Hungary and Romania, interest rates tend to adjust to movements in exchange rates, while results for the Czech Republic and Poland are less clear. Finally, for all four countries we find that interest rates are rather persistent. That is, changes in interest rates appear to happen gradually rather than abruptly.

The remainder of this paper is structured as follows: Section 1 provides a brief literature review on time-varying monetary policy rules on the one hand and monetary policy rules in CESEE on the other hand. In section 2 we introduce an augmented monetary policy rule in the spirit of Taylor (Taylor, 1993) and the econometric framework that allows for variable selection and time-varying parameters in a coherent way. Section 3 presents the empirical results and section 4 concludes.

1 Literature review

Changes in monetary policy reaction functions over time have been taken into account in several ways. Early work used sample splits at presumed break dates to estimate both samples independently (see, e.g., Clarida et al., 2000) or used dummy intercepts to account for different periods. While splitting samples is a simple way of dealing with time-varying coefficients, it has two disadvantages: First, the optimum point in time for a split has to be assumed, which is not always an obvious choice, and second, this implies that changes in coefficients are modeled as a change-point process, implying abrupt changes of the underlying structural coefficients. Another strand of literature makes use of Markov-switching regimes, typically distinguishing between a high- and a low-inflation regime (see, e.g., Assenmacher-Wesche, 2006). Compared to simple sample splitting, this method offers the advantage that no break date or threshold has to be predefined. The nature of Markov-switching regimes, however, does not allow for gradual adjustments in monetary policy but assumes an abrupt change of how monetary policy is conducted instead. While there may be occasions that would justify abrupt changes (such as a general switch from exchange rate to inflation targeting) a more realistic approach would allow for gradual changes. Such time-varying monetary policy rules can be estimated by drifting coefficients within a state-space model, brought

into the monetary policy rule literature by Boivin (2006), who examined Taylor rules estimates for the U.S.A. It enables the estimation of smooth changes without any prior assumption of sample breaks.

Since most empirical work on monetary policy rules has focused on the U.S.A. or other advanced economies, literature on monetary policy rules in CESEE is rather limited. Most studies (e.g. Mohanty and Klau, 2005; Paez-Farell, 2007; or Ghatak and Moore, 2011) use coefficients that are constant over the whole time span, and in some of these papers the time range starts already in the 1990s. The assumption of a linear feedback rule with coefficients that are constant over the sample period seems especially unrealistic for CESEE economies for two reasons: First, these countries underwent a major economic transition in the 1990s, and second, three out of four countries in our sample switched from an exchange rate-targeting framework to inflation targeting in the late 1990s or early 2000s. Petreski (2011) lists inferred and official switch dates from exchange rate to inflation targeting for the Czech Republic, Hungary[2] and Poland. Romania currently follows a managed floating exchange rate regime, in line with using inflation targets since 2005 as a nominal anchor for monetary policy. In addition, the economic environment changed considerably during the last decade, from economic boom phases in the mid-2000s to periods of severe contraction brought about by the global financial crisis and a period of recovery and loose international monetary policy. Given these facts, it seems unlikely that a linear feedback rule can appropriately characterize monetary policy for the countries considered in this study.

In the early literature addressing the non-linear nature of the monetary policy rule, usually sample splits or simple dummy intercepts are used to capture different regimes of monetary policy (see, e.g., Frömmel and Schobert, 2006; Yilmazkuday, 2009). More recently, Frömmel et al. (2011) and Petreski (2011) both used a Markov-switching approach to estimate monetary policy rules for CESEE economies. All these authors find that the countries under consideration in this study reacted strongly to inflation in the recent past, reflecting the move from exchange rate targeting to inflation targeting. With regard to the Czech Republic and Hungary, there is some evidence that contemporary monetary policy decisions are additionally driven by the exchange rate, the output gap and foreign interest rates (only for the Czech Republic). In contrast, for Poland and Romania most authors find that none of these additional variables significantly influence the interest-setting process.[3]

2 Data and empirical methodology

In this section we describe the methodology and data we use to evaluate Taylor rules for the Czech Republic, Hungary, Poland and Romania. Specifically, we estimate an augmented forward-looking Taylor rule where the coefficients are allowed to change over time.

[2] *Hungary kept an official exchange rate band of +/–15% against the euro until the beginning of 2008 (Frömmel et al., 2011).*

[3] *The impact of the leu's exchange rate on interest rate setting in Romania has not been investigated yet.*

2.1 Monetary policy rules

In his seminal paper, John Taylor (Taylor, 1993) proposed a rule that describes the reaction function of the U.S. Federal Reserve (Fed). In its simplest version, the Taylor rule postulates nominal interest rates as a function of inflation and real activity:

$$i_t^T = \overline{r} + \pi^* + \kappa\left(\pi_{t+12} - \pi^*\right) + \gamma(y_t - \overline{y}) \quad (1)$$

The rule states that interest rates (i_t^T) should rise if expected inflation (π_{t+12}) exceeds the inflation target set by the central bank (π^*) or if output (y_t) increases above its trend value $(\bar{y})$. As a by-product, the Taylor rule pins down the long-run neutral interest rate $\bar{r}$.[4] Note that this is a forward-looking version of the Taylor rule[5] since we include expected inflation as opposed to historical or contemporaneous inflation. These rules, as advocated in Clarida et al. (1998, 2000), provide a more realistic characterization of monetary policy as the interest-setting behavior of central banks is generally forward looking. Related to this debate is a discussion raised by Orphanides (2001) about the importance of using real-time data instead of data that are revised after a monetary policy decision has been made. While we do not have real-time data on output available for use, we construct the output gap based on an expanding-window estimation to better mimic the central bank's historical information set at each point in time compared to filtered data using all ex-post available data.

By estimating κ and γ it is possible to investigate whether the central bank stabilizes deviations of inflation from target and the output gap (Assenmacher-Wesche, 2006). More specifically, and to fulfill the Taylor principle, the coefficient on the deviation of inflation from target should be greater than 1 and the one attached to the output gap positive (see e.g., Woodford, 2001). Since nominal interest rates naturally respond one-for-one to increases in inflation (if the Fisher equation holds), a coefficient that is exactly unity would imply that the central bank would not sufficiently counteract inflation movements.

Since the Taylor rule was originally put forward to characterize monetary policy in the U.S.A., we extend the rule to account for structural features of small open economies. Albeit all countries considered in this study switched from exchange rate to inflation targeting prior to our sample period, several authors have argued more generally to control for exchange rates when estimating a Taylor rule (e.g. Clarida et al., 1998; Taylor, 2001), and especially so when looking at catching-up economies. In these economies, the exchange rate plays a more vital role than in advanced economies, where most domestic and foreign transactions are in local currency, markets are deeper, and the private sector is better equipped for absorbing exchange rate changes (Ostry et al., 2012). A number of studies have found that emerging market inflation targeters often (implicitly) include the exchange

[4] *Estimating the neutral rate via a Taylor rule has recently been shown to improve inflation forecasts (Horváth, 2009). See also Mehrotra and Slacík (2009), who show that actual money growth in relation to a McCallum rule provides information about inflation developments over a horizon of ten quarters for four CESEE economies.*

[5] *In the literature, the Taylor rule is often stated as* $i_t^T=\pi_t+2+0.5(\pi_t-2)+0.5(y_t-\bar{y})$. *This is simply a rearranged version of equation (1), in its backward-looking form using actual inflation along with Taylor's findings for* $\kappa=1.5$, $\gamma=0.5$. *Moreover Taylor assumed that the Fed effectively followed an inflation target of 2% between 1987 and 1993, and that the long-run real interest rate was also 2%.*

rate in their interest rate reaction function (see, e.g., Mohanty and Klau, 2005, and Aizenman et al., 2011).[6] In line with the majority of the literature, we include a contemporaneous measure and a lagged measure for exchange rate movements (er_t, er_{t-1}), where generally an increase in the measure signals a depreciation of the local currency. This comes with the benefit that, depending on the size and sign of the estimated coefficients, the reaction of the central bank can be classified by rules of thumb discussed in Taylor (2001) and Frömmel and Schobert (2006). For example, a country pursuing an explicit exchange rate target is expected to cut interest rates when faced with appreciation pressures. This would be reflected in significant positive coefficients on both the current and lagged values of the exchange rate. In other cases discussed in Frömmel and Schobert (2006), the central bank reacts only temporarily to movements in the exchange rate. With respect to CESEE countries, Frömmel et al. (2011) experiment with different indicators of exchange rate movements and conclude that augmenting the Taylor rule improves estimation results. In what follows we will experiment with levels and period-on-period growth rates of the national nominal exchange rate against the euro, the real effective exchange rate, and the exchange market pressure index (EMP), which captures depreciation pressures on the currency and deviations of the currency from its long-run trend. In addition we include the three-month EURIBOR as a foreign interest rate and the EONIA as a robustness check.[7] Including a foreign interest rate in the Taylor rule can be motivated from an uncovered interest rate parity perspective. The uncovered interest rate parity basically relates domestic interest rates to foreign interest rates and exchange rate expectations. While it is hard to empirically find evidence of the validity of the uncovered interest parity in the short run, Chinn and Meredith (2004) show for a set of G-7 countries that it holds in the long run. Consequently, and with respect to the countries considered in this study, foreign interest rates have been frequently included in monetary policy rules (see e.g., Arlt and Mandel, 2014; Horváth, 2009; Vašíček, 2010).

To complete the model, note that policy rates typically show a very persistent pattern over time since central banks tend to adjust interest rates gradually. The dynamics of adjustment of the actual level of the interest rate to its target is then given by:

$$i_t = \left(1 - \sum_{p=1}^{2} \rho_p\right) i_t^T + \sum_{p=1}^{2} \rho_p i_{t-p} \qquad (2)$$

That is, the central bank conducts interest rate smoothing by setting the actual rate as a linear combination of what is implied by the Taylor rule (first part in

[6] *Albeit in a textbook inflation-targeting setting, the exchange rate should only affect an inflation-targeting central bank's interest rate through its impact on expected inflation, a more pragmatic approach should admit the importance of the exchange rate in the case of catching-up economies and provide some leeway within the inflation target framework (Ostry et al., 2012).*

[7] *The results based on the EONIA are available from the authors upon request. Replacing the three-month EURIBOR by the EONIA yielded very similar results to those presented in section 3. In fact, correlations of estimated coefficients based on the estimations presented in section 3 and the robustness exercise including the EONIA are mostly close to 0.99. Only in Hungary, the coefficient on exchange rate developments shows a slightly smaller correlation of about 0.7.*

equation 2) and the historical interest rate (second part of equation 2). Our augmented Taylor rule that includes interest rate smoothing then becomes:

$$i_t = \left(1-\sum_{p=1}^{2}\rho_p\right)\left[\bar{r}+\pi^*+\kappa\left(\pi_{t+12}-\pi^*\right)+\gamma(y_t-\bar{y})+\psi er_t+\psi_1 er_{t-1}+\lambda eur_t\right]+\sum_{p=1}^{2}\rho_p i_{t-p} \quad (3)$$

where we have opted to include $p=2$ lags based on the marginal likelihood (approximated through the deviance information criterion).[8] Following Assenmacher-Wesche (2006), we assume that the central bank is able to control interest rates only up to a stochastic error u_t and subsume the long-run real interest rate and the inflation target into the constant $\alpha=\bar{r}-(\kappa-1)\pi^*$. We can then re-write equation (3) into

$$i_t = \left(1-\sum_{p=1}^{2}\rho_p\right)\left[\alpha+\kappa\pi_{t+12}+\gamma(y_t-\bar{y})+\psi er_t+\psi_1 er_{t-1}+\lambda eur_t\right]+\sum_{p=1}^{2}\rho_p i_{t-p}+u_t \quad (4)$$

Equation (4) is the model that is typically estimated in the literature. In this paper, however, we pursue a more flexible approach that allows estimated coefficients to vary over time. Accordingly the model becomes:

$$i_t = \left(1-\sum_{p=1}^{2}\rho_{t,p}\right)\left[\alpha_t+\kappa_t\pi_{t+12}+\gamma_t(y_t-\bar{y})+\psi_t er_t+\psi_{t,1} er_{t-1}+\lambda_t eur_t\right]+\sum_{p=1}^{2}\rho_{t,p} i_{t-p}+u_t \quad (5)$$

2.2 Econometric framework

This section introduces the econometric setting for the empirical analysis that follows. Let us assume that a time series $\{z_t,\ t=1,\ldots,\ T\}$ is described by the observation equation

$$\mathrm{z}_t = \mathrm{x}'_t\mathrm{b}_t + \mathrm{u}_t, \quad (6)$$

where $x_t=(x_{1t},\ldots,\ x_{Kt})'$ denotes a K-dimensional vector of possible explanatory variables measured in time t and $b_t=(b_{1t},\ldots,\ b_{Kt})'$ is a $K\times 1$ vector of dynamic regression coefficients. Furthermore, let u_t be a normally distributed white noise error with zero mean and variance σ^2.

Following Nakajima and West (2013), we assume that the elements of b_t, b_{jt} $(j=1,\ \ldots,\ K)$ are related to a latent stochastic process β_{jt} as follows

$$\mathrm{b}_{jt} = \beta_{jt}\mathrm{s}_{jt},\ \ \mathrm{s}_{jt} = \mathrm{I}\left(\left|\beta_{jt}\right| > \mathrm{d}_j\right). \quad (7)$$

Here $I(|\beta_{jt}|>d_j)$ denotes the indicator function which equals unity if the latent parameter β_{jt} exceeds a threshold $d_j\in\mathbb{R}$ to be estimated from the data. This implies that if β_{jt} is small, $s_{jt}=0$ and thus $b_{jt}=0$. Since β_{jt} evolves over time, this implies that b_{jt} could be non-zero for some points in time whereas for other periods it could equal zero. Thus $s_{jt}=0$ implies that there is no regression relationship between z_t and x_{jt} in time. The assumption that b_{jt} arises as a thresholded variant of β_{jt} provides a flexible and parsimonious means of modeling dynamic relationships and account-

[8] *Results based on p=3 lags generally yielded a worse fit and are available from the authors upon request.*

ing for model uncertainty, effectively capturing the notion that some variables may be relevant in some periods and less relevant in other periods.

To complete the model we also have to impose a law of motion for $\beta_t=(\beta_{1t},\ldots,\beta_{Kt})'$ where we assume for simplicity that it follows a random walk process

$$\beta_t = \beta_{t-1} + e_t. \tag{8}$$

Here, e_t is a vector white noise process with zero mean and a $K\times K$ dimensional variance-covariance matrix $V=diag(\vartheta_1, \ldots, \vartheta_K)$. Equation (8) is typically called a state equation for β_t. In general, it would be straightforward to assume that equation (8) is a stationary process. However, since the data are typically not very informative on the actual persistence of the latent states, and the length of our data is rather limited we take the simpler route and estimate equation (8).

The model outlined above belongs to the general class of state-space models and provides a large degree of flexibility in terms of modeling. For instance, our model is capable of unveiling changes in the underlying structural behavior of the time series analyzed. This proves to be of prime importance when used to investigate the behavior of a central bank. It is worth noting that our model nests a plethora of simpler models. For instance, if e_t equals the zero vector for all t then $\beta_t=\beta_{t-1}$ and equation (6) collapses to a standard linear regression model.

The model is estimated in a Bayesian fashion. This implies that suitable prior distributions have to be imposed on all parameters, which are described in more detail in the next section. Estimation is done using the Markov chain Monte Carlo (MCMC) algorithm described in Nakajima and West (2013). This implies that conditional on all other parameters the latent states are sampled by means of a Metropolis Hastings (MH) step. Likewise, we adopt a direct MH step to simulate the thresholds controlling the inclusion/exclusion of a given variable. For the remaining steps simple conditional posterior distributions exist, which makes Gibbs sampling feasible.

2.3 Prior distributions and implementation

We take a Bayesian stance to estimation and inference. More specifically, this implies that we have to specify suitable prior distributions on the parameters of the model. In the present application we use the following prior setup. We specify a standard inverted gamma prior on σ^2,

$$\sigma^2 \sim IG(\underline{a}_0, \underline{a}_1), \tag{9}$$

with $\underline{a}_0=0.01$, $\underline{a}_1=0.01$ being hyperparameters. The specific values chosen render the prior effectively uninformative. Furthermore, following Griffin and Brown (2010) and Bitto and Frühwirth-Schnatter (2014), we impose the following hierarchical priors on the elements of the initial state β_0,

$$\beta_{0j} \sim N\left(0, \tau_j^2\right),\ \tau_j^2 \sim G\left(a_\tau, \frac{a_\tau \xi_i}{2}\right),\ \xi_i \sim G(\underline{b}_0, \underline{b}_1). \tag{10}$$

Let $a_\tau=0.2, \underline{b}_0=0.01$ and $\underline{b}_1=0.01$ denote scalar hyperparameters chosen by the researcher. Finally, we impose a prior setup similar to the one described in the previous equation on the square root of the variances in V,

$$\sqrt{\vartheta_j} \sim N\left(0,\phi_j^2\right),\ \phi_j^2 \sim G\left(a_\phi, \frac{a_\phi \zeta_i}{2}\right),\ \zeta_i \sim G\left(\underline{c}_0, \underline{c}_1\right). \tag{11}$$

Similar to the prior on β_{0j} we set $a_\phi=0.2$, $\underline{c}_0=0.01$ and $\underline{c}_1=0.01$. This prior controls the degree of time-variation in the parameters and is thus crucial for the present application. Recently, researchers searched for priors that allow for significant variation in the autoregressive parameters while keeping the model stable, which is also the approach we follow in this paper.

The traditional prior setup used in this model relies on a pre-sample that is used to inform the corresponding prior distributions. In addition, hyperparameters are typically set such that the model is stable. Typically, this can be regarded as one of the main limitations of this modeling approach. However, in this paper we pursue a hierarchical approach that limits the problems regarding prior elicitation.

2.4 Data

For each country, our sample contains 131 monthly observations spanning the period from 7/2004 to 5/2015. The reason we do no extend the sample span to cover the period prior to 2004 is mainly due to the lack of reliable data on inflation expectations. Moreover, and as discussed above, by the start of the time period considered, all countries had already switched to inflation targeting. Hence, we use the longest sample available for estimating a forward-looking Taylor rule and cover various economic regimes and different phases of movements in exchange rates, prices and, ultimately, monetary policy. Data on policy rates are from Bloomberg and provided in percent per annum. As a measure for real activity, we select industrial production (including the construction sector, seasonally and working day adjusted, volume index). We calculate the output gap using a standard Hodrick-Prescott (HP) filter with $\lambda=129,600$. As mentioned earlier, the HP filter is estimated by using an expanding window to ensure that we take only information into account that was available to the central bank at that particular point in time. Inflation expectations for all countries considered are taken from the Hungarian central bank's quarterly inflation report and reflect one year-ahead inflation expectations of households. We use various measures to capture interest rate sensitivity to exchange rate movements. To this end, we collected data on the real (CPI-based) effective exchange rate from the Bank for International Settlements. Moreover we include national currencies' nominal bilateral exchange rates vis-à-vis the euro and the exchange market pressure index as advocated in Aizenman and Pasricha (2012) and Feldkircher et al. (2014). The EMP is defined as $emp_t=\left(\frac{e_t-e_{t-1}}{e_{t-1}}-\frac{ir_t-ir_{t-1}}{ir_{t-1}}\right)$, with e_t denoting the local nominal exchange rate per EUR 1 and ir_t standing for international reserves (minus gold) in U.S. dollars. Finally, we also include deviations of national currencies' nominal exchange rates vis-à-vis the euro from their long-run trend, estimated by the HP filter (fx_gap_t). For all measures related to exchange rates, an increase reflects a depreciation (pressure) of the currency and the related coefficient in the Taylor rule is expected to be positive. Lastly, data on the three-month EURIBOR, the average interest rate at

which European banks are willing to lend money to each other over the three-month horizon, and the EONIA, the interbank rate at which banks lend overnight, are retrieved directly from the ECB's statistical data warehouse.

3 Empirical results

In this section we provide evidence whether monetary policy in the Czech Republic, Hungary, Poland and Romania can be adequately described by time-varying augmented Taylor rules. The main results are summarized in charts 1 to 4. In each chart, the first plot in the top panel on the left-hand side shows the domestic policy rate. The subsequent charts show, on the left-hand scale, the policy rate, coefficients attached to inflation expectations, the output gap, a measure for the exchange rate and the three-month EURIBOR. Moreover, we plot the coefficients associated with the lagged interest and exchange rates. In general, we present results for the specification that yielded the best in-sample fit and coefficients with the expected sign. The long-run transform of the coefficients is achieved by multiplying the respective coefficients with $\left[\frac{1}{1-\sum^{2}\widehat{\rho_{t-p}}}\right]$. In the charts, we further show the associated posterior inclusion probabilities (PIPs) plotted on the right-hand scale. They indicate how often a particular variable has been included in the Taylor rule model compared to how often the coefficient has been pushed toward zero. Variables with PIPs greater than 0.5 – as indicated by exceeding the dashed blue line in the charts – are considered as important regressors (Barbieri and Berger, 2004). Finally, the vertical line marks the collapse of the investment bank Lehman Brothers in September 2008, which is generally seen to mark the outbreak of the global financial crisis.

Chart 1 shows the results for the *Czech Republic*. Looking at the main components of the Taylor rule first, we see that our findings reveal both positive coefficients on inflation expectations and deviations of output from its long-run trend. The coefficient on deviations from trend output, however, receives only little support in the data as indicated by small posterior inclusion probabilities throughout the sample period. This finding is in line with Horváth (2009), who examines a broad range of Taylor rules, including backward- and forward-looking rules and rarely finds evidence for output stabilization for the Czech Republic. Vašíček (2010) argues that output gaps are typically estimated with a large noise component, which can be even more pronounced for emerging economies. The coefficient on inflation expectations is small but positive and hovers around 0.2 at the beginning of the sample. Hence our estimation results differ from those of Ghatak and Moore (2011) and Mehrotra and Sánchez-Fung (2011), who employ a linear regression framework and report insignificant (and negative) responses of interest rates to inflation. With the outbreak of the crisis, the coefficient on inflation starts to decrease strongly. Note that the size of the estimated coefficients does not directly reflect the central bank's preferences regarding output relative to inflation stabilization. This is so because the coefficients and the weight the central bank puts on inflation versus output stabilization are related in a non-linear fashion (Svensson, 1998).[9] To recover central banks' preferences, we would have to pin down an optimal monetary policy rule in connection with a structural model for

[9] *Assenmacher-Wesche (2006) estimates non-linear Taylor rules for three industrialized countries and, in a second step, maps the reduced form estimates into a loss function of the central bank.*

the economy under consideration first. This is in contrast to the modeling approach taken in this paper, which relies on a standard but ad hoc formulation of the monetary policy reaction function. See, among others, Assenmacher-Wesche (2006), Favero and Rovelli (2003) and Castelnuovo and Surico (2003) for structural estimations of central banks' preferences. The small coefficient attached to inflation (expectations) might be seen as a violation of the Taylor principle which ensures inflation stabilization only if the respective coefficient exceeds unity. However, as noted in Linnemann and Schabert (2006), the interpretation for small open economies is less straightforward since here additional variables typically appear as important ingredients in empirical Taylor rules. Estimates of sensitivity to inflation are close to those of Petreski (2011), who uses a Markov switching approach to estimate the Taylor rule, and Vašíček (2010) using a linear model. Next, we examine the reaction of the central bank to deviations of the exchange rate from its long-run trend. A positive coefficient would indicate that the central bank increases the policy rate in case the exchange rate depreciates. Looking at the coefficients, we see that the contemporaneous coefficient is negative. However, the coefficient on lagged deviations of the exchange rate from its long-run trend is positive and the sum of both coefficients seems balanced. In general, this finding indicates that the exchange rate is not included in the central bank's reaction function – or, put differently, that the interest rate is not used to stabilize the exchange rate. This does not, however, imply that the Czech central bank does not pay close attention to exchange rate movements. In fact, it decided in autumn 2012 to use the exchange rate as a monetary policy instrument and commenced foreign exchange interventions a year later with the aim of letting the koruna not appreciate well above CZK/EUR 27. That is, while the policy rate is the main monetary instrument to achieve price stability, foreign exchange market operations are used as an additional central bank instrument operating alongside the monetary policy/Taylor rule. See Ostry et al. (2012) and Benes et al. (2013) for analytical approaches to model exchange rate interventions for inflation-targeting economies. Our empirical results on the exchange rate also comply with the findings of Petreski (2011), who reports a positive coefficient for the period when the Czech central bank pursued an exchange rate target – not covered in our sample – whereas during the period of inflation targeting the coefficient on the exchange rate is not significant. Next, and in line with Arlt and Mandel (2014), Horváth (2009) and Vašíček (2010), we find that the policy rate of the Czech central bank reacts significantly and strongly to the three-month EURIBOR. More specifically, a 100 basis point increase in the three-month EURIBOR would suggest a 65 basis point increase in domestic interest rates at the beginning of the sample. In the most recent sample period this effect has somewhat abated, which might be related to the fact that the EURIBOR does not adequately reflect the monetary policy stance in times when monetary policy is characterized by unconventional measures. Lastly, our results point to a significant degree of interest rate smoothing. The sum of the two lags of the policy rate is close to unity. This implies that even in the case of strong and significant coefficients on some of the variables discussed above, the Czech central bank does not necessarily change its policy rate immediately. Whereas the other coefficients have shown marked variation over time, coefficients on lags of the policy rates have been pretty stable over the period considered. This implies that while the relative importance of output, inflation and

exchange rate stabilization has varied over time, there has been no general regime shift of the Czech central bank's interest rate-setting behavior such as altering rates more frequently and aggressively during our sample period.

In chart 2 we depict the results for *Hungary*. Both estimated coefficients on inflation and output are positive but the latter is not significant throughout the sample period. In contrast to the results for the Czech Republic, the coefficient on inflation is close to unity at the beginning of the sample, after which sensitivity to inflation declines to about 0.7 in the most recent period of the sample, which is close to estimates provided in Frömmel et al. (2011) and slightly higher compared to the findings of Petreski (2011). The lack of significance in sensitivity of domestic interest rates to the output gap is in line with the existing literature (see e.g., Ghatak and Moore, 2011; Petreski, 2011; Jung and Kiss, 2012; Vašíček, 2010). Both coefficients on the exchange rate are positive, complying with the results of Yilmazkuday (2009) and Frömmel et. al (2011). Although the exchange rate band of +/–15% against the euro was officially not abandoned until February 2008, our results reveal that the importance of the exchange rate already decreased in the run-up to this date. The coefficient then rose slightly again in the subsequent years but lost ground from 2012 – a period in which the Hungarian authorities started to implement several measures to reduce household foreign currency debt and, consequently, decoupled movements in the exchange rate from national financial stability considerations. Note, however, that posterior inclusion probabilities throughout the sample period are above but close to the 0.5 threshold, indicating positive but weak support in the data. Next, we find that the Hungarian short-term interest rate adjusted also to the short-term interest rate in the euro area. This effect declined strongly from the beginning of the sample period to late 2007. From that period on, the effect started to rise again (to about 0.85). This result might indicate that – due to the abandonment of the exchange rate band – the euro area interest rate replaced the exchange rate in influencing monetary policy decision making. Compared to the Czech Republic, adjustments to the interest rate in the euro area are in general more pronounced in Hungary. Lastly, we also find a high degree of persistence in interest-setting behavior.

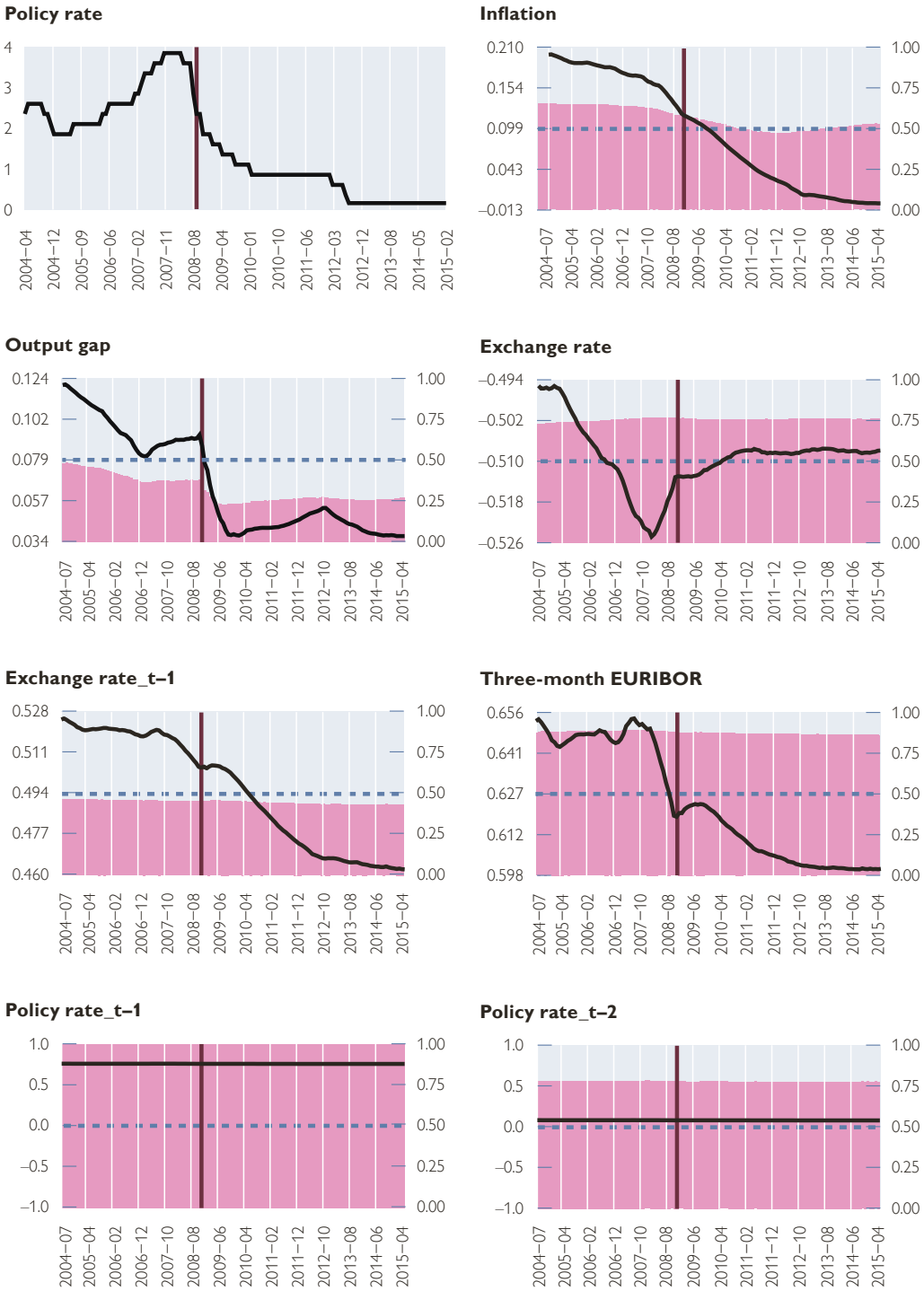
Chart 3 summarizes the results for *Poland*. Two observations stand out: First, among all the countries under investigation, the-long term inflation coefficient is the highest exceeding unity throughout the sample period. Moreover, the coefficient receives strong posterior support in the data, with PIPs of close to 0.8 for all time periods considered. In other words, Polish monetary policy is characterized by a strong commitment to inflation stabilization (Frömmel and Schobert, 2006). Second, the estimated Taylor rule for Poland is the only one that yields a positive and significant coefficient on the deviation of real activity from its long-term trend. This evidence is rather weak, however, since posterior inclusion probabilities exceed the 0.5 threshold only marginally. Next, we investigate whether interest rates adjust to movements in the exchange rate. As in the Taylor rule for the Czech Republic, the best fit is achieved by including a measure that indicates deviations from the złoty's long-run trend. Both coefficients on current and lagged exchange rate deviations are estimated with opposite signs, while they sum up to approximately zero, which indicates a zero net effect. Non-significant effects of interest rate adjustment to exchange rate movements are also reported in Petreski (2011). Interest rate adjustments in Poland are also influenced by short-term inter-

est rates in the euro area. However, the estimated coefficient is much smaller relative to that of the other countries covered in this study. Also, the coefficient decreased somewhat, from about 0.4 at the beginning of the sample period to 0.36 in the most recent period of the sample. In line with the results for the Czech Republic and Hungary, we find evidence for interest rate smoothing with lagged coefficients on the policy rate summing up to close to unity.

Lastly, we investigate estimates for the augmented Taylor rule in *Romania* (chart 4). The variables affecting interest setting are inflation expectations, period-on-period change in the nominal exchange rate and short-term interest rates in the euro area. Interest rates adjust significantly to movements in inflation expectations. In line with the results for the other countries covered in this study, the coefficient attached to inflation was more pronounced at the beginning of the sample period (about 0.6) and less so during the most recent part of the sample period (about 0.4). There is no evidence in the data that monetary policy reacted to deviations of output from the long-run trend, as indicted by PIPs far below the 0.5 threshold; this complies with the results of Frömmel et al. (2011), Ghatak and Moore (2011) and Vašíček (2010). Both coefficients on exchange rate changes are positive, indicating that monetary policy tightened when the Romanian leu weakened against the euro, but of low significance. The estimated policy rule for Romania also features short-term interest rates in the euro area. Compared to the other countries in this study, the estimates for Romania are by far the most pronounced, ranging from about 1.5 at the beginning of the sample period to about 0.8 in the most recent period. The decline in the estimated coefficient is similar to the results for the Czech Republic, Poland and, partially, Hungary. Finally, we find evidence for interest rate smoothing. Compared to the other countries, the degree of smoothing is somewhat smaller and interest rate setting seems less persistent.

Chart 1

Czech Republic

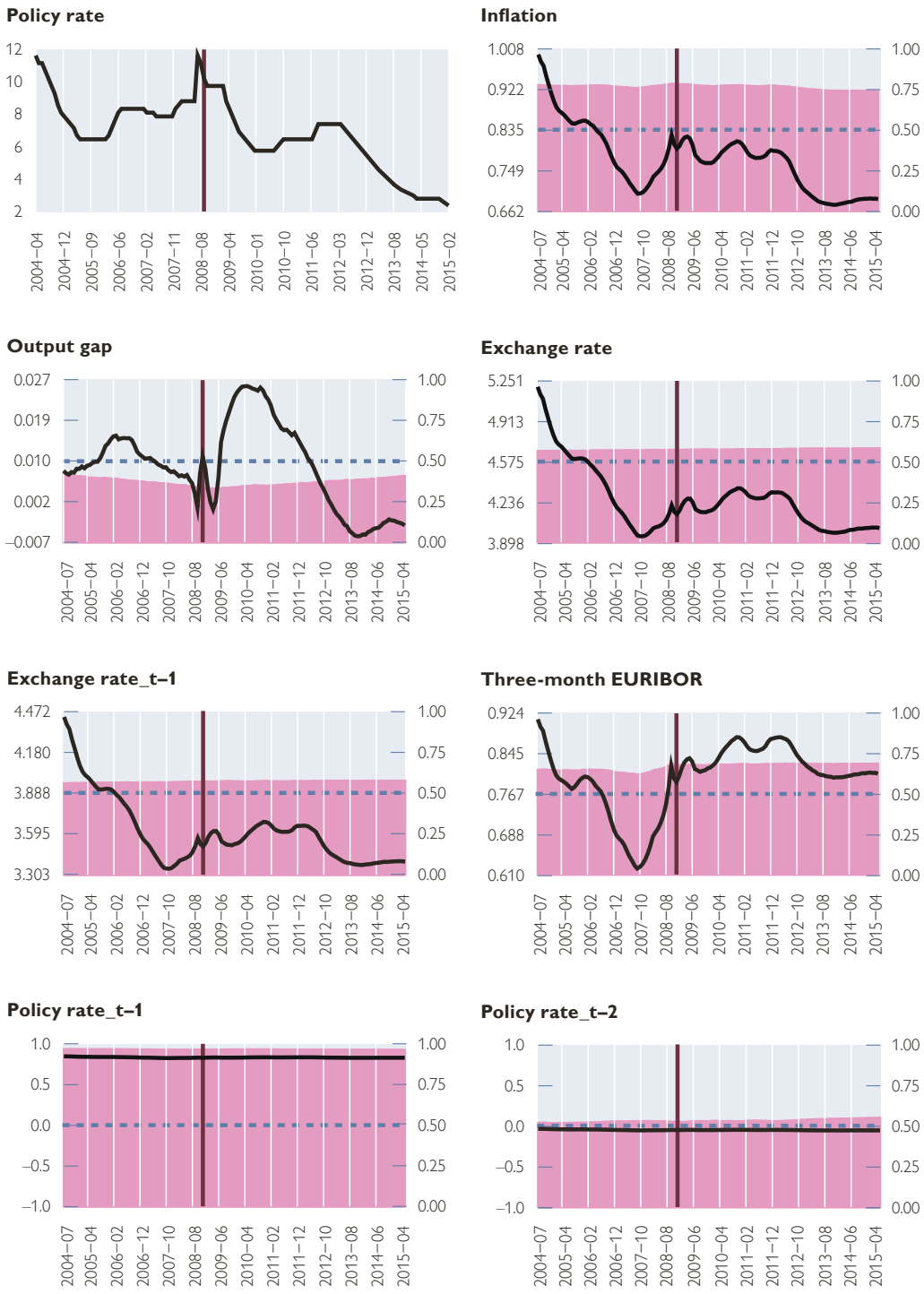


Source: Authors' calculations.

Note: The first plot depicts the domestic policy rate, whereas the remaining plots show the long-run coefficients (black solid line, left-hand scale) of inflation expectations (12 months ahead), the output gap, deviations of the exchange rate from its long-run trend, the three-month EURIBOR and short-run coefficients for two lags of the policy rate. The pink area refers to the time-varying posterior inclusion probability of the respective variable (right-hand scale). A variable should be included in the model if it receives a higher posterior inclusion probability than 0.5 (dashed blue line). The vertical line marks the outbreak of the global financial crisis.

Chart 2

Hungary

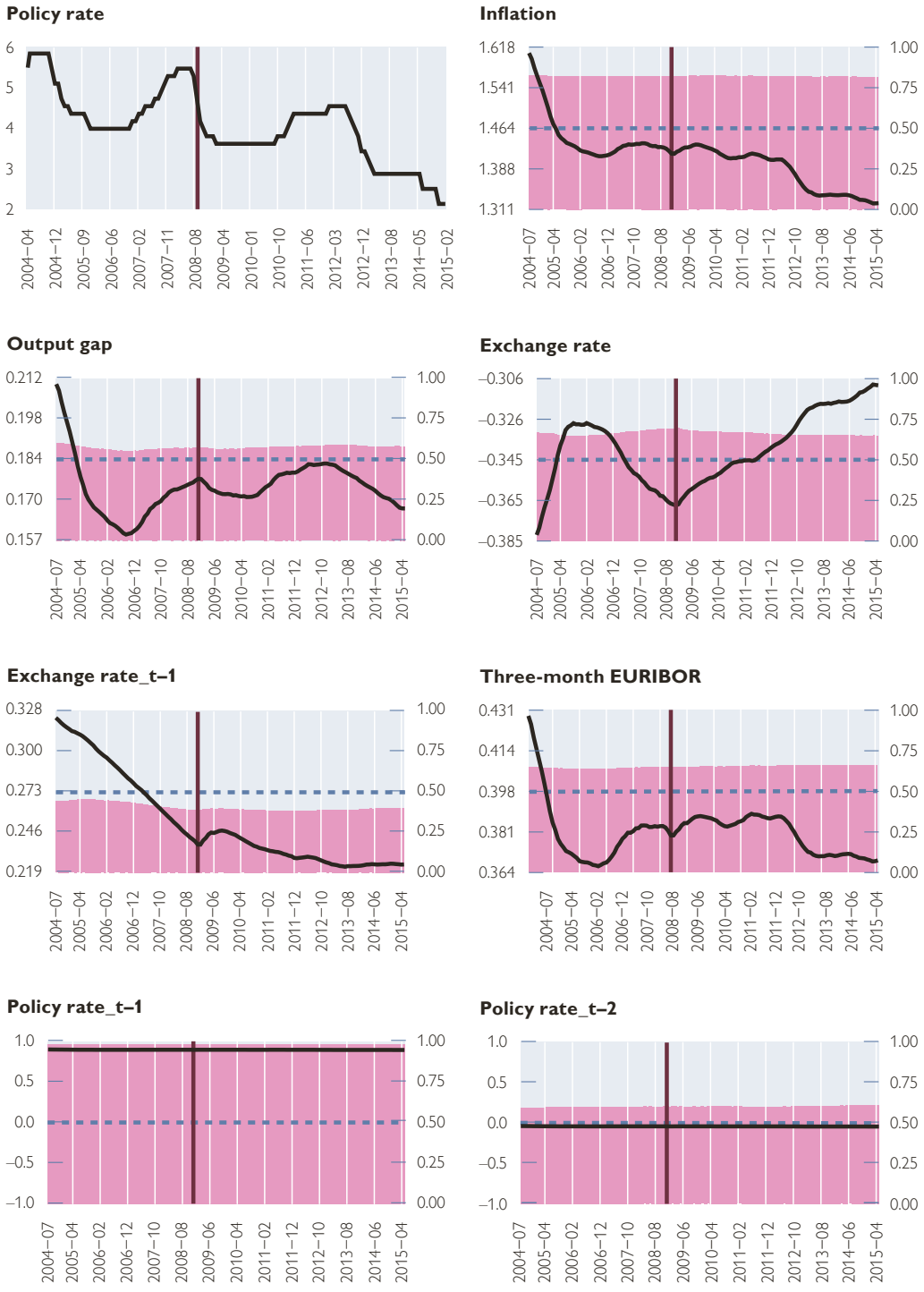


Source: Authors' calculations.

Note: The first plot depicts the domestic policy rate, whereas the remaining plots show the long-run coefficients (black solid line, left-hand scale) of inflation expectations (12 months ahead), the output gap, deviations of the exchange rate from its long-run trend, the three-month EURIBOR and short-run coefficients for two lags of the policy rate. The pink area refers to the time-varying posterior inclusion probability of the respective variable (right-hand scale). A variable should be included in the model if it receives a higher posterior inclusion probability than 0.5 (dashed blue line). The vertical line marks the outbreak of the global financial crisis.

Chart 3

Poland

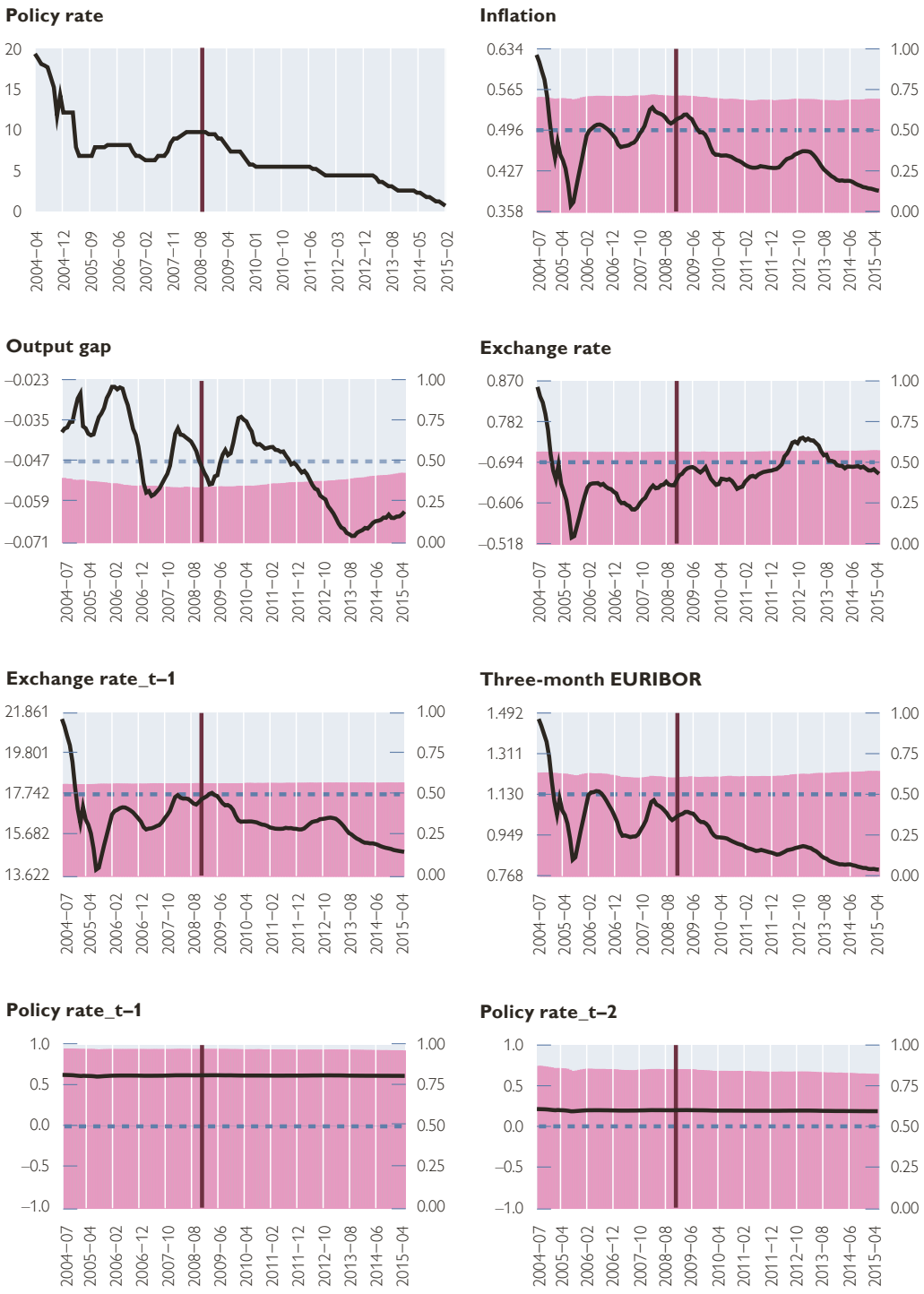


Source: Authors' calculations.

Note: The first plot depicts the domestic policy rate, whereas the remaining plots show the long-run coefficients (black solid line, left-hand scale) of inflation expectations (12 months ahead), the output gap, deviations of the exchange rate from its long-run trend, the three-month EURIBOR and short-run coefficients for two lags of the policy rate. The pink area refers to the time-varying posterior inclusion probability of the respective variable (right-hand scale). A variable should be included in the model if it receives a higher posterior inclusion probability than 0.5 (dashed blue line). The vertical line marks the outbreak of the global financial crisis.

Chart 4

Romania



Source: Authors' calculations.

Note: The first plot depicts the domestic policy rate, whereas the remaining plots show the long-run coefficients (black solid line, left-hand scale) of inflation expectations (12 months ahead), the output gap, nominal exchange rate vis-à-vis the euro (period-on-period growth), the three-month EURIBOR and short-run coefficients for two lags of the policy rate. The pink area refers to the time-varying posterior inclusion probability of the respective variable (right-hand scale). A variable should be included in the model if it receives a higher posterior inclusion probability than 0.5 (dashed blue line). The vertical line marks the outbreak of the global financial crisis.

4 Conclusions

In this study we estimate monetary policy rules for four of the largest CESEE economies, namely the Czech Republic, Hungary, Poland and Romania. During the last decade, these economies underwent different business cycle regimes, ranging from prolonged boom phases in the mid-2000s[10] to periods of contraction during the global financial crisis and a recovery phase characterized by a loose monetary policy stance of major central banks, including the ECB. Hence, we estimate monetary policy reaction functions that make it possible to take into account those changes by drawing on a novel econometric framework that features time-varying coefficients and variable selection in a coherent fashion.

Our results can be summarized as follows: First, we find evidence for a significant response of the policy rate to domestic expected inflation. Among all countries considered, the responses were most pronounced in Poland and less so in the Czech Republic. This does not necessarily imply that monetary policy in Poland has been more strongly anchored to price stability than in the Czech Republic. To clarify this issue, it would be necessary to assess how the estimated coefficients of the Taylor rule depend on the preference parameters of the central bank and on the structural parameters of the economy (Hayo and Hofmann, 2006).[11] Second, we find that output stabilization seems to play a minor role for interest setting in the countries considered, with the exception of Poland (Vašíček, 2010). Our results hence comply with the official communication of these countries' central banks, which seem to follow a comparatively strict version of inflation targeting. Third, we find that in Hungary and Romania, interest rates tend to adjust to movements in exchange rates. That is, if the national currencies face depreciation pressures, central banks will respond by increasing the policy rate. For the Czech Republic and Poland, the results are less clear. Fourth, we find that euro area short-term interest rates feature prominently in the estimated Taylor rules for all countries considered in this study. This does not necessarily imply that these countries' central banks aim to stabilize the domestic exchange rate via anchoring monetary policy to euro area short-term rates – which would contradict the weak evidence for exchange rate stabilization discussed above – since additional factors besides foreign interest rates determine exchange rates (e.g. country risk premia). Rather, this result shows how strongly these countries are integrated with the euro area and that changes in the euro area's monetary policy stance are likely to feed directly into the domestic economies (Babecká-Kucharčuková et al., 2014). Finally, looking at the time variation of the estimated monetary policy rules, we find that interest rates' adjustment to inflation has decreased in all CESEE countries under consideration, mirroring the low interest rate and low inflation environment in the most recent part of our sample period. That is, during the boom years in the mid-2000s interest rates adjusted more strongly to inflation than during the global financial crisis and its aftermath. Also, the importance of the three-month EURIBOR in the estimated Taylor rules decreased over time for all countries (except Hungary), but to different extents. Note that even against the

[10] *Hungary, which experienced a higher degree of volatility of economic growth, represents the only exception from this pattern.*

[11] *In other words, estimating central banks' preferences requires modeling the economy under consideration jointly with the Taylor rule – and even in this case some restrictive assumptions have to be made to recover the preferences of the monetary policymaker. (see, e.g., Castelnuovo and Surico, 2003, or Favero and Rovelli, 2003).*

backdrop of a declining trend of coefficients attached to inflation expectations and the EURIBOR, the data still show pronounced evidence of both variables being important components of monetary policy rules. The decline in the estimates for the three-month EURIBOR might be explained by the fact that this rate may be a good proxy for overall monetary policy conditions in the euro area during normal periods, but less so during times when monetary conditions are driven by unconventional measures (Babecká-Kucharčuková et al., 2014). Future work might look more closely into the direct effects of euro area quantitative easing on interest rate setting and, more generally, monetary policy in CESEE economies.[12]

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[12] *One approach that has resurfaced recently is capturing unconventional monetary policy via a "shadow interest rate" as proposed e.g. in Lombardi and Zhu (2014) and Whu and Xia (2015) that resembles a conventional policy rate during "normal" times and can turn negative during times the central bank provides additional stimulus by implementing nonstandard measures. However, as yet, there is no theoretical underpinning for embedding shadow rates into monetary policy rules.*

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A monetary policy rule for Russia, or is it rules?

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Riikka Nuutilainen[1]

We estimate several monetary policy rules for Russia for the period 2003–2015. We find that the traditional Taylor rule describes the conduct of monetary policy in Russia reasonably well, whether coefficients are restricted to being the same or allowed to change over the sample period. We find that the Bank of Russia often overshot its inflation target and that extensive overshooting is associated with large depreciations of the ruble, testifying to the importance of the exchange rate in the conduct of monetary policy in Russia.

JEL classification: E31, E43, E52, P33
Keywords: monetary policy rule, Taylor rule, McCallum rule, Russia, inflation

Introduction and motivation

In this paper, we estimate several monetary policy rules for Russia. Such estimations are standard for most OECD countries and are often used in policy debates. However, estimates of monetary policy rules for emerging markets are much less common, and this is where we make a contribution.

As one of the G-20 countries, Russia is one of the largest emerging markets, and its financial markets are relatively advanced.[2] Russia remains to a large extent dependent on exports of hydrocarbons, with approximately two-thirds of its export revenue coming from sales of crude oil, oil products and natural gas. This feature of the Russian economy obviously has implications also for the conduct of monetary policy.

Using data from 2003 to 2015, we estimate several different specifications for monetary policy rules in Russia. For the whole period, we find that an augmented Taylor rule seems to depict the data reasonably well. Russian monetary authorities seem to focus on maintaining the stability of both inflation and the output gap and respond to changes in these aggregates by adjusting the interest rate. However, the results are somewhat sensitive to whether the exchange rate or both the exchange rate and the oil price are included in the empirical specification. In contrast to earlier studies, however, the McCallum rule does not seem to fit the data very well, indicating a change in the conduct of monetary policy in Russia.

The Taylor rule seems to be in congruence with the data also when we allow the coefficients of monetary policy rule to change over time. It is noteworthy that at the very end of our sample, the Bank of Russia (CBR) seems to have placed much greater weight on output stabilization. This is perhaps understandable given the fall in output in 2014 and 2015. Also, one must note that the Bank of Russia moved officially into full-fledged inflation targeting only in 2015 and that exchange rate targeting was officially abandoned in November 2014. In 2014, the Bank of Russia stated that "Starting from 2015, the monetary policy will be conducted under the inflation targeting regime" (Bank of Russia, 2014). At that time the central bank expected to achieve the medium-term inflation target of 4% in 2017. While

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[2] *For example, by 2014, domestic credit provided by the financial sector had reached 52% of GDP, higher than e.g. in Romania (38%), an EU country, or Kazakhstan (37%), a large former Soviet republic.*

Russia's exchange rate targeting was already very flexible, this change could affect our results at the very end of the sample.

We must highlight the role of the exchange rate and the oil price in our estimates. Including them in the estimated policy functions sometimes leads to counterintuitive results, which may be explained by some of the idiosyncratic features of Russian economic policy.

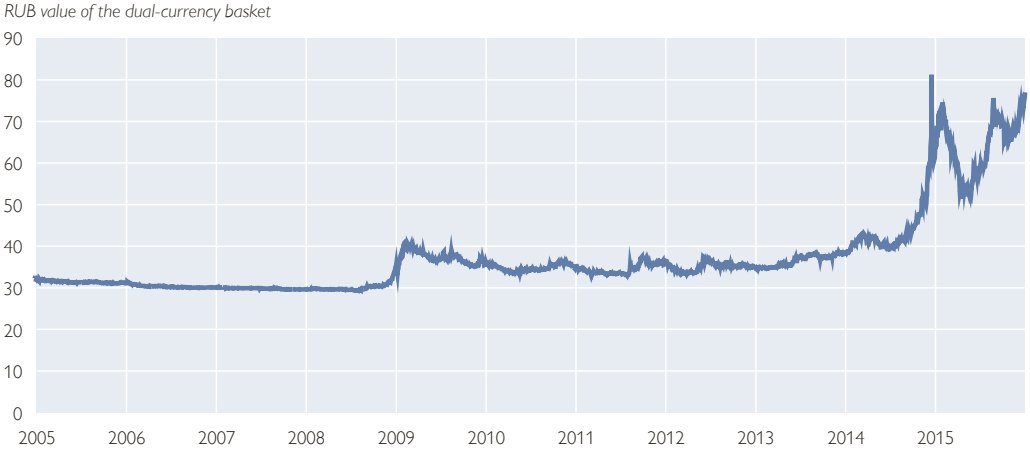
The study is structured as follows. In the first section, we describe the conduct of monetary policy in Russia during our sample period and present a short literature survey. The second section introduces the monetary policy rules to be estimated as well as the data. The third section discusses the empirical estimates, and the fourth section concludes.

1 Monetary policy rules for Russia

Our data sample runs from 2003 to 2015. During this time, the Bank of Russia had several goals for its policy, although the whole period was marked by a gradual shift toward more full-fledged inflation targeting, which was officially introduced from the beginning of 2015. At the same time, the central bank explicitly pursued exchange rate stability as one of its key targets for almost the whole sample period. The Bank of Russia gave up the exchange rate target only in November 2014, although it announced then that it would stand ready to intervene in the foreign exchange markets to dampen undue volatility. However, it should be noted that the Bank of Russia had continuously widened the allowed fluctuation band around the central parity of its exchange rate basket, which consists of the U.S. dollar and the euro to reflect both Russia's foreign trade orientation and the dollar's traditionally large role in the Russian economy. Moreover, the targeted exchange rate was also allowed to change to reflect underlying market pressures, especially after 2008, as is evidenced in chart 1.[3]

Chart 1

CBR dual-currency basket

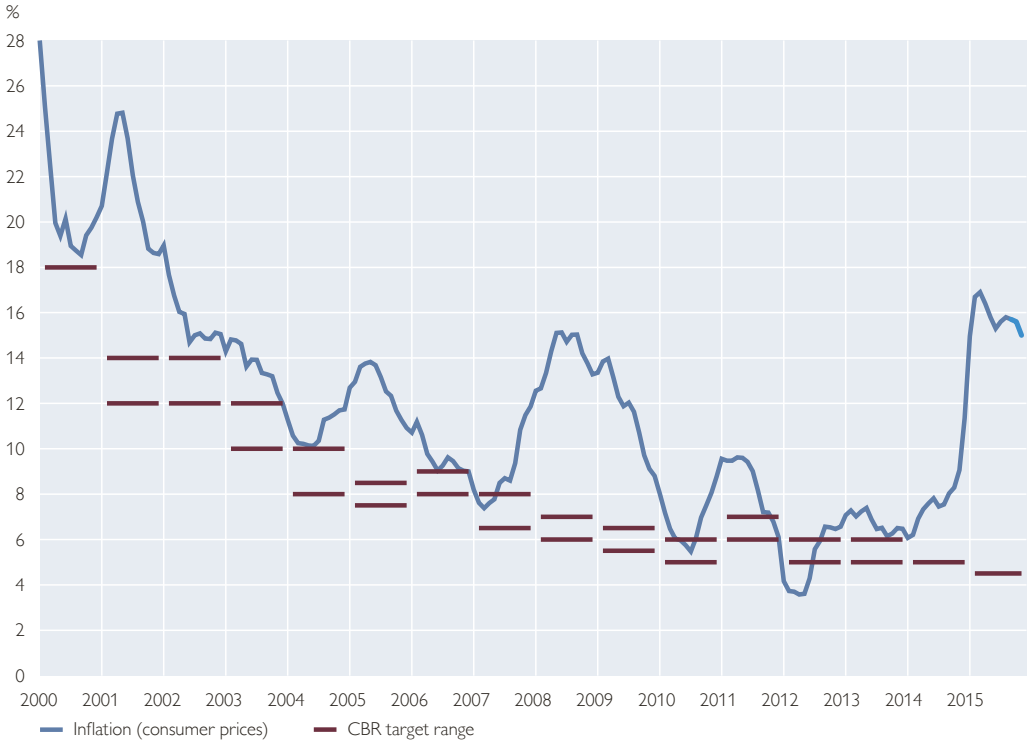


Source: Bank of Russia (CBR).

[3] *The value of the Bank of Russia's dual-currency basket is calculated as a weighted sum of the ruble values in U.S. dollars and euro at official exchange rates. The weights varied in 2005–2007 after the dual-currency basket was adopted. An initial weight of USD 0.90/EUR 0.10 was used in February 2005, and the weight of the euro was gradually increased to the current USD 0.55/EUR 0.45 in February 2007. This weighting remains in place to this day, but the currency basket lost its relevance for exchange rate policy in November 2014.*

Chart 2

Inflation and annual inflation targets



Source: IMF, CBR, Rosstat.

The Bank of Russia first stated price stability as its primary policy objective in its 2007 monetary policy guidelines (Bank of Russia, 2006). This can be seen as the starting point for the gradual move to inflation targeting in Russia.

Chart 2 shows inflation targets (or target ranges) set by the Bank of Russia as well as actual inflation from 2000 to 2015. Note that especially during earlier periods, it was sometimes difficult to discern inflation targets from inflation forecasts, although the ranges were called inflation targets in the Bank of Russia's annual monetary policy guidelines. One can see that actual inflation overshot inflation targets on several occasions, and that the greatest deviations from inflation targets happened in the aftermath of large currency depreciations, for example in 2009 and 2015. This empirical regularity can be used to justify inclusion of an exchange rate variable in the empirical estimates of Russia's monetary policy rules, which is further corroborated by the official role of the exchange rate basket.

While empirical estimates of different monetary policy rules are relatively common in advanced OECD countries, similar exercises for emerging market countries are still quite rare. Moreover, there are only a handful of published papers on monetary policy rules in Russia, and their data samples usually end more than a decade before our data. Esanov et al. (2005) estimate several monetary policy rules for Russia for the period starting in 1993 or 1994 and ending in 2002. For a large part of this data sample, Russia had a fixed exchange rate target. The authors find that the McCallum rule with the monetary base as a target fits the

data best. In their estimation, the U.S. dollar exchange rate is also used as a control variable. The results are plausible in the sense that monetary aggregates were explicitly used as intermediate targets during much of this period. However, there is a structural break in the data in 1995, when the ruble was officially pegged to the U.S. dollar. This reminds us of the importance of the exchange rate for the conduct of Russian monetary policy.

Vdovichenko and Voronina (2006) estimate monetary policy rules for a period starting only after the crisis of 1998, but their sample is very short, from 2000 to 2003. They also find that the McCallum rule with the monetary base seems to reflect the underlying data reasonably well, but only when the exchange rate is included as well.

Drobyshevsky et al. (2008) look at the conduct of monetary policy in Russia between 1999 and 2007. They find that commercial banks' correspondent accounts in the central bank seem to be the instrument of choice for monetary policy. This would speak for a variant of the McCallum rule for Russia.

One may also note that a somewhat stable link between monetary aggregates and other economic variables, i.e. the money demand function, is needed for the McCallum rule to be a viable strategy for a central bank to follow. For Russia, e.g. Korhonen and Mehrotra (2010) find such a stable money demand function, but again, the exchange rate needs to be included in the estimated empirical relationship.

2 Methodology and data

We estimate two types of monetary policy reaction functions to evaluate the Bank of Russia's behavior in 2003–2015. We utilize the literature on monetary policy rules to formulate the reaction functions. For a timely capture of the recent policy changes, we use monthly data in the estimations. This section introduces the policy rules estimated and the data used in the empirical analysis. Data and their original sources are listed in table A1 in the annex.

The estimated interest rate rule is a version of the famous rule proposed by John Taylor (1993), according to which a central bank reacts to output gaps and deviations of inflation from a target rate. Following Taylor (2001), we select an open economy version of the rule, accounting also for exchange rate developments, because of the strong emphasis on exchange rate stabilization in the monetary policy of the Bank of Russia. In addition, oil prices strongly impact the behavior of output, inflation and the exchange rate in Russia. It is reasonable to assume that the Bank of Russia takes oil prices directly into account in monetary policy decisions. Therefore, oil prices are added to the policy rule as one of the macroeconomic variables to which the central bank may directly react when setting its policy.

Taylor (1993, 2001) assumes that the central bank reacts to deviations of output from a potential level. Determining potential output in practice, however, is very difficult even for developed countries that have long time series, let alone emerging economies like Russia that display structural changes. Following Orphanides and Williams (2007), we estimate the so-called "difference rule" that considers changes in output growth from long-run trend growth. There is much less controversy in determining the trend growth rate than potential output for an economy.

Following the empirical literature, policy smoothing is added to the estimated rules to increase their empirical fit.[4]

We estimate the Taylor interest rate rule of the form:

$$i_t = \alpha_0 + \alpha_1(\pi - \pi^*)_{t-1} + \alpha_2 \widehat{\Delta y}_{t-1} + \alpha_3 \widehat{reer}_{t-1} + \alpha_4 \widehat{oil}_{t-1} + \alpha_5 \widehat{oil}_{t-2} + \alpha_6 i_{t-1} + \varepsilon_t. \quad (1)$$

In the empirical estimations, we use the Bank of Russia key policy rate (the one-week repo credit rate) as the policy interest rate from February 2011 onward, when the central bank adopted this instrument and started to publish the data. We select the refinancing rate as the policy interest rate prior to that date.[5] The inflation deviation term $(\pi-\pi^*)_{t-1}$ is determined as the year-on-year growth of consumer prices over the annual CPI growth target determined by the central bank for the year.[6] We use the inflation target observed at the beginning of the year in question, as this should be the most relevant e.g. for formulating expectations for monetary policy. On some occasions when it became obvious that original target could not be reached, the Bank of Russia changed the target toward the end of the year. We do not take these changes into account.

Output growth deviation $\widehat{\Delta y}_{t-1}$ is calculated by removing the Hodrick-Prescott (HP) filtered trend from the estimated monthly GDP year-on-year growth series published by the Russian Ministry of Economic Development.[7] Similarly, the exchange rate deviation $\widehat{reer}_{t-1}$ and oil price deviation $\widehat{oil}_{t-1}$ are calculated by removing the HP trend from the real effective exchange rate (REER) index and the index for Urals oil prices, respectively. In equation (1), α_0 is a constant term and ε_t stands for the estimation error. Parameters α_1 to α_5 are the estimated policy reaction coefficients and α_6 measures the strength of policy smoothing. For the policy to be countercyclical, we should observe that $\alpha_1>0$, $\alpha_2>0$, $\alpha_3<0$ and $\alpha_4,\alpha_5>0$.

In addition to the interest rate rule, we also estimate a money supply rule introduced by McCallum (1988). The McCallum rule is defined in nominal terms. McCallum (1988, 2000) suggests that the central bank should react to nominal output growth deviation from the target rate. This way, the policy would not be biased in the short run to the errors arising when separating the realized nominal output growth into real growth and inflation. We follow McCallum (1988, 2000) and use base money growth as the policy instrument, because it is the monetary aggregate over which the central bank has full control. The estimated McCallum

[4] *The majority of the empirical studies include policy smoothing in the estimated policy rules. Examples include Clarida et al. (1998), who estimate such rules for large developed countries, Mehrotra and Sánchez-Fung (2011) for 20 emerging countries as well as Vdovichenko and Voronina (2006) and Esanov et al. (2005) for Russia.*

[5] *The level of the policy rate is shifted up to match the refinancing rate in February 2011, so that only true policy changes affect the interest rate variable (see the upper left panel in chart A4).*

[6] *For a robustness check, an HP-filtered inflation deviation series is also considered. There is not much difference, except at the very end of our sample, between using the official inflation target or HP filtering to determine the trend inflation rate (see middle left panel in chart A4).*

[7] *Hodrick-Prescott filtering is a standard method for removing trend levels and calculating the output gap. However, it has an obvious shortcoming of unreliability at the beginning and end of the data sample. In calculating the de-trended series, we used data starting in January 1999, wherever available. Our HP-filtered data may still suffer from the endpoint problem at the end of our sample. However, in an earlier version of the paper, we used data only up to February 2015, and results were almost identical. This makes us confident that our results do not depend on the very last observations. We use the year-on-year GDP growth data, as month-on-month data are not available. By using only the cycle component of year-on-year growth, the output growth deviation variable is better able to capture the sudden changes than the year-on-year growth rate itself.*

rule is also formulated to take into account possible policy reactions to exchange rate and oil price changes as well as to account for policy smoothing.

The McCallum rule estimated is of the form:

$$\Delta bm_t = \beta_0 + \beta_1 \widehat{\Delta x}_{t-1} + \beta_2 \widehat{neer}_{t-1} + \beta_3 \widehat{oil}_{t-1} + \beta_4 \widehat{oil}_{t-2} + \beta_5 \Delta bm_{t-1} + \varepsilon_t \qquad (2)$$

The nominal base money growth Δbm_t is the year-on-year change in the monetary aggregate M0. Fortunately, the Russian Ministry of Finance publishes a monthly GDP estimate in rubles.[8] We use this series to calculate the year-on-year nominal GDP growth rate and use the HP filter to get the nominal output growth deviation $\widehat{\Delta x}_{t-1}$. The exchange rate gap and oil price gap are calculated similarly to (1), but using the nominal effective exchange rate (NEER) index. Again, β_0 is a constant term, β_1 to β_4 measure the strength of policy reactions in base money to the macroeconomic variables and β_5 measures policy inertia. The error term ε_t captures elements of random behavior that might be present at time t and potential omitted variables and specification errors. Increases in base money indicate policy easing. Therefore, the signs in the countercyclical policy reaction are the opposite of those in the Taylor rule: $\beta_1<0$, $\beta_2>0$, and $\beta_3, \beta_4>0$.

The estimated policy rules are formulated to retain the rules' operationality. Policy is assumed to react to the macroeconomic variables prevailing in the previous period, which thus are available at time t. Traditionally, Taylor rules have been estimated with realized data, which is also one of the strengths of the approach, as one does not need to take a stand on expectation formation. This is also the approach we follow here. Also, HP filtering is performed using data available until the time of estimation. To adequately account for policy reactions to oil prices, the second lag of the oil price deviation also needs to be added to the policy rules.

Chart A4 in the annex depicts the data series used in the empirical estimations. All variables used in the estimations can be considered to be stationary in levels.[9] Descriptive statistics and unit root test statistics of the variables are presented in table A2. Last, correlations between the variables are presented in table A3.

3 Estimation results

The policy reaction functions are empirically estimated using the generalized method of moments (GMM) estimator. The use of the GMM is fairly standard in estimating policy reaction functions with inertia and possible measurement errors in the variables. Estimation results are presented in table 1 and table 2. Our data sample spans January 2002 to November 2015. In addition, we have the December 2015 values for the policy variables, which enables us to estimate the monetary policy rules until end-2015. The McCallum rule is estimated using data from

[8] *The monthly GDP estimate can deviate a few percentage points from Rosstat's final ruble GDP value in annual terms. But the monthly estimate by the Ministry of Finance is available to the central bank for its policy decisions much sooner than Rosstat's official quarterly GDP.*

[9] *An augmented Dickey-Fuller (ADF) unit root test cannot reject the null hypothesis of a unit root in the inflation deviation variable, but the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test does not reject the null for stationarity either. Moreover, in the case of the reference policy rate, although the null hypothesis of unit root is rejected by the ADF test, the null for stationarity cannot be rejected by the KPSS test. All other variables are stationary at least at the 5% level of significance on the basis of both the ADF and the KPSS tests.*

January 2003.[10] The Taylor rule is estimated from 2004 onward. Prior to 2004, the Taylor rule residuals are not well-behaved and suffer from non-normality and autocorrelation. As a robustness check, the policy reactions are also estimated for a more recent time period starting from 2007, when inflation targeting was initiated as the main policy goal of the Bank of Russia. The previous literature on estimating monetary policy rules for Russia (Esanov et al., 2005; Vdovichenko and Voronina, 2006) does not take into account central bank policy reactions to oil prices. To maintain comparability with these earlier results, the Taylor and McCallum rules are estimated also without the oil price variable.

3.1 Time-invariant policy rules

The estimated policy reactions of the Taylor rule (equation 1) are presented in table 1. The policy reactions are generally in line with the theoretical assumptions showing a stabilizing policy in terms of reactions to both inflation and output growth deviations. The reactions are also statistically significant. The estimation results differ little whether we use the time period starting in 2004 or the more recent period from 2007 onward.

The policy reactions to exchange rate developments and oil prices are harder to interpret, as these two variables are largely interrelated. The interest rate reactions to the exchange rate and oil prices are statistically significant, but the signs of the estimated reactions are the opposite of those assumed beforehand. In the policy rules literature, policy easing is assumed to follow exchange rate appreciation. Here, we find the opposite. Also, an increase in oil prices is assumed to lead to policy tightening, as it boosts future output growth and increases inflation. When disregarding the effect of oil prices, the estimated policy reactions to inflation and output gaps are smaller and statistically less significant in some instances. In addition, the reaction to exchange rate changes remains positive. Our results may be explained by the fact that a rise in the oil price also leads to exchange rate appreciation. We might not be able to completely disentangle these two effects in our estimation, which may lead to the observation that exchange rate appreciation is followed by monetary policy tightening, even if the oil price increase is the original cause of the appreciation.[11]

[10] *Data availability partly limits the selection of the estimation period, as the base money aggregate is available only from 2003 onward.*

[11] *In our dataset, the correlation between the REER gap and the oil price gap is 0.55 (and the correlation between the NEER gap and the oil price gap is 0.58). Also, the lagged oil price gap correlates strongly with the REER and NEER gaps (see table A3 in the annex). As a robustness check, we have also estimated the Taylor rule using the nominal effective exchange rate (NEER). The results are largely similar to the ones using the REER. In our estimation period, there is no considerable difference between the REER and NEER gap series (see lower left panel in chart A4).*

Estimated McCallum rule policy reactions are presented in table 2. Russian monetary policy reacts countercyclically to nominal output growth deviation. The nominal output gap reaction parameter is statistically significant only in the period starting from 2003 and when including oil prices in the estimated policy rule. After 2007, the reactions to nominal output as well as all other macroeconomic variables except the exchange rate become statistically insignificant. Therefore, the McCallum rule does not describe the Bank of Russia's policy in the more recent period.

Oil prices are important among the variables the Bank of Russia considers in making its policy decisions. In terms of the strength and significance of the estimated policy reactions to output growth and inflation deviations, policy rules that do account for oil prices perform better than those that do not. Policy inertia as measured by the autoregressive (AR) lag coefficient is also less pronounced in the Taylor rule when oil prices are added to the estimated equation. In general, policy smoothing behavior is strong in the estimated rules. This is common in the empirical estimations of policy rules, especially when using higher-frequency monthly data (see, for example, Clarida et al., 1998, Mehrotra and Sánchez-Fung, 2011, Vdovichenko and Voronina, 2006, as well as Esanov et al., 2005).

Table 1

Taylor rule estimation results

	Taylor rule 01/2004–12/2015		Taylor rule 01/2007–12/2015	
	i_t	i_t	i_t	i_t
c	0.399***	0.249**	0.590***	0.260*
	(0.140)	(0.124)	(0.172)	(0.144)
$(\pi-\pi^*)_{t-1}$	0.026**	0.018**	0.029**	0.013
	(0.010)	(0.009)	(0.013)	(0.012)
$\widehat{\Delta y}_{t-1}$	0.023**	0.020**	0.028***	0.024***
	(0.009)	(0.008)	(0.008)	(0.008)
$\widehat{reer}_{t-1}$	0.025***	0.022***	0.023***	0.019***
	(0.007)	(0.006)	(0.007)	(0.005)
$\widehat{oil}_{t-1}$	–3.120***		–3.033***	
	(0.586)		(0.518)	
$\widehat{oil}_{t-2}$	2.709***		2.577***	
	(0.672)		(0.551)	
i_{t-1}	0.949***	0.965***	0.929***	0.965***
	(0.015)	(0.013)	(0.021)	(0.018)
SSR	78.90	79.51	73.01	73.90
J-stat.	14.13	9.10	11.57	7.74
	(0.70)	(0.91)	(0.87)	(0.96)

Source: Authors' calculations.

*Note: The table presents GMM estimates. Standard errors are given in parentheses. ***, ** and * denote the 1%, 5% and 10% level of significance, respectively. The instrument list includes a constant and second, third, fourth, fifth and sixth lags of the variables. The instrument lag selection is based on the autocorrelation behavior of the dependent variable. Standard errors and covariances are computed using a Newey-West weighting matrix. SSR=sum of squared residuals.*

Table 2

McCallum rule estimation results

	McCallum rule 05/2003–12/2015		McCallum rule 01/2007–12/2015
	Δbm_t	Δbm_t	Δbm_t
c	0.683	1.049*	–0.252
	(0.572)	(0.564)	(0.540)
$\widehat{\Delta x}_{t-1}$	–0.285*	–0.192	–0.181
	(0.163)	(0.141)	(0.133)
$\widehat{reer}_{t-1}$	0.037	0.161*	–0.195*
	(0.136)	(0.090)	(0.107)
$\widehat{oil}_{t-1}$	–8.931		8.750
	(16.073)		(10.937)
$\widehat{oil}_{t-2}$	17.801		9.595
	(13.595)		(11.370)
Δbm_{t-1}	0.949***	0.927***	1.002***
	(0.031)	(0.028)	(0.045)
SSR	4387.6	4331.8	3231.8
J-stat.	4.54	4.17	7.23
	(0.92)	(0.90)	(0.70)

Source: Authors' calculations.

*Note: The table presents GMM estimates. Standard errors are given in parentheses. ***, ** and * denote the 1%, 5% and 10% level of significance, respectively. The instrument list includes a constant and second, third, fourth and fifth lags of the variables. The instrument lag selection is based on the autocorrelation behavior of the dependent variable. Standard errors and covariances are computed using a Newey-West weighting matrix. SSR=sum of squared residuals.*

3.2 Time-varying policy reactions

In this section, the behavior of monetary policy reaction is allowed to vary over time. To this aim, the monetary policy rules (equations 1 and 2) are estimated in a rolling window. We select an eight-year rolling window (96 observations)[12] and use data starting from January 2002. The estimation window is moved one observation forward at each step, and the policy reactions parameters are re-estimated. Proceeding this way, the policy reactions are estimated for 73 subsamples. Hodrick-Prescott filtering is performed at each step prior to the estimation for time t, using data available until $t–1$, to ensure that the variables in the estimation do not depend on future releases of the data.

The Taylor rule rolling parameter estimates and their 90% confidence bounds are displayed in chart A1 in the annex. The confidence bounds are calculated based on the estimated standard errors computed using a Newey-West weighting matrix. The interest rate exhibits a statistically significant reaction to inflation deviation from the first subsample (January 2002–December 2009) to the subsample ending in February 2015. Reactions to the output gap are statistically significant for the entire estimation period, except during a short period at the beginning of the estimation sample. At end-2014, the Bank of Russia started to react very strongly to the output gap, and at the same time, reactions to inflation became less significant. Interest rate reactions to inflation remain insignificant until the end of our estimation sample, and the AR policy-smoothing parameter also shows values larger than one. One may of course interpret the strong reaction to the output gap as signifying that the Bank of Russia is reacting to the output gap, but eventually, the reaction will also have an effect on inflation via Phillips curve. During the turbulent times at end-2014 and in 2015, however, the Taylor rule does not seem to fit the Russian data as well as before end-2014.

McCallum time-varying estimates are depicted in chart A2. Base money does not seem to react to oil prices; therefore, chart A3 presents the time-varying parameters for the McCallum rule without oil prices. The McCallum rule fits the Russian data until the 2004–2011 sample. Reactions in base money to nominal output growth deviation are negative and statistically significant at the 10% level. Reactions to the exchange rate are also significant and positive, as is assumed in the literature. The time-varying estimation confirms our earlier finding. After around 2012, the McCallum rule performs very poorly in describing Russian monetary policy.

To illustrate the difference between the Bank of Russia's inflation and output objectives, chart 3 shows the Taylor rule time-varying long run responses.[13] The chart presents long-run parameters only for the subperiods they are statistically significantly different from zero. In addition, it omits the most recent periods during which the Taylor rule does not describe monetary policy in Russia well and during which the value of the policy-smoothing AR parameter in the rolling Taylor rule estimation is equal to or above one. The long-run estimated coefficients for

[12] *The estimation window selection is subject to the tradeoff between estimation accuracy with a large enough sample size and the ability of the rolling estimates to detect policy changes occurring in the most recent data in a timely fashion. A seven-year (84 observations) and a nine-year (108 observations) window is also considered, and the results remain largely robust to the window selection.*

[13] *The long-run response parameters are calculated as* $\beta^{LR} = \frac{\beta}{1-\rho}$, *where* β *is the estimated short-run reaction and* ρ *is the estimated policy smoothing parameter.*

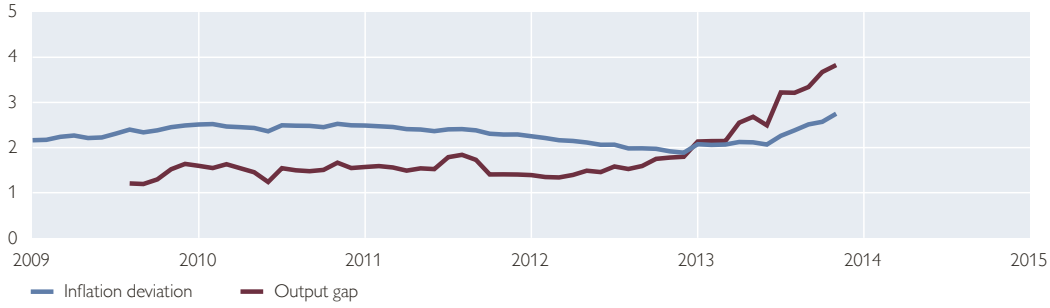
inflation are not very far from 1.5, the value Taylor (1993) selected to describe U.S. monetary policy from the late 1980s to the early 1990s. In 2010–2014, the long-run inflation coefficient is larger than one, thus fulfilling the "Taylor principle." The Bank of Russia's interest rate policy seems to place a relatively large weight on output stabilization, however, as the long-run output gap reaction parameter is higher than the 0.5 suggested by Taylor (1993). Interestingly, chart 3 indicates a change in the tradeoff between the two policy objectives. Prior to 2014, monetary policy was more concerned with price stability, but since then, output growth stability has become relatively more important.

The long-run policy response to nominal output in the McCallum rule without oil prices is displayed in chart 4. Again, the response coefficient is depicted only for the subperiods in which the short-run parameter is statistically significant (parameter ρ is always significant). Quantitatively, the strength of the policy response is stronger than the value of 0.5 suggested by McCallum (2000) for the growth-type policy rules.

Chart 3

Taylor rule long-run response over time

Estimated coefficients of long-run response parameters

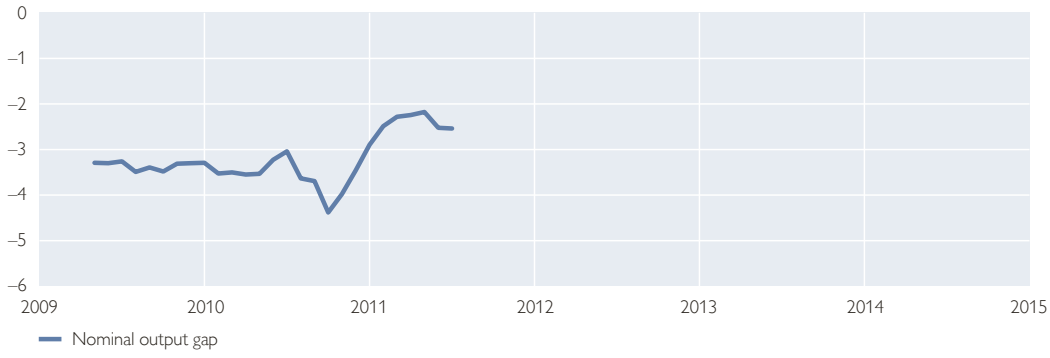


Source: Authors' calculations.

Chart 4

McCallum rule long-run response over time

Estimated coefficients of long-run response parameters



Source: Authors' calculations.

4 Concluding remarks

We have estimated different monetary policy rules for Russia for the period 2003–2015. As no recent papers have undertaken a similar exercise, our contribution is able to illustrate the challenges Russian policymakers faced during both calm and very turbulent periods. We can see that the traditional Taylor rule seems to describe monetary policy in Russia reasonably well, even though the Bank of Russia has moved to full-fledged inflation targeting only recently. Even if exchange rate stability has also been important, the Bank of Russia has stabilized inflation in a manner consistent with the experience of many other central banks in the world. Moreover, monetary authorities have clearly also tried to dampen output fluctuations, and the weight of the output gap in the central bank's objective function seems to have increased during the very turbulent period of 2014 and 2015. For this reason, central bank interest rate policy seems to have stopped reacting statistically significantly to inflation in 2015, so that the traditional Taylor rule does not provide as good a description of Russian monetary policy as prior to 2015.

It is noteworthy that earlier papers on Russia's monetary policy rules emphasized the role of monetary aggregates. The low level of development in Russia's financial markets was often cited as the reason for this monetary policy choice. Our results with more recent data suggest that Russian monetary policy has changed, and the observed move toward inflation targeting also tells us that Russia's financial markets have become more mature.

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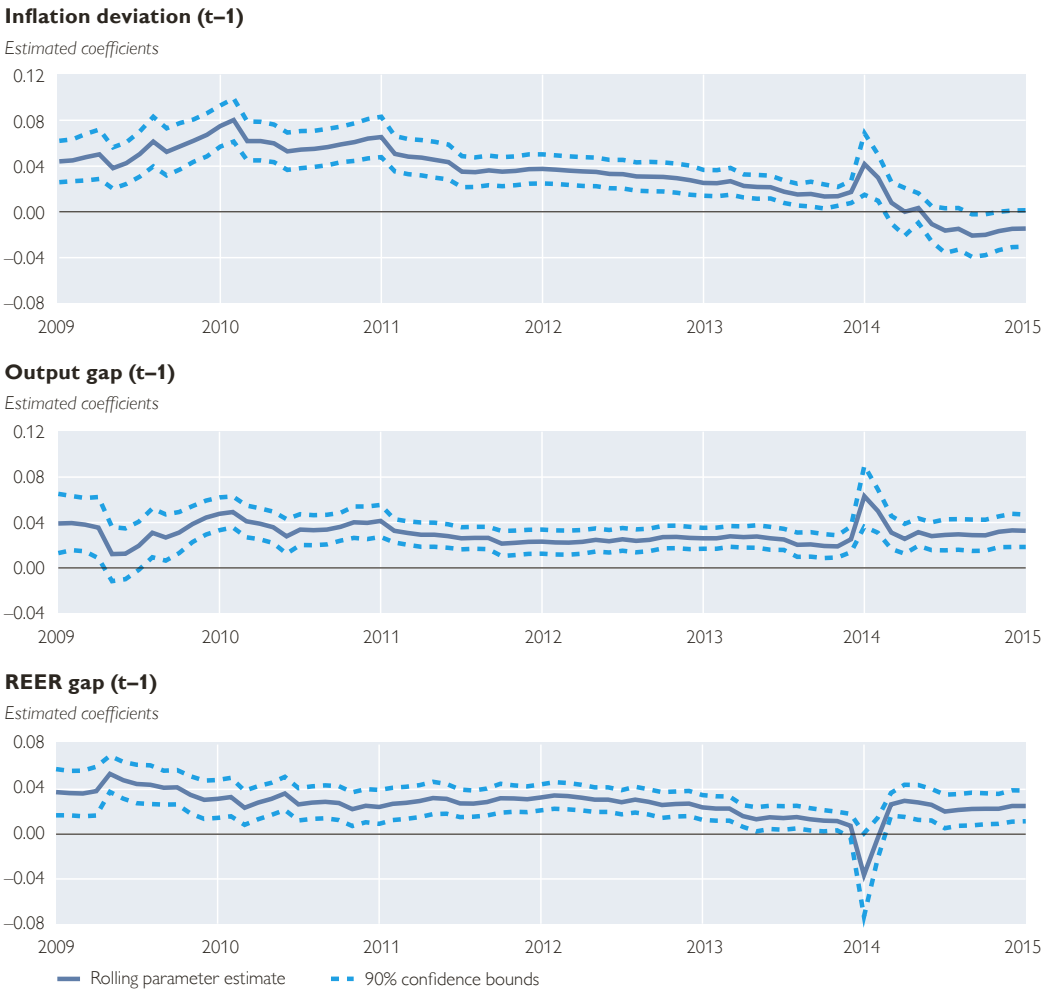
Annex

Chart A1

Rolling Taylor rule parameter estimates



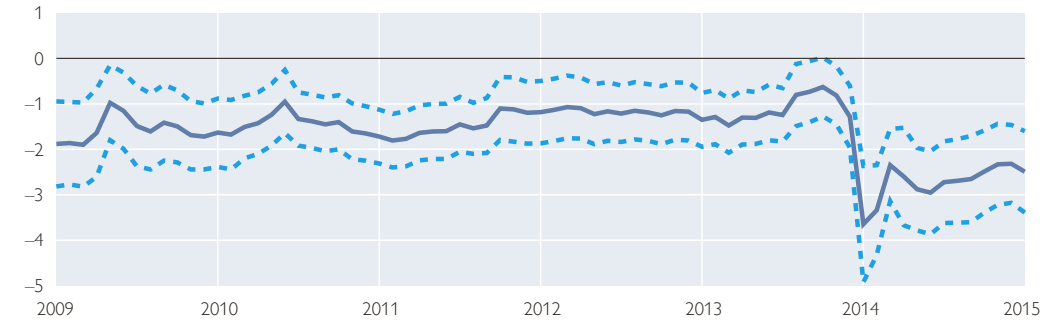
Source: Authors' calculations.

Chart A1 continued

Rolling Taylor rule parameter estimates

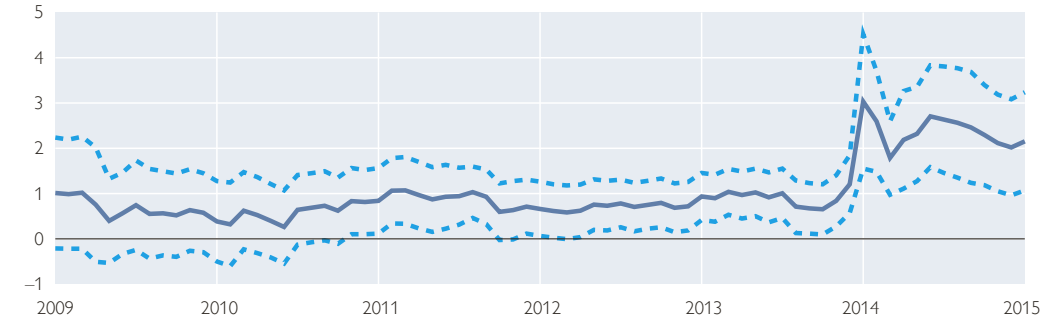
Oil price gap (t–1)

Estimated coefficients



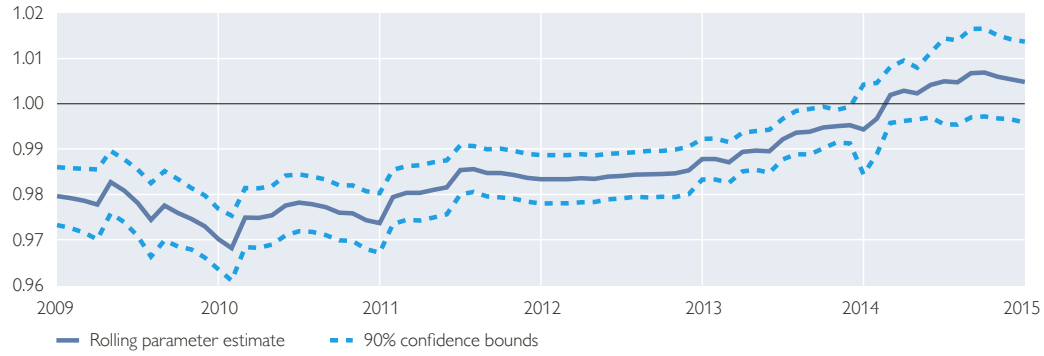
Oil price gap (t–2)

Estimated coefficients



Policy smoothing parameter

Estimated coefficients



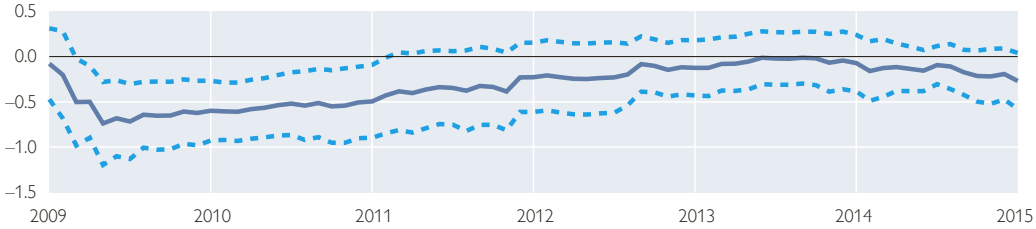
Source: Authors' calculations.

Chart A2

Rolling McCallum rule parameter estimates

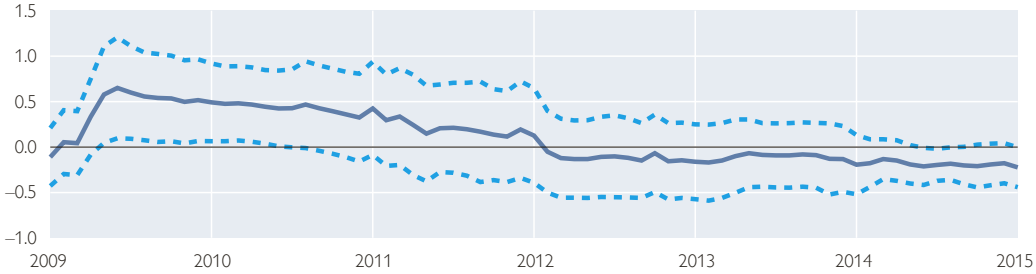
Nominal output gap (t–1)

Estimated coefficients



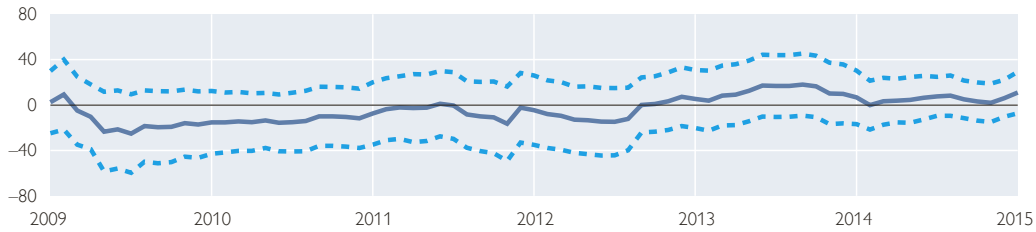
NEER gap (t–1)

Estimated coefficients



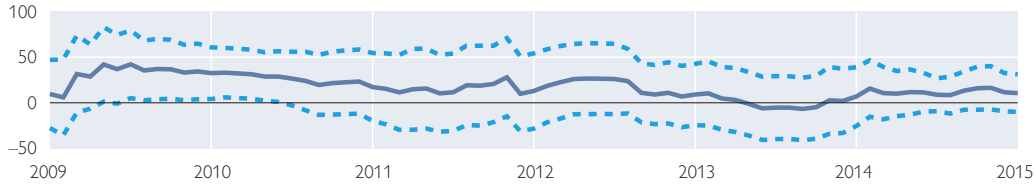
Oil price gap (t–1)

Estimated coefficients



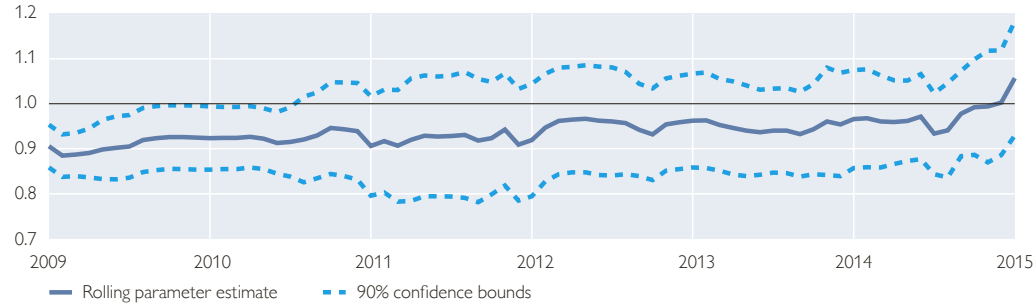
Oil price gap (t–2)

Estimated coefficients



Policy smoothing parameter

Estimated coefficients



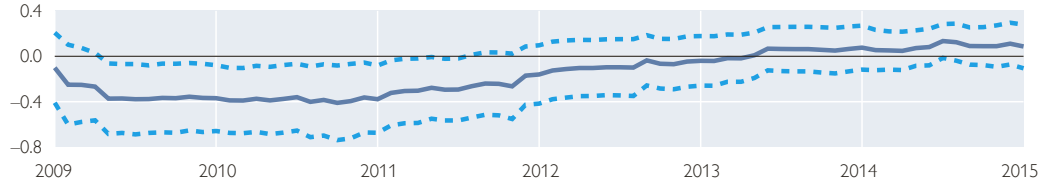
Source: Authors' calculations.

Chart A3

Rolling McCallum rule parameter estimates without oil prices

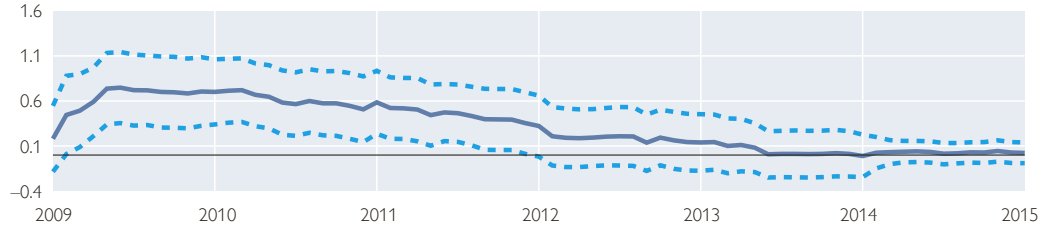
Nominal output gap (t–1)

Estimated coefficients



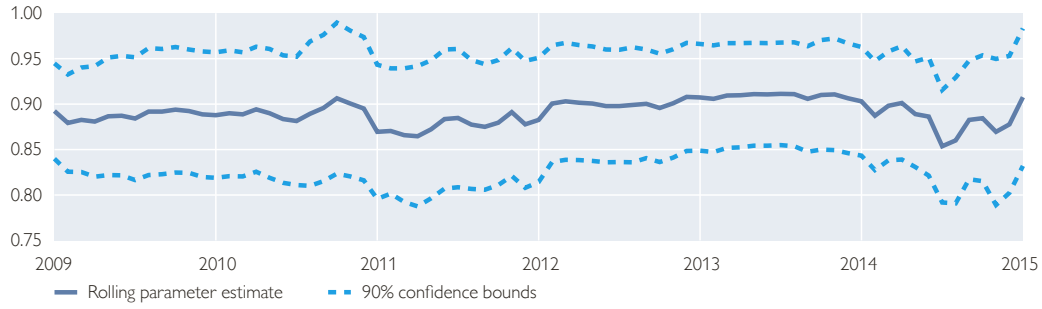
NEER gap (t–1)

Estimated coefficients



Policy smoothing parameter

Estimated coefficients



Source: Authors' calculations.

Table A1

Variables

Variable	Measure	Source[1]	Availability
Interest rate			
Refinancing rate	% per annum (end of period)	CBR	01/2000–11/2015
Central bank policy rate	% per annum (end of period)	CBR	02/2011–12/2015
Monetary aggregate			
Monetary base growth	year-on-year change (%) in RUB monetary base (broad definition)	CBR	12/2002–12/2015
Inflation			
Consumer price inflation deviation	CPI year-on-year inflation (%) less the average of the annual target range for CPI inflation[3]	FSSS, CBR	01/2000–11/2015
Output gap			
Real GDP growth gap	real year-on-year GDP growth (estimate) less HP trend[2]	EM	01/2001–11/2015
Nominal GDP growth gap	year-on-year change (%) in GDP in RUB less HP trend[2]	MF	01/2000–11/2015
Exchange rate gap			
Real effective exchange rate gap	REER index (2010=100) less HP trend[2]	BIS	01/2000–11/2015
Nominal effective exchange rate gap	NEER index (2010=100) less HP trend[2]	BIS	01/2000–11/2015
Oil gap			
Crude oil price gap	Urals oil price in USD, monthly average (index 2010=100) less HP trend[2]	OPEC	01/2000–11/2015

Source: Authors' compilation.

[1] BIS=Bank for International Settlements, CBR=Bank of Russia, EM=Ministry of Economic Development of the Russian Federation, FSSS=Russian Federal State Statistics Service (Rosstat), MF=Ministry of Finance of the Russian Federation, OPEC=Organization of the Petroleum Exporting Countries.

[2] Hodrick-Prescott filter applied to data series starting from 01/1999, where data are available. Smoothing parameter λ=14400.

[3] Inflation target may be changed during the year. When calculating the inflation deviation series we use the target inflation rate (range) available at the start of the year.

Table A2

Descriptive statistics

Variable	Obser-vations	Mean	Minimum	Maximum	Standard deviation	Skewness	Kurtosis	ADF t-stat.[1]	KPSS LM-stat.[2]
Interest rate									
Reference policy rate	167	12,04	7,75	25.00	4.17	1.30	4.22	–3.288**	0.288***
Monetary aggregate									
Base money growth	156	17.78	–13.45	54.77	15.28	0.14	2.16	–2.028**	0.044
Inflation									
CPI inflation deviation	167	3.28	–1.93	12.42	3.12	1.15	3.83	–1.095	0.125
Output gap									
Real GDP growth gap	167	–0.02	–12.22	5.47	3.13	–1.83	7.52	–2.939***	0.035
Nominal GDP growth gap	167	–0.22	–36.90	35.77	7.80	–0.45	8.35	–4.900***	0.043
Exchange rate gap									
REER gap	167	–0.09	–20.32	9.93	4.29	–1.25	7.18	–7.897***	0.058
NEER gap	167	–0.21	–18.32	9.08	4.70	–1.07	4.95	–6.609***	0.083
Oil gap									
Oil price gap	167	–0.00	–0.37	0.53	0.13	0.68	6.68	–4.589***	0.035

Source: Authors' calculations.

[1] The table presents the augmented Dickey-Fuller (ADF) unit root test statistic with a maximum of 13 lags. The intercept is included in the test equation if it is statistically significant.

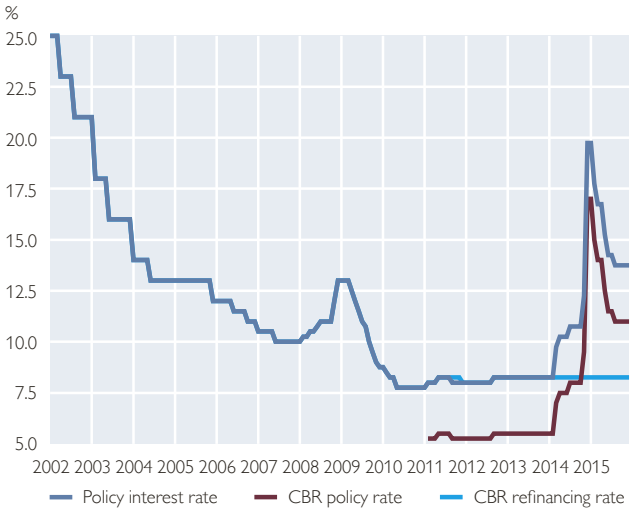
[2] The Kwiatkowski-Phillips-Schmidt-Shin (KPSS) Lagrange Multiplier (LM) test statistic evaluates the null hypothesis that the series is stationary. The trend term is included in the test equation if it is statistically significant.

*Note: Data for 01/2002–11/2015. ***, ** and * denote the 1%, 5% and 10% level of significance, respectively.*

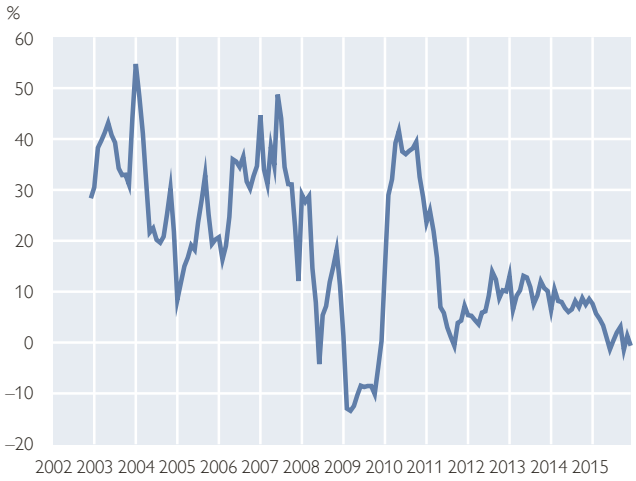
Chart A4

Data used in the policy rule analysis

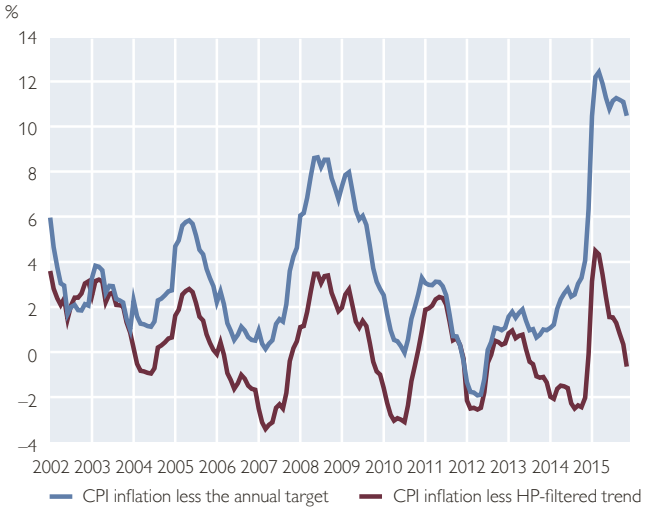
Policy interest rate



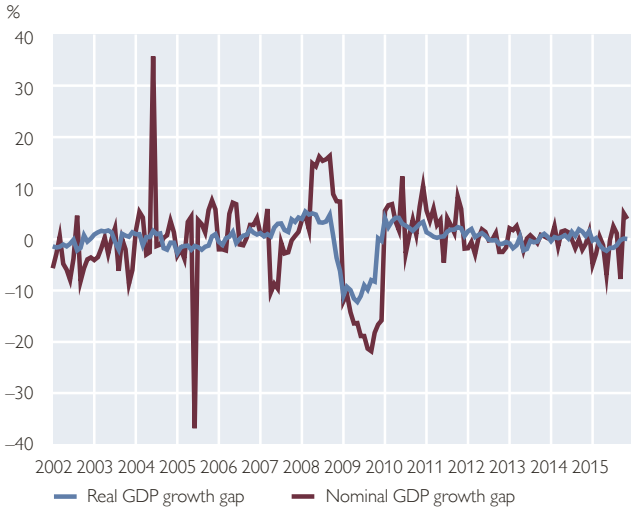
Growth in monetary base



Inflation deviation

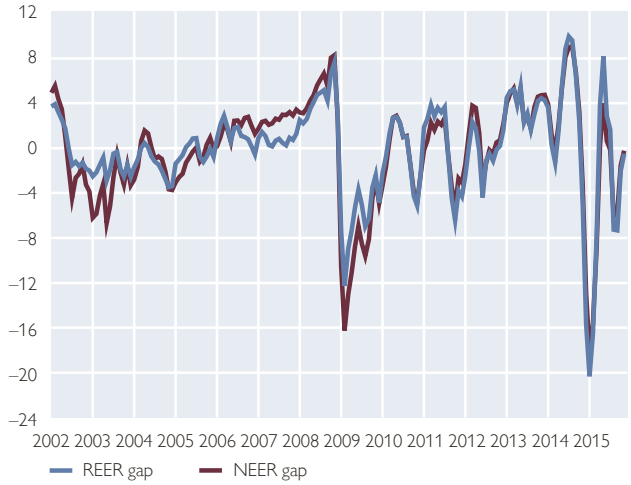


Output gaps



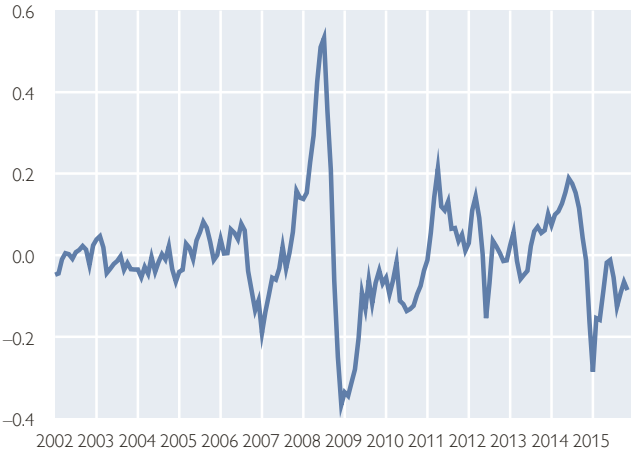
Exchange rate gaps

Deviation from trend



Oil price gap

Deviation from trend



Source: Authors' calculations and sources of variables stated in table A1.

Table A3

Correlations between the variables (individual samples, 01/2002–11/2015)

Correlation [t-stat.]	i_t	Δbm_t	$(\pi-\pi^*)_t$	$\Delta\hat{y}_t$	$\Delta\hat{x}_t$	$\widehat{reer}_t$	$\widehat{neer}_t$	$\widehat{oil}_t$
i_t	1.00							
	[]							
Δbm_t	0.18**	1.00						
	[2.27]	[]						
$(\pi-\pi^*)_t$	0.47***	−0.40***	1.00					
	[6.66]	[−5.38]	[]					
$\Delta\hat{y}_t$	−0.17**	0.50***	−0.27***	1.00				
	[−2.20]	[7.18]	[−3.51]	[]				
$\Delta\hat{x}_t$	−0.13	0.25***	−0.05	0.59***	1.00			
	[−1.64]	[3.23]	[−0.61]	[9.08]	[]			
$\widehat{reer}_t$	−0.39***	0.14*	−0.23***	0.33***	0.34***	1.00		
	[−5.22]	[1.73]	[−2.91]	[4.36]	[4.49]	[]		
$\widehat{neer}_t$	−0.47***	0.19**	−0.31***	0.49***	0.43***	0.95***	1.00	
	[−6.52]	[2.36]	[−4.01]	[6.92]	[5.93]	[36.10]	[]	
$\widehat{oil}_t$	−0.22***	−0.03	−0.06	0.52***	0.38***	0.55***	0.58***	1.00
	[−2.84]	[−0.35]	[−0.75]	[7.50]	[5.08]	[8.07]	[8.82]	[]

Source: Authors' calculations.

*Note: ***, ** and * denote the 1%, 5% and 10% level of significance, respectively.*

Table A3 continued

Correlations between the variables (individual samples, 01/2002–11/2015)

Correlation [t-stat.]	i_{t-1}	Δbm_{t-1}	$(\pi-\pi^*)_{t-1}$	$\Delta\hat{y}_{t-1}$	$\Delta\hat{x}_{t-1}$	$\widehat{reer}_{t-1}$	$\widehat{neer}_{t-1}$	$\widehat{oil}_{t-1}$
i_t	0.97***	0.18**	0.46***	−0.15*	−0.10	−0.37***	−0.43***	−0.19**
	[47.42]	[2.23]	[6.49]	[−1.93]	[−1.26]	[−4.85]	[−5.82]	[−2.45]
Δbm_t	0.19**	0.94***	−0.42***	0.52***	0.21***.	0.13	0.18**	−0.02
	[2.39]	[33.49]	[−5.69]	[7.45]	[2.71]	[1.63]	[2.25]	[−0.23]
$(\pi-\pi^*)_t$	0.49***	−0.37***	0.98***	−0.23***	−0.04	−0.27***	−0.33***	−0.06
	[6.97]	[−4.85]	[57.98]	[−2.92]	[−0.52]	[−3.48]	[−4.26]	[−0.70]
$\Delta\hat{y}_t$	−0.20**	0.46***	−0.32***	0.90***	0.45***	0.30***	0.45***	0.55***
	[−2.47]	[6.43]	[−4.11]	[26.27]	[6.18]	[3.93]	[6.32]	[8.19]
$\Delta\hat{x}_t$	−0.15*	0.26***	−0.06	0.64***	0.51***	0.32***	0.42***	0.47***
	[−1.83]	[3.36]	[−0.76]	[10.18]	[7.33]	[4.23]	[5.78]	[6.59]
$\widehat{reer}_t$	−0.36***	0.14*	−0.16**	0.36***	0.31***	0.82***	0.77***	0.55***
	[−4.77]	[1.71]	[−2.05]	[4.80]	[4.10]	[17.65]	[15.13]	[8.06]
$\widehat{neer}_t$	−0.46***	0.19**	−0.26***	0.51***	0.39***	0.82***	0.87***	0.62***
	[−6.36]	[2.40]	[−3.40]	[7.36]	[5.29]	[17.58]	[21.36]	[9.66]
$\widehat{oil}_t$	−0.23***	−0.02	−0.06	0.46***	0.30***	0.46***	0.47***	0.90***
	[−2.93]	[−0.31]	[−0.69]	[6.45]	[3.89]	[6.34]	[6.61]	[25.40]

Source: Authors' calculations.

*Note: ***, ** and * denote the 1%, 5% and 10% level of significance, respectively.*

Weathering global shocks and macrofinancial vulnerabilities in emerging Europe: Comparing Turkey and Poland

Markus Eller,
Florian Huber,
Helene Schuberth[1]

While both Turkey and Poland weathered the 2008/2009 crisis relatively well compared to other countries in Central, Eastern and Southeastern Europe (CESEE), their macrofinancial indicators responded fairly strongly to the Federal Reserve System's tapering announcement in May 2013. Among other things, marked currency depreciation and reversals in capital flows challenged policymakers in the region. To get a deeper understanding of the transmission of global shocks to macrofinancial variables in CESEE, we use a Bayesian global vector autoregressive model to investigate the effects of a U.S.-based monetary policy shock. Our simulation results suggest that both Turkey and Poland exhibit pronounced short- to medium-run macroeconomic responses to U.S. monetary policy shocks. The responses of Turkey tend to be somewhat stronger than those of Poland and the CESEE average, signaling the structurally different nature of the Turkish economy. Overall, the identified impulse responses largely resemble the actual post-tapering announcement developments.

JEL classification: C32, F41, F44, E52
Keywords: external vulnerabilities, international shock transmission, monetary policy shock, tapering, capital flows, GVAR, Turkey, Poland, CESEE

Central, Eastern and Southeastern Europe (CESEE[2]) was one of the fastest-growing regions in the world, with average annual real GDP growth rates of around 6.5% between 2003 and 2007. This outstanding growth performance was mainly driven by large foreign capital inflows, which fueled domestic credit growth, led to a surge in asset prices (in particular housing prices) and considerably boosted domestic demand. However, sizable GDP growth was generated on the back of rising vulnerabilities. Soaring prices and wages were one of the consequences of sharply rising domestic demand. As a matter of fact, double-digit inflation rates were not unusual during the pre-2008 boom years in several countries. Not only was lending growth in the years preceding the crisis very strong, but a large part of domestic loans to households and nonfinancial corporates was also denominated in foreign currency in a number of CESEE countries. Growing internal imbalances were also reflected in the development of the external sector. Increasing internal demand led to an appreciation of the exchange rate in countries with floating exchange rate regimes, which made exports more expensive and led to the buildup of substantial current account deficits in a number of countries.

Apparently, CESEE countries that were subject to stronger vulnerabilities and imbalances up to 2007 were hit harder during the 2008/2009 global financial crisis (see EBRD, 2009, or Bakker and Klingen, 2012). A sudden stop of capital

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[2] *Unless mentioned otherwise, our data is for the 11 CESEE EU Member States (CESEE-11: Bulgaria, Croatia, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Romania, Slovakia and Slovenia) and Turkey (referred to as the 12 CESEE economies in the following).*

inflows in the fall of 2008 triggered a sharp contraction of domestic demand, just when the slump in global trade hit the region's exports. This halt in capital inflows was a combination of a liquidity (credit supply) shock and a slump in export demand. The "Vienna Initiative," which ensured that banks maintained an exposure to subsidiaries in CESEE, together with stabilization packages of international financial institutions and the European Union (EU), was decisive in avoiding a much sharper contraction.

Only a few economies managed to escape relatively unscathed. Poland and Turkey share a rather favorable management of the 2008/2009 global financial crisis. In contrast to all the other EU countries, Poland did not experience a recession, while Turkey, after a short-lived contraction in 2009, quickly returned to sizable GDP growth rates. From 2010 to 2012, when foreign investors searched for yields in a low-interest global environment, both countries were among the main magnets for foreign capital in CESEE, with inflows approaching pre-crisis magnitudes. But the U.S. tapering announcement in May 2013 triggered a sharp repricing of risks and had large effects on emerging markets, resulting in substantial drops in stock market indices and large exchange rate depreciations (IMF, 2013; Aizenman et al., 2014). Both Poland and Turkey were affected, reflecting that CESEE is one of the regions which are perceived to be vulnerable to "risk on" and "risk off" modes (Bernanke, 2012) in global financial markets. These countries' particular way of managing the series of global economic and financial shocks since 2008, together with the fact that Poland and Turkey are the two largest economies in the investigated CESEE region, motivated a focus on these two countries in this article.

The paper is structured as follows: Section 1 discusses the reasons behind the relative success of Poland and Turkey in managing the impact of the 2008/2009 global financial crisis comparatively well. Section 2 looks at the post-2009 evolvement of macrofinancial risks in a comparative perspective, with a special emphasis on external vulnerabilities and banking sector risks. Section 3 studies the transmission of a contractionary monetary policy shock in the U.S.A. – exemplary for a global-scale, external shock – to domestic macroeconomic variables by means of a global vector autoregressive (GVAR) model. Section 4 concludes.

1 Poland and Turkey perform outstandingly in weathering the 2008/2009 crisis

During the 2008/2009 crisis, the CESEE region as a whole suffered larger output declines than any other region in the world (Berglöf et al., 2009). However, cross-country variation in crisis-related output declines was large. While several CESEE countries have still not been able to achieve significantly larger GDP than in 2008, Turkey and Poland in 2014 surpassed their 2008 GDP levels (at market prices) by 24% and 19%, respectively.

Poland was the only country in the EU to avoid a recession in 2009. Thereafter, it posted average annual real GDP growth of around 3% through 2014. Several factors are responsible for this success (EBRD, 2009; Bakker and Klingen, 2012): In the years before the crisis, Poland avoided significant macroeconomic imbalances, reduced fiscal deficits to 2% of GDP in 2007, ensured that inflation expectations were well-anchored, and restrained strong lending in foreign

currency early on. As a consequence, at about 24%, the share of domestic foreign exchange loans in total loans to the nonbank private sector was one of the lowest shares in the region in 2007. A Financial Stability Committee that issues early warnings and recommendations was established already in 2008.

When the crisis hit, exports contracted sharply and asset prices declined amid a sharp slowdown of capital inflows. But Poland had built up enough space to conduct countercyclical policies, implementing fiscal stimulus measures and lowering policy rates from 6% to 3.5% and reserve requirement rates from 3.5% to 3%. Prior to the crisis of 2008/2009, the Polish government lowered taxes in an effort to curb domestic consumption. As the Polish economy's export dependence is comparatively low, these fiscal stimulus measures helped to diminish the impact of the crisis significantly. The exchange rate appreciated by 50% against the euro between 2004 and 2008 and played a shock-absorbing role during the crisis when the złoty depreciated by 30% against the euro through February 2009 (see Stątżka-Gawrysiak, 2009). Furthermore, liquidity (foreign exchange swaps) and banking stabilization measures (increase in the deposit insurance limit, credit guarantee program) were taken. Also, the (unused) IMF Flexible Credit Line of 2009 was effective in stabilizing market expectations and helped maintain access to international capital markets (Bakker and Klingen, 2012).

Like Poland, Turkey experienced robust growth in the run-up to the crisis. Given low domestic savings, growth was largely fueled by capital inflows. But the real exchange rate appreciation led to the buildup of current account deficits, with the current account moving from a surplus in 2001 into a deficit of about 6% of GDP in 2007. In 2009, Turkey suffered a relatively moderate recession – compared to other countries in the region – followed by an immediate and very significant recovery in 2010 and 2011 and a renewed moderation of growth thereafter. As a result, Turkey's average annual real GDP growth rate of about 5.5% between 2010 and 2014 was not only one of the largest in the region but also associated with comparatively strong volatility.

The resilience of the Turkish economy and in particular of the Turkish banking sector during the short, V-shaped recession can be partly explained by the legacy of financial system restructuring and by the early implementation of macroprudential tools in the aftermath of the crisis of 2001, which led to benign credit growth (Akkoyun et al., 2013). In 2007, the share of foreign currency loans in total loans to resident private nonbanks was about 30%, somewhat below the CESEE average. Foreign currency lending of corporates was restrained, and in 2009, banks were prohibited from lending in foreign currency (or foreign-indexed loans) to households. In 2008, dividend payouts were curtailed to bolster bank-retained earnings and capital. During the crisis, banks' capital adequacy ratios were higher than the Basel II requirement of 8%. Another factor behind Turkey's relative crisis resilience was the swift and decisive use of countercyclical macroeconomic policies during 2008/2009 (Bakker and Klingen, 2012). Turkey implemented a fiscal stimulus package, and the Turkish central bank (Türkiye Cumhuriyet Merkez Bankası, TCMB) lowered the policy rate by 10.25 percentage points between the end of 2008 and 2009. Reserve requirements were reduced as well. As a consequence, the exchange rate depreciated and the current account deficit declined.

In the run-up to and during the 2008/2009 crisis, Poland and Turkey shared some similarities. Both instituted macroeconomic reforms that created room for maneuver to conduct countercyclical policies during the crisis. Another parallel is the early use of macroprudential tools, in particular to curb foreign currency lending.

2 Macrofinancial risks in a comparative perspective

Picking up the argument that countries might be hit more strongly by external shocks if domestic macrofinancial vulnerabilities are more pronounced (e.g. EBRD, 2009; Bakker and Klingen, 2012; IMF, 2013; Mishra et al., 2014), this section aims to provide a brief overview of the macrofinancial risk profiles of Turkey and Poland by investigating the developments of capital flows, exchange rates, cross-border banking and domestic banking sector stability indicators since 2009 in general and since the Federal Reserve System's (Fed's) tapering announcement in May 2013 in particular.

Macrofinancial developments in emerging market economies after 2009 can be characterized by two different stages. During the 2010 to 2012 period, capital inflows resumed quite strongly, associated with a shift of capital from low yields in advanced economies to higher returns in emerging markets. At the same time, private sector credit growth regained momentum. However, Federal Reserve Chairman Ben Bernanke's remark before the U.S. Congress on May 22, 2013, that the Federal Open Market Committee could take a step down in the pace of asset purchases if economic improvement appeared to be sustained (Bloomberg, 2013) stood for a new turning point. This indication of a phaseout of the Fed's expansionary monetary policy stance and the related expectation of an increase in interest rates in the U.S.A. implied sizable capital outflows and/or a reduction in capital inflows as well as a depreciation of currencies in emerging markets. As a result, several emerging market economies, including Turkey, sharply hiked policy rates in early 2014 to stabilize their exchange rates and to rein in capital outflows. However, macrofinancial pressure on emerging markets has continued not only due to tapering in the U.S.A., eventually followed by the first hike in the federal funds rate in December 2015 in seven years from near zero, but also due to geopolitical tensions and a cooling-off of the Chinese economy. At the same time, expansionary monetary policy in the euro area was intensified in March 2015 with the start of the Eurosystem's Public Sector Purchase Programme, which is likely to have cushioned, at least to a certain extent, the international spillovers of tighter monetary policy in the U.S.A.[3]

The spillovers of advanced economies' monetary policy decisions to emerging markets point to pronounced global macrofinancial interdependencies. It should be noted that domestic macroeconomic fundamentals in emerging markets play a decisive role, too. Apparently, emerging market economies with stronger macroeconomic fundamentals, deeper financial markets, and a tighter macroprudential

[3] *So far, there is hardly any empirical evidence on the international spillovers of combined monetary policy shocks in the U.S.A. and the euro area. Chen et al. (2015) use a global vector error correction framework to compare the impact of unconventional monetary policy measures both in the U.S.A. and in the euro area. They find that U.S. unconventional monetary policy generally has stronger domestic and cross-border impacts than euro area nonstandard measures; this partly also holds for the cross-border transmission to selected CESEE countries. Feldkircher (2015) resorts to a global vector autoregression model and shows that the real economy in CESEE reacts nearly equally strongly to a contractionary U.S. monetary policy shock and to a corresponding euro area shock.*

policy stance (including capital flow management measures) in the run-up to the tapering announcement experienced smaller currency depreciations and smaller increases in government bond yields in 2013 to 2014 (IMF, 2013; Mishra et al., 2014).

Turning to the countries of interest in this study, chart 1 shows that the Turkish lira depreciated comparatively strongly against the U.S. dollar from May 2013 until the end of January 2014, before the TCMB raised the one-week repo rate by 550 basis points from 4.5% to 10%. In the summer of 2014, when the downward pressure on the currency subsided and the risk premium on Turkish assets fell, the TCMB was in a position to cut the policy rate (by a total of 175 basis points until early summer 2015). However, in the third quarter of 2014, renewed depreciation set in and continued until very recently. Overall, since the beginning of 2013, the Turkish lira has lost about 40% of its value against the U.S. dollar and about 25% against the euro.

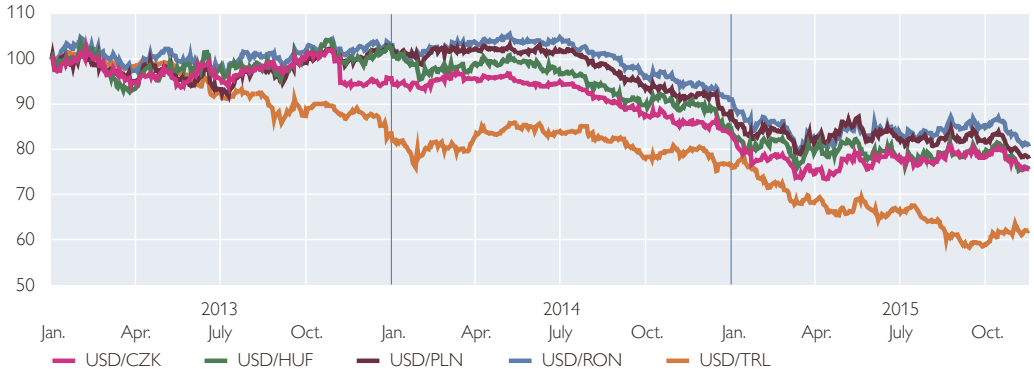
In contrast to the Turkish economy, the Polish economy is more affected by developments of the euro than of the U.S. dollar, given the structure of foreign trade and foreign exchange liabilities. Immediately after the Fed's tapering announcement, the Polish złoty experienced only a short-lived depreciation against the euro; thus, Narodowy Bank Polski (NBP) did not have to raise the policy rate. Since then, the NBP has kept the currency's value against the euro more or less unchanged, while the value weakened against the U.S. dollar because the euro depreciated against the U.S. dollar.

Chart 1

CESEE exchange rates and policy rates

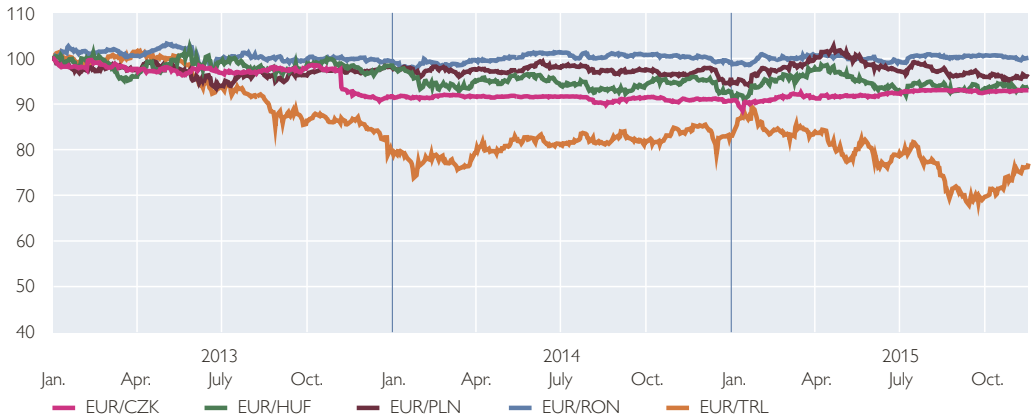
Exchange rate developments versus the U.S. dollar

January 1, 2013=100 (rise=appreciation) *Latest observation: November 17, 2015*



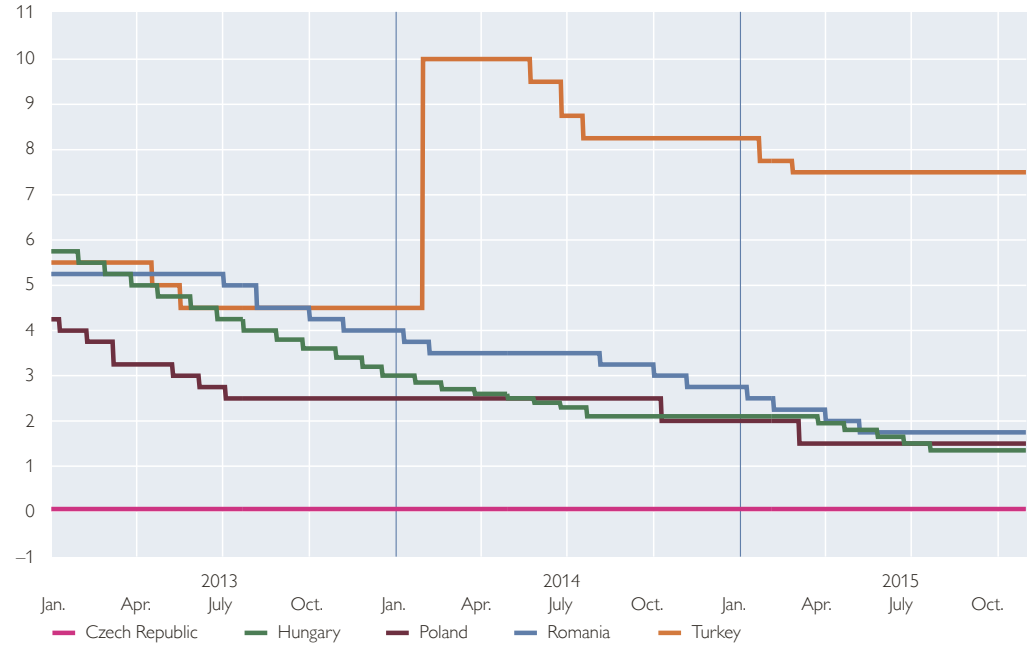
Exchange rate developments versus the euro

January 1, 2013=100 (rise=appreciation) *Latest observation: November 17, 2015*



Policy rate developments in CESEE

% *Latest observation: October 30, 2015*



Source: Thomson Reuters, national central banks.

Capital flows to Poland and Turkey surged considerably again after the marked drops in 2009 (see chart 2). In Poland, these dynamics lasted until mid-2011 and largely reflected net portfolio inflows. Spillovers from the euro area sovereign debt crisis were apparently responsible for a pronounced net outflow of currency and deposits in Poland from late 2011 until early 2013. Following the Fed's tapering announcement, net portfolio inflows also declined quite substantially and ultimately resulted in a financial account deficit in the first half of 2014. Since then, portfolio and other investment inflows have not yet resumed considerably. It should be noted, however, that this reduction in net capital inflows also went hand in hand with a correction of the current account deficit. While the Polish income balance deficit (much of which can be explained by repatriated earnings of foreign-owned firms) is still quite sizable, the goods and services balance has recorded surpluses since 2013.

Turkey was able to keep the positive capital flow dynamics until early 2013. Net portfolio inflows widened steadily during this period, but inflows of loans and deposits (e.g. remittances) also played a considerable role as part of "other investment." However, after the tapering announcement, portfolio investment inflows steadily declined, while inflows of loans largely kept their level. In the first half of 2015, Turkey was confronted with some additional reduction in portfolio investment flows. In contrast to Poland, Turkey has so far not been able to substantially correct its current account deficit (largely a deficit in the goods and services balance). The current account deficit moderated somewhat in the first half of 2014, falling to around 6% of GDP (down from 8% at the end of 2013). However, this rebalancing was mostly driven by the normalization in the gold balance along with weak domestic demand. Moreover, the financing of the current account deficit remains rather fragile, given the comparatively large share of short-term (non-FDI) flows in the financial account.

Parts of the discussed changes in capital flows consist of changes in banking capital flows, i.e. direct cross-border lending activities. Another channel of international shock transmission via banks consists of lending through foreign-owned affiliates, which is generally perceived to be less volatile than direct cross-border lending (Milesi-Ferretti and Tille, 2011). Chart 3 shows that claims of BIS-reporting banks on CESEE economies have declined since the 2008/2009 crisis primarily through cross-border lending (right-hand panel), while consolidated claims (including lending through affiliates, left-hand panel) have also clearly lost momentum but have on average remained unchanged. This development is partly indicative of the success of the Vienna Initiative.[4]

In the search-for-yield period from 2010 to 2011, cross-border lending to both Poland and Turkey experienced a remarkable revival. However, in 2012, when countries in CESEE were increasingly confronted with contagion effects from the euro area sovereign debt crisis, cross-border claims on CESEE declined again (even more strongly than in 2009), with the notable exception of those on Turkey, which was able to avoid a reduction in both consolidated and cross-border claims. Finally, associated with the Fed's tapering announcement in May 2013, cross-

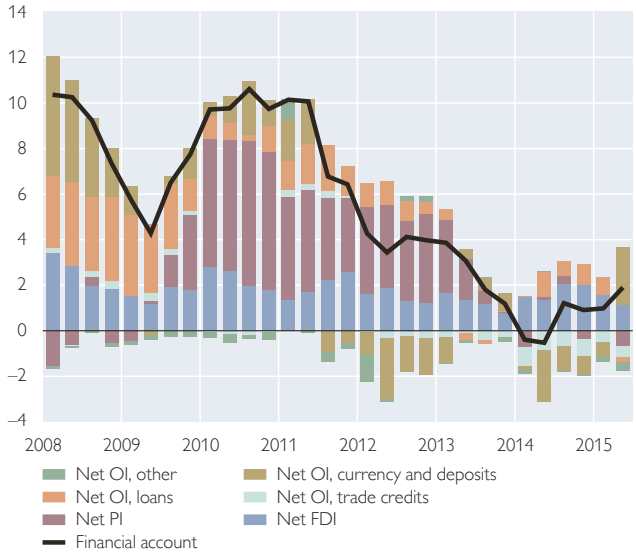
[4] *See http://vienna-initiative.com/. De Haas et al. (2012) show that foreign banks that participated in the Vienna Initiative were relatively stable lenders in CESEE. This is also confirmed by Hameter et al. (2012), who find that intragroup cross-border credit from Austrian banks was more stable than lending to nonaffiliated borrowers in CESEE during the 2008/2009 crisis.*

Chart 2

Decomposition of annualized balance of payments

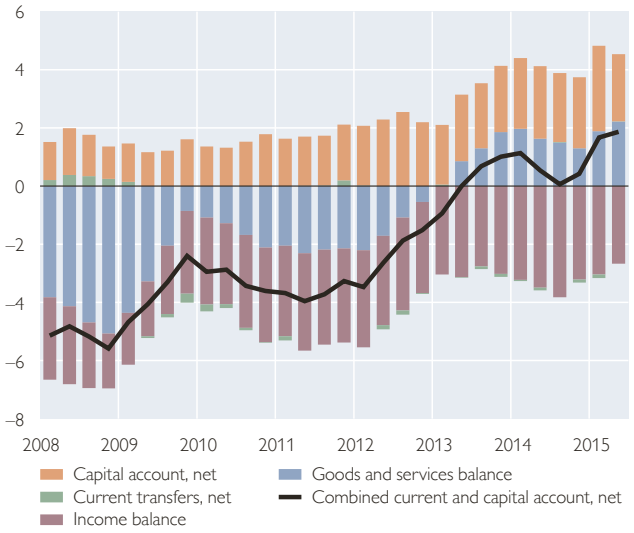
Poland: financial account

% of GDP (cumulative four-quarter moving sums)



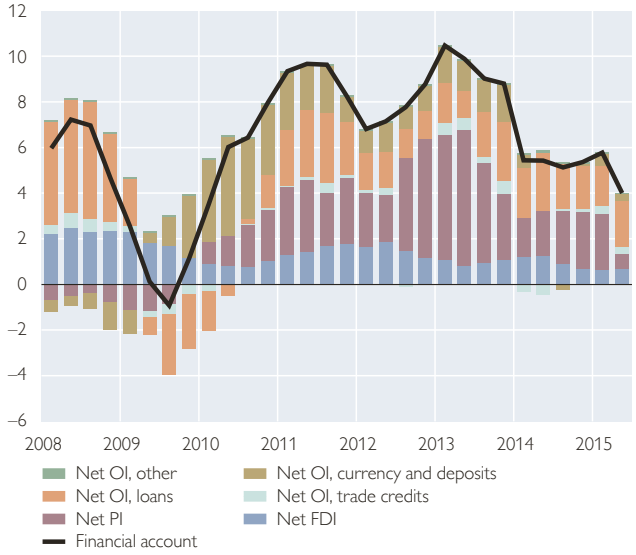
Poland: current and capital account

% of GDP (cumulative four-quarter moving sums)



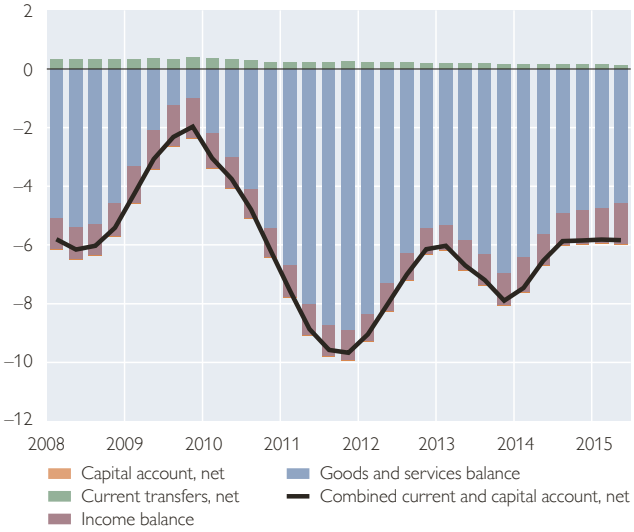
Turkey: financial account

% of GDP (cumulative four-quarter moving sums)



Turkey: current and capital account

% of GDP (cumulative four-quarter moving sums)



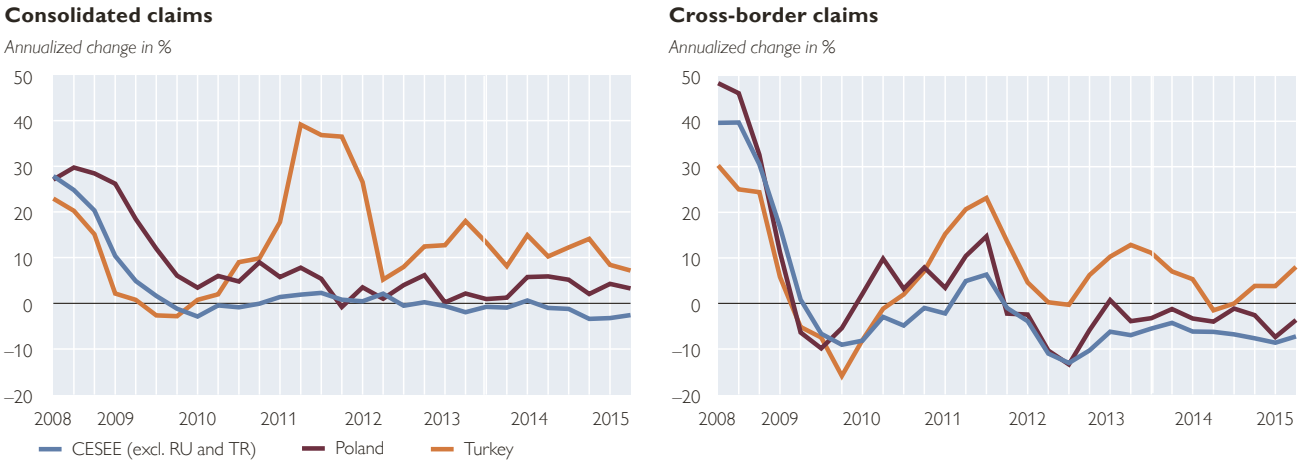
Source: IMF, Eurostat, OeNB, national central banks.

Note: OI=other investment; PI=portfolio investment; FDI=foreign direct investment.

border bank lending has continued to decline in Poland, while Turkey experienced a pronounced slowdown from an annual growth rate of 13% in the first quarter of 2013 to –1.5% in the first quarter of 2014 before growth rebounded remarkably to 8% in the first quarter of 2015. At the same time, it should be noted that consolidated claims have not lost considerable momentum since early 2013.

Chart 3

Outstanding stocks of all BIS reporting banks vis-à-vis CESEE



Source: national central banks, IMF, OeNB.

As emphasized by the EBRD (2009) or Bakker and Klingen (2012) for the 2008/2009 crisis and by Mishra et al. (2014) for the impact of the Fed's tapering announcement, structural banking sector variables are crucial in explaining the intensity of domestic macroeconomic responses to an external shock.

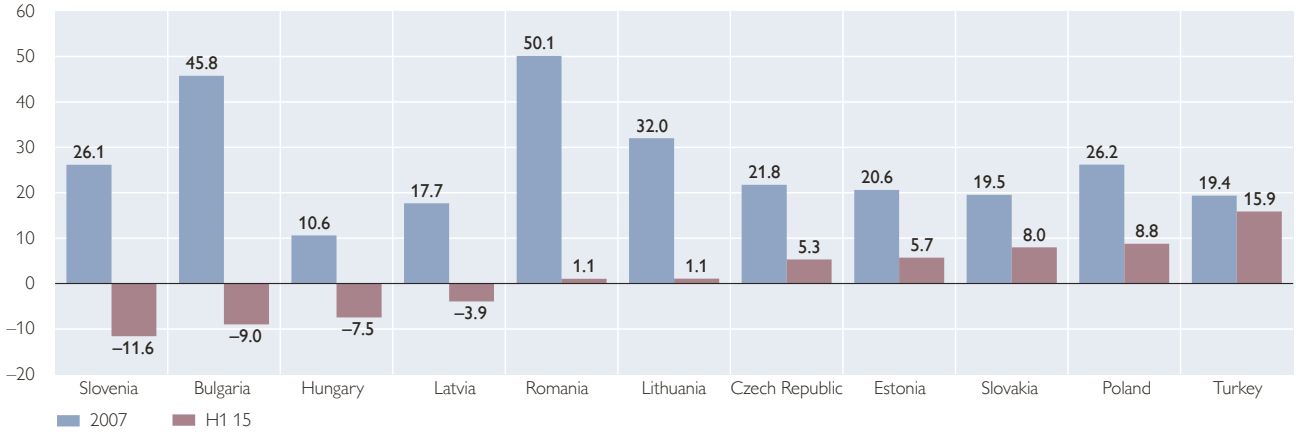
Compared to very strong private sector credit growth in several CESEE economies before 2008, we have seen subdued or negative credit growth rates in the region since 2009 (chart 4). Clear signs of a revival in credit in the region as a whole have yet to emerge. Turkey is an important exception and has continued to record respectable credit growth rates after 2009. Although it fell somewhat in the first half of 2014, growth of domestic credit to the nonbank private sector experienced some pickup in Turkey in 2015 and expanded by more than 15% (year on year, inflation-adjusted) in the first half of 2015. Looking at the composition of domestic credit, the share of foreign currency loans in total loans to resident private nonbanks reached about 37% in Turkey and 27% in Poland in September 2015. As mentioned before, Turkish banks are no longer permitted to lend to households in foreign currency; they can offer foreign currency loans only to corporates that have income in foreign currency. Consequently, new foreign exchange loans have been mainly extended to large trading firms that have sufficient access to financial hedging (OECD, 2014). Credit dynamics also have to be seen relative to the development of deposits. While at the end of 2009, nearly 85% of loans to resident nonbanks were covered by deposits in Turkey, the above-mentioned pronounced credit growth caused this ratio to deteriorate steadily to just 70% in September 2015. Poland, on the other hand, was not able to significantly raise the coverage of loans by deposits (75% in 2009 compared to about 77% in September 2015).

Basel II standards have been implemented in Turkey since July 2012, so far with a limited impact on capital adequacy in the banking sector. While the tier 1 capital adequacy ratio (CAR) stood at more than 17% in Turkey at the end of 2009, it steadily declined amid a marked credit expansion to a bit more than 12%

Chart 4

Domestic credit to the nonbank private sector: pre-crisis versus post-crisis growth

Year-on-year percentage change in period-end stock, HICP-deflated



Source: national central banks, OeNB.

in June 2015. In the same period, Poland was able to raise the the tier 1 CAR from 12% to 14%. The profitability of banks has deteriorated in Turkey since 2009, though it is still large compared to that in other countries in the region. Whereas the return-on-assets ratio stood at 1.3% in Turkey in June 2015, it has halved compared to end-2009. In Poland, in turn, the return-on-assets ratio improved somewhat from 0.8% to 1% in the same period.

3 GVAR simulation of the economic transmission of a U.S. monetary policy shock

While in the previous section, we reviewed domestic macrofinancial vulnerabilities that are considered relevant for the intensity of country-specific responses to external shocks, in this section we try to get a better understanding of the possible macroeconomic responses of CESEE countries – in particular Turkey and Poland – to a global-scale external shock. For this purpose, we use a global vector autoregressive (GVAR) model and simulate the impact of a contractionary monetary policy shock in the U.S.A. in recognition of that country's pivotal role in shaping the global business cycle (see Feldkircher, 2015). Given the comparably stronger trade integration of the CESEE region with the euro area than with the U.S.A., the ongoing monetary accommodation in the euro area might have some counterbalancing impact,[5] but a systematic comparison of the impact of Fed- versus ECB-induced monetary policy shocks would be beyond the scope of the present paper.

In recent years, several authors have started to focus on the international economic transmission of U.S. monetary policy shocks across the globe. Among others, Canova (2012) studies the influence of U.S.-based shocks on Latin American economies. He finds that monetary policy shocks produce significant fluctuations

[5] *Globan (2015) analyzes the spillover effects of monetary policy shocks in the euro area to seven non-euro area CESEE EU Member States. He finds that over the last years, macroeconomic developments in the euro area have become increasingly important drivers of capital inflows in CESEE.*

abroad, while demand and supply shocks tend to produce insignificant responses. The IMF (2014) detects a lagged, relatively short-lived, negative GDP growth response in emerging market economies (including Turkey and Poland) to a contractionary U.S. monetary policy shock. The impact of external shocks on capital flows has been explicitly analyzed for Turkey in Özen et al. (2013). The authors determine that external financial stress (proxied by a positive shock in the Chicago Board Options Exchange's Volatility Index, VIX) results in a marked decline in net portfolio investment.

While the literature surveyed above explicitly investigates the international transmission of shocks by means of conventional vector autoregressive (VAR) models, most of these studies remain confined to two-country models, neglecting second- and third-round spillover effects. This exclusion implies that these simpler specifications potentially provide biased estimates, underestimating relevant effects by ignoring reactions stemming from other countries. Thus, modeling approaches that simultaneously model a large set of economies have gained popularity recently. Georgiadis (2015) studies global spillovers from identified U.S. monetary policy shocks in a GVAR model and finds that U.S. monetary policy generates sizable output spillovers to the rest of the world. Feldkircher and Huber (2016) use a Bayesian GVAR model to analyze international spillovers of a contractionary U.S. monetary policy shock and of expansionary U.S. aggregate demand and supply shocks. They show that the monetary policy shock has strong cross-border spillovers on output and prices.

In the present paper, we apply the same methodological framework as in Feldkircher and Huber (2016), but we differ by using an updated dataset, by including financial account variables and by explicitly showing country-specific evidence for Poland and Turkey.

3.1 The GVAR model

The GVAR model put forward by Pesaran et al. (2004) constitutes a flexible means of incorporating large information sets in the modeling framework. Successful applications of the GVAR methodology range from the analysis of global shocks (see e.g., Dees et al., 2007, Pesaran et al., 2007, Feldkircher and Huber, 2016) to forecasting (Crespo Cuaresma et al., 2016).

The point of departure is the individual country model for country $i = 0,...,N$, which is assumed to be a VAR(1,1) model[6] featuring exogenous regressors

$$x_{it} = a_{i0} + a_{i1}t + \Psi_{i1}x_{it-1} + \Lambda_{i0}x^{*}_{it} + \Lambda_{i1}x^{*}_{it-1} + \epsilon_{it} \quad (1)$$

where x_{it} is a $k_i \times 1$ vector of endogenous variables measured at time t and a_{ij} $(j = 1,2)$ denotes $k_i \times 1$ vectors of coefficients associated with the constant and trend. Furthermore, Ψ_{i1} is a $k_i \times k_i$ parameter matrix corresponding to the first lag of the endogenous variables and Λ_{ik} $(k=0,1)$ are $k_i \times k^{*}_i$ dimensional parameter matrices corresponding to the (weakly) exogenous variables x^{*}_{it} defined as:

[6] *Our model is heavily parameterized, and even in the presence of a Bayesian approach, the limited time span available suggests that higher lag orders lead to a proliferation of parameters, ultimately producing unstable and imprecise results. Thus, we have opted to include only one lag of each variable type showing up in the VAR model.*

$$x_{it}^{*} = \sum_{j \neq i} w_{ij} x_{jt} \tag{2}$$

where w_{ij} are weights between countries i and j, usually set to bilateral trade linkages.[7] These weakly exogenous variables aim to approximate cross-country linkages. It can easily be seen that the specific structure of the model in (1) implies parametric restrictions on variables of other countries. Finally, $\epsilon_{it} \sim N(0, \sum_i)$ is a standard vector white noise error term.

It is straightforward to show that a sequence of the models described in equation (1) can be solved to yield a global representation of the model. As a consequence, (weakly) exogenous variables become effectively endogenous, and the global system resembles a standard large dimensional VAR given by

$$x_t = b_0 + b_1 t + F\ x_{t-1} + e_t \tag{3}$$

where x_t denotes the global vector, consisting of the stacked endogenous variables of all countries, i.e. $x_t = (x'_{0t}, \ldots, x'_{Nt})'$. The coefficient matrices of the deterministic part b_0, b_1 and the matrix corresponding to the lagged endogenous variables F are complex functions of the underlying estimates originating from the local models and the weightings used.

Note that equation (3) is a standard VAR(1,1) model with a deterministic constant and trend. All textbook formulas for functions of the parameters like impulse responses, forecasts or forecast error variance decompositions apply. To ensure stationarity of the model, we have to impose that $eig(F) < 1$. Technically, this rules out explosive behavior of the model. From an economic point of view, this restriction states that policymakers try to smooth possible impacts of shocks hitting the economy.

3.2 Prior setup and estimation

The GVAR model is usually estimated using standard techniques like maximum likelihood. However, recently Bayesian methods have proved to be a good alternative (see Crespo Cuaresma et al., 2016; Feldkircher and Huber, 2016). While standard techniques are easy to use, they are prone to overfitting the data. This directly translates into the well-known "curse of dimensionality," which implies that a strong in-sample fit leads to weak out-of-sample forecasting performance. Hence, following the literature on Bayesian VARs (Sims and Zha, 1998), we use a conjugate Minnesota prior, which has a proven track record in forecasting applications. This implies using a Gaussian prior on the coefficients in equation (1) and an inverted Wishart prior on $\sum_i$. Intuitively, the mean and variance for the prior on the coefficients are set such that the model in equation (1) is shrunk toward a random walk, implying that the first own lag of a variable is perceived to be an important predictor. Higher lag orders are assumed to be less important, implying that the prior variances on the corresponding coefficients are set to small values.

[7] *For a very similar dataset, Crespo Cuaresma et al. (2016) and Feldkircher and Huber (2016) find that while mixtures of weights (i.e. using trade weights for real variables and financial weights for financial variables) outperform other alternatives in terms of marginal likelihoods, the final impact on the results is rather negligible.*

Estimation and computation of the impulse response functions is a straightforward application of Monte Carlo integration. Because impulse responses are highly nonlinear functions of the parameters, we have no closed-form posterior solutions for them. However, in the natural conjugate case, the (conditional) posteriors for the coefficients have well-known distributional forms. Thus, it is straightforward to set up a simple Markov chain Monte Carlo scheme to estimate the local models and compute the corresponding impulse response schedules. More detailed information on the Minnesota prior in a GVAR framework can be found in Crespo Cuaresma et al. (2016).

3.3 Data overview

We rely on an updated variant of the dataset put forward in Feldkircher and Huber (2016). This dataset covers 42 economies and the euro area as a regional aggregate (representing over 90% of global output) for the time period from Q1 1995 to Q4 2013. Table 1 presents the countries included in the analysis.

Table 1

Country coverage

Rest of the world (11): US, EA, GB, CA, AU, NZ, CH, NO, SE, DK, IS
CESEE (12): CZ, HU, PL, SK, SI, BG, RO, HR, LT, LV, EE, TR
CIS & Western Balkans (6): RU, UA, BY, GE; AL, RS
Asia (9): CN, KR, JP, PH, SG, TH, ID, IN, MY
Latin America (5): AR, BR, CL, MX, PE

Source: Authors' compilations.

Note: Abbreviations generally represent the two-digit ISO country code; EA denotes euro area.

We use a standard set of macroeconomic aggregates, including GDP, inflation, real exchange rates measured vis-à-vis the U.S. dollar, short- and long-term interest rates, trade and financial account balances and finally the price of oil as a global control variable. Table 2 provides a brief description of the variables included.

Table 2

Variable description

Variable	Description	Source
y	Real, seasonally adjusted GDP, 2005=100, in logarithms	IMF
Δp	Rate of consumer price inflation, based on seasonally adjusted CPI	IMF
e	Real exchange rate relative to the U.S. dollar (deflation based on national CPI levels)	IMF
i_s	Typically, three-month money market rate (annualized)	IMF
i_L	Typically, yield on 10-year-government bonds (annualized)	IMF
tb	Ratio of real exports to real imports, in logarithms	IMF
fa	Financial account (excl. reserve assets) relative to GDP, cumulative moving annual values	IMF, NCBs
$poil$	Price of oil, seasonally adjusted, in logarithms	IMF
w_{ij}	Bilateral average trade flows between countries i and j	OECD

Source: Authors' compilations.

Note: For more details on data definitions and compilation, see Feldkircher (2015).

The choice of the variables is standard in the literature on GVAR modeling. However, inclusion of the financial account allows us to gain a deeper understanding

of the role of capital movements in the transmission of economic shocks. Note that in this part of the analysis, we have not included the additional variables discussed in section 2, such as structural banking sector indicators, although they could affect the intensity of domestic macroeconomic responses to an external shock. The main reasons are limited data availability for the large country sample and limited degrees of freedom in the estimations.

The set of weakly exogenous variables is constructed using bilateral average trade flows over the estimation window. This choice aims to approximate the underlying relationship between countries. Other possible choices include weighting schemes based on financial or geographical weighting. However, we focus exclusively on a weighting scheme based on trade weights because this seems to deliver more robust results than financial weights (as proxied through bilateral banking exposure), as the latter usually prove to be more volatile.

It is worth noting that the individual country models are constructed to include all variables described in table 2 for all countries, if available. One exception is the long-term interest rate, which is not available for some emerging market economies. Moreover, in the case of the U.S. country model, we obviously did not use the real exchange rate vis-à-vis the U.S. dollar but the real effective exchange rate (based on the CPI). All weakly exogenous variables except the weakly exogenous real exchange rate are included. The latter is included only in the U.S. country model. For more information, see Feldkircher and Huber (2016).

3.4 Shock identification

The model described in section 3.1 is completely atheoretical, as reduced-form impulse responses generally report the response of some interesting variable of interest to a weighted average of different structural shocks. To identify the effects of a monetary policy shock, researchers have opted for several possible identification schemes. However, we follow Eickmeier and Ng (2015) and Feldkircher and Huber (2016) and impose sign restrictions on the impulse response functions of the U.S. country model to retrieve the structural GVAR representation. In contrast to other identification schemes, this scheme gives us more flexibility than restrictions on the short-run behavior of the impulse response functions. As alternatives to structural identification schemes, Pesaran et al. (2004) advocated the use of generalized impulse response functions (GIRFs). These GIRFs, however, have no theoretical interpretation, rendering the use of this approach unfeasible for our research objectives.

Loosely speaking, sign restrictions rotate a given set of orthogonal responses until a prespecified set of restrictions is fulfilled. This is achieved by sampling orthonormal rotation matrices using the algorithm outlined in Rubio-Ramirez et al. (2010). Using such a rotation matrix, we compute the corresponding impulse response schedules using Monte Carlo integration.[8]

[8] *More details on how this procedure works can be found in Feldkircher and Huber, 2016.*

Specifically, we impose the following set of sign restrictions:

Table 3

Sign restrictions

Shock	y	Δp	i_S
AD	↑	↑	↑
AS	↑	↓	↓
MP	↓	↓	↑

Source: Authors' compilations.

Note: AD refers to aggregate demand, AS to aggregate supply and MP to monetary policy. We impose the restrictions as ≥ and ≤. The restrictions are binding for one quarter after impact.

As more restrictions typically lead to stronger identification in a sign restriction framework (see Fry and Pagan, 2011), we not only identify the monetary policy shock, but simultaneously also identify an aggregate demand and an aggregate supply shock. The orthonormal rotation matrices establish a relationship between our reduced-form GVAR and the underlying structural representation of the model. In light of the sign restrictions described in table 3, our structural GVAR model shares features commonly observed in the standard dynamic stochastic general equilibrium models usually employed by central banks and policy institutions. In particular, the contractionary monetary policy shock is defined as an unexpected increase in the U.S. short-term interest rate that is assumed to trigger a decline in output and inflation in the U.S.A. at least until the first quarter after impact.

3.5 Impulse response analysis

Chart 5 depicts the responses of key macroeconomic variables to a contractionary U.S. monetary policy shock for Poland, Turkey and the CESEE average.[9] Interestingly, in several cases, the responses in Turkey deviate from those in Poland and the CESEE average. This heterogeneity can most likely be explained by the fact that in the observation period, Turkey was characterized by stronger economic volatility (recall the 2001 crisis) than the CESEE region on average (recall the introduction and section 1).

Examining the response of output, we see a pronounced decline in real GDP that has a persistent nature (corroborating the findings of Feldkircher and Huber, 2016, and Willems, 2013). Compared to other CESEE economies, Turkey displays the strongest GDP drop on impact. Output contracts by –0.7% and then recovers somewhat until the end of the first year after the shock but continues to decline at a steady rate of about –0.3% (statistically significant until three years after the shock). CESEE countries reach their minimums of output declines on average within the first year after the shock (–0.4%) and are then able to relieve the pressure only slowly.[10]

[9] *CESEE responses are shown as simple unweighted averages across the 12 CESEE economies. Purchasing power parity (PPP) weighted responses would be an alternative, but they limit the responses to those of dominant countries (such as Poland or Turkey).*

[10] *The finding of a stronger GDP decline in Turkey compared to Poland in response to a contractionary monetary policy shock in the U.S.A. is consistent with IMF findings (IMF, 2014).*

The significant output decline is also mirrored by the developments of other variables in the model. Consumer prices decline, with this effect being statistically significant only in Turkey. A pronounced hike in the short-term interest rate that lasts at least for two quarters after the shock is only briefly able to reverse the price decline in Turkey. Inflation and interest rates in Poland, on the other hand, do not show a statistically significant pattern.

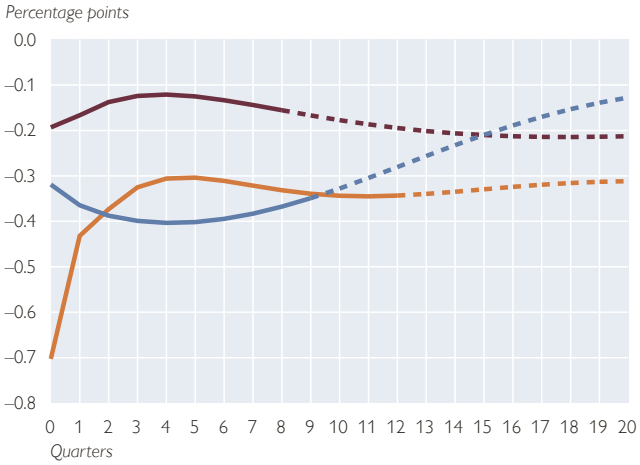
The contractionary U.S. monetary policy shock also results in a marked real depreciation of CESEE currencies vis-à-vis the U.S. dollar, which remains persistent in the CESEE region on average at least up to one-and-a-half years after the shock. Having the nominal depreciation figures of chart 1 in mind, interestingly, the real depreciation in Turkey is apparently not as strong and persistent as in Poland. In line with these currency depreciations, trade balances improve, though we are able to identify a statistically significant improvement only for Turkey up to one quarter after the shock.

Finally, we see a strong medium-run deterioration of the financial account in Turkey, reflecting capital outflows (or a reduction in capital inflows) right after the monetary contraction in the U.S.A. The mentioned initial hike in the Turkish short-term interest rate might be a reflection of domestic policymakers' attempts to contain these capital outflows. A short-run deterioration in the financial account can also be observed in Poland, though it is less pronounced than in Turkey.

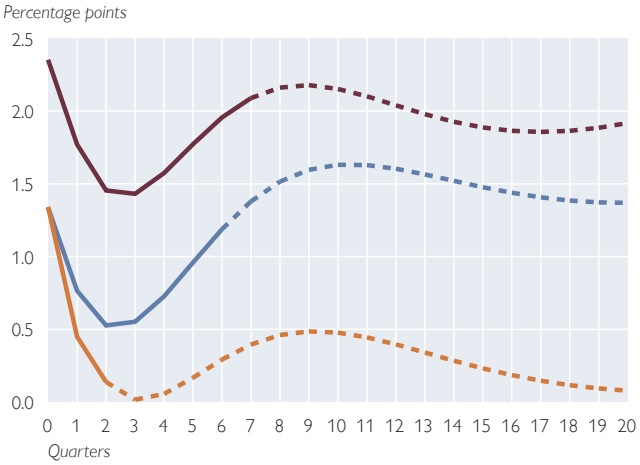
Chart 5

Response to a (one standard deviation) contractionary U.S. monetary policy shock

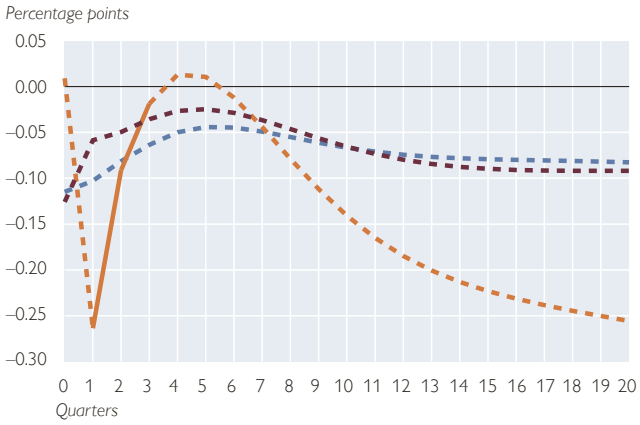
Median response of real GDP



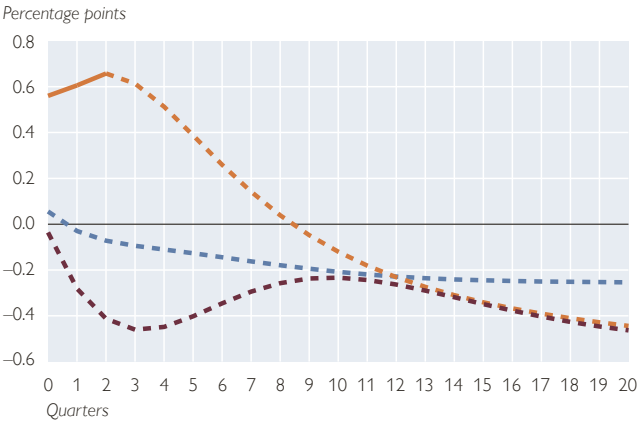
Median response of real exchange rate (vis-à-vis USD)



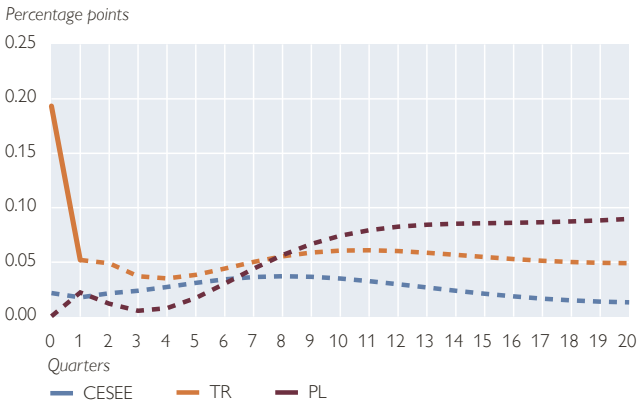
Median response of consumer price inflation



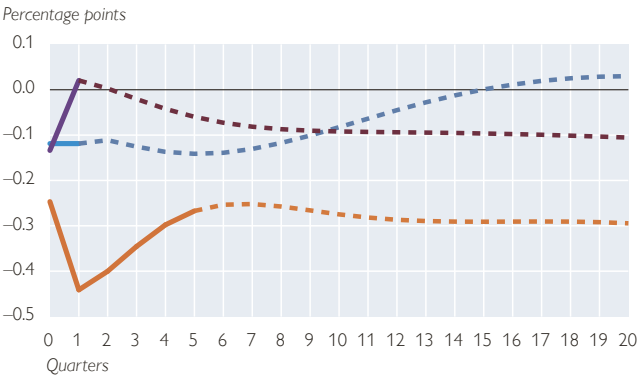
Median response of short-term interest rate



Median response of trade balance



Median response of financial account



Source: Authors' calculations.

Note: Median impulse response in percentage points, based on 1,000 draws that are randomly extracted from the posterior of the impulse responses. Solid lines represent statistically significant responses in the sense that zero does not lie between the 25th and 75th percentile, whereas dotted lines indicate statistically insignificant responses. t=0 denotes the quarter in which the shock occurs. CESEE figures are based on unweighted averages across the 12 CESEE economies.

4 Concluding remarks

In this paper, we have tried to describe Turkey's and Poland's relative economic performance in the situation of two recent global shocks: first, the global economic crisis in 2008/2009 and second, the Fed's tapering announcement in May 2013. Our description places an emphasis on the underlying macrofinancial vulnerabilities.

While both Turkey and Poland weathered the 2008/2009 crisis comparatively well, macrofinancial indicators responded fairly strongly to the Fed's tapering announcement. Among other things, marked currency depreciation, reversals in capital flows and a slowdown in cross-border bank lending challenged policymakers in the region.

To improve our understanding of the actual responses of domestic macroeconomic variables to a global-scale external shock, we investigate the international transmission mechanism of a contractionary U.S. monetary policy shock by means of a Bayesian GVAR model. This multicountry model provides a more coherent picture of the underlying transmission channels by taking cross-country effects seriously. We investigate the economic responses in the CESEE region, with a specific focus on Turkey and Poland, to a U.S.-based contractionary monetary policy shock.

Our simulation results suggest that both Turkey and Poland tend to exhibit significant short- to medium-run responses to an unexpected increase in the short-term interest rate in the U.S.A., while long-run responses tend to become insignificant after a few quarters for most variables under scrutiny. Taking a regionally comparative stance unveils somewhat stronger responses for Turkey than for Poland or the CESEE average, signaling the structurally different nature of the Turkish economy. More specifically, as a traditional emerging market economy outside the EU, Turkey is less interlinked with the euro area, suggesting a business cycle decoupling from Poland and the other CESEE countries. In addition, the higher volatility of the macroeconomic fundamentals in Turkey might translate into different risk profiles, leading to more pronounced responses. Moreover, the strong trade ties between Turkey and the U.S.A. suggest a stronger transmission mechanism for U.S.-based shocks. Overall, for most of the studied macroeconomic variables, the identified responses mimic the actual developments that we have observed since early 2013. It remains to be seen whether potential further interest rate hikes in the U.S.A. will still lead to pronounced short-term macroeconomic responses in CESEE or whether financial markets have already largely priced in such increases.

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CESEE-related abstracts from other OeNB publications

The abstracts below alert readers to studies on CESEE topics in other OeNB publications. Please see *www.oenb.at* for the full-length versions of these studies.

The Russian banking sector – heightened risks in a difficult environment

Stephan Barisitz

The Russian banking sector has passed from excessive retail credit growth (up to early 2014) to a general contraction of credit (2015). This recent decrease is to a certain degree attributable to Western sanctions cutting off leading banks from cheap refinancing, but most notably to the steep fall of the oil price. The latter caused the ruble to plunge and pushed the Russian economy into recession. Temporary financial instability was reined in by the Bank of Russia's sharp increase of the key rate (largely reversed recently) and by expanding deposit insurance coverage. Liquidity injections, foreign currency repurchase agreements, and the recapitalization of a number of systemic banks also helped. Moreover, a degree of regulatory forbearance was introduced. As the economy shrinks, nonperforming loans are inevitably rising and profitability is declining. The banking sector is primarily exposed to credit risk, followed by liquidity and exchange rate risk. Connected lending is a structural problem now being finally tackled. Shock absorbers have eroded but still provide leeway: deposits are sizable and depositor confidence seems to have returned. Russian banks have a positive net external creditor position. Public debt is low and the country – notwithstanding terms-of-trade losses – continues to boast current account surpluses. Foreign currency reserves – after declining – have restabilized and remain substantial.

Published in *Financial Stability Report 30*.

Event wrap-ups and miscellaneous

10th Conference of the South-East European Monetary History Network (SEEMHN)

Compiled by
Clemens Jobst and
Thomas Scheiber

The general theme of the conference, which took place in Vienna on October 1 and 2, 2015, was "Financial development and economic growth in South-East Europe." The conference was jointly organized by the Department for Economic and Social History of the University of Vienna and the Oesterreichische Nationalbank (OeNB). The purpose of the SEEMHN conference was to gather scholars researching the historical development of banks, central banks and financial markets as well as the economic development of countries in Southeastern Europe to gain new insights into the growth-finance nexus.

In his keynote address, Professor *Philipp Ther* (Department for East European History, University of Vienna) focused on economic reforms in Eastern Europe in the 1990s within a wider context of political and economic developments in the advanced European economies. According to Ther, the success of reforms has remained mixed and today economic inequality between rich and poor regions within Eastern European countries is as big as between Eastern and Western European economies.

The second keynote lecture also addressed regional inequalities with respect to the questions of why and how countries industrialized. Professor *Max-Stephan Schulze* (London School of Economics and Political Science) presented new evidence on regional growth in the late 19th century Habsburg Empire. Here again, inter-regional differences in levels and growth of GDP per capita were far larger than previously assumed; Eastern regions economically lagged behind Western regions mainly due to the absence of nearby consumer markets.

The first session, chaired by *Julia Wörz* (Foreign Research Division, OeNB) dealt with the evolution of financial markets in the Balkans. *İrfan Kokdaş* (Izmir Kâtip Çelebi University) looked at the connections between the formation of large agricultural estates and the concentration of tax farms in the hands of a few families as well as the workings of financial markets in the Ottoman Balkans. By way of example, he focused on three towns during the period from 1690 to 1850. *John R. Lampe* (University of Maryland) presented a comprehensive overview of the evolution of the banking and financial system in Southeastern Europe since independence. Moreover, he pointed out a lot of open questions – some of which will be addressed in his contribution to the forthcoming six-volume publication on the history of Southeastern Europe of the University of Regensburg.

The second session, chaired by Professor *Carsten Burhop* (Department for Economic and Social History, University of Vienna), explored the origins and propagation mechanisms of financial crises, taking 1931 as an example. *Stefan Nikolic* (University of York) analyzed the propagation of the 1931 financial crisis from the core to the periphery, using weekly data looking for "fast and furious" effects as posited by Kaminsky. Nikolic provided evidence that the 1931 crisis was also a sovereign debt crisis, which forced Balkan countries to introduce exchange controls. While the collapse of *Creditanstalt* had no significant impact on the long-term government bond yields of Balkan countries, significant effects can be detected for the ensuing German crisis and in particular for the weeks after the U.K. had left the gold standard. *Flora Macher* (London School of Economics and Political Science) presented evidence that the Hungarian crisis in 1931 originated in the banking system and was caused by a tight monetary policy after the currency

crisis in 1928, an agricultural crisis in 1930 and a fiscal policy that promoted the provision of state-guaranteed loans. Ongoing research will try to explain what triggered the crisis and how it interacted with the crisis which hit neighboring Austria at the same time.

The final session, chaired by *Martin Ivanov* (Bulgarian Academy of Science and Sofia University), focused on the interplay between government debt and financial markets and the importance of sound government finances for the sustainability of fixed exchange rate regimes. The evidence presented by *Matthias Morys* (University of York) showed that the Greek debt crisis has to be seen within the context of the Balkans' long history of reliance on external financial support as well as international financial supervision. *Michael Pammer,* Professor at the Johannes Kepler University in Linz, scrutinized the determinants of the higher risk premium on Hungarian state bonds as compared to Cisleithanian state bonds. For the period 1881–1914 the relevant factors determining the spread in yields were strictly fiscal ones, he argued, particularly the difference in the ability of both states to finance their interest payments out of their tax, fee and property revenues. Political crises affected relative price movements from time to time, but only in the short run and to a negligible extent compared with ordinary fluctuations. 19th century Balkan economies needed to attract foreign capital for large-scale investment in various infrastructure and other key sectors, which led to the accumulation of enormous public foreign debt. Professor *Nikolay Nenovsky* and *Jacques-Marie Vaslin* (both CRIISEA, University of Picardie Jules Verne, Amiens) elaborated on the incompatibilities of shadowing the Latin Monetary Union and public debt dynamics at the Balkans periphery.

The details of the program as well as a link to selected presentations can be found at: https://www.oenb.at/en/Monetary-Policy/Research/conferences.html

Box 1

What is the South-East European Monetary History Network (SEEMHN)?

The South-East European Monetary History Network (SEEMHN) was established in 2006 and brings together financial and monetary historians, economists and statisticians working on South-East Europe. Its main objective is to increase the visibility of the region in historical research and promote research on the region as an integral part of European history.

An important outcome of the long-standing cooperation of the central banks involved in the SEEMHN has been the compilation of datasets of monetary and financial variables for seven South-East European countries, including Austria, covering the period from the 19th century to World War II. This data volume was published in December 2014 and is available for free download on the websites of the central banks involved:

https://www.oenb.at/en/Publications/Economics/south-east-european-monetary-history-network-data-volume.html

13th ESCB Emerging Markets Workshop

Compiled by
Florian Huber

On November 5 and 6, 2015, the Oesterreichische Nationalbank (OeNB) hosted the 13th ESCB workshop on emerging markets. Since its inception in the early 2000s, this workshop series has served as a forum for researchers working on macroeconomic and financial issues related to emerging economies – a forum which allows them to present and discuss analytical work undertaken in the ESCB on these topics. From a large number of high-quality submissions, a two-day program was compiled, covering diverse topics such as monetary and fiscal policy, macrofinancial stability and economic growth. Given the high policy relevance of the papers presented at the workshop, lively discussions ensued in all sessions, providing useful insights and guidance for researchers and policymakers alike.

In her introductory statement, *Doris Ritzberger-Grünwald,* Director of the OeNB's Economic Analysis and Research Department, emphasized the increasing relevance of emerging markets for the global economy in general and for Austria in particular. Emerging and developing countries now account for more than 57% of global GDP in terms of purchasing power parities, which symbolizes a shift in the balance of economic power from advanced to emerging economies. For Austria, deepening economic integration with emerging economies in Central, Eastern and Southeastern Europe (CESEE) has historically paid off in terms of GDP growth and job creation. However, intensifying integration and the rise of emerging economies on the global scene has also led to new challenges for both advanced and emerging economies. The recent euro area recession has once more brought up the question to what extent economic shocks in developed countries trigger downturns in emerging markets.

Keynote lecture I: "Back to normal. Which normal for Latin America?"

Enrique Alberola, Chief Representative of the Americas Office of the Bank for International Settlements (BIS), talked about the structural shift of Latin American economies toward a more resilient and sustainable growth path. The overall economic expansion that lasted over several decades was strongly driven by internal and external factors, namely the commodity boom, wide access to external financing and increases in structural demand coupled with expansionary policies. However, five consecutive years of decreasing growth and a continuous decline in growth expectations coupled with deteriorating external positions have led to a more pessimistic picture for the region as a whole. Alberola provided several possible reasons for recent crises in Latin America. For instance, he identified loose financial conditions in advanced economies as a key driver behind recent asset price booms and sharp increases in credit. In addition, international investors' sentiment has proved to be another important determinant of economic instability in the region. Recent capital flow data display stark differences between balance of payments figures and EPFR data, and need to be monitored closely. Since commodity prices and capital flows tend to be of prime importance for several countries in Latin America, Alberola has attempted to tackle shortcomings of existing measures of the output gap in a recent research paper: together with his coauthors he has constructed a commodity- and capital flow-adjusted measure of the output gap. From 2003 onward, this augmented measure yields lower estimates of the output gap, suggesting that existing methods tend to underestimate the current output gap. Alberola also mentioned that most countries in the region now

possess stronger lines of defense than in the past and stressed the challenge of a generally lower level of economic growth and its impact on financial stability.

The discussion that followed centered on whether a rise of the middle class increases the demand for public services, which in turn could lead to more pressure on public balances. Moreover, while the presentation of Alberola showed that a few Latin American countries including Brazil are currently especially vulnerable, there are also counterexamples like Mexico, where reform-oriented policies in the last years have led to a higher resilience with regard to macrofinancial shocks.

Session I: Monetary policy and capital flows (part I)

The first session on monetary policy and capital flows tackled two main questions. First, *Riikka Nuutilainen* (BOFIT) talked about the relationship between reserve requirements and the bank lending channel in China. The study she presented was the result of joint work with *Zuzana Fungáčová* (BOFIT, currently European Central Bank) and *Laurent Weill* (BOFIT). In the study, panel fixed-effect estimations are used to model the change in banks' loan supply as a function of a monetary policy indicator, GDP growth and a set of bank-specific characteristics for the time period between 2004 and 2013. To approximate the monetary policy stance of the People's Bank of China, the reserve requirements ratio is included in the model. The main findings are that a tightening of reserve requirements leads to a significant decline in loan growth, underpinning the effectiveness of reserve requirements as a monetary policy instrument. Other determinants like bank-specific characteristics tend to play a minor role in explaining changes in loan growth. *Dubravko Mihaljek* (BIS), who acted as the discussant for this session, emphasized that the issue with credit in China is not so much its growth but its allocation; perhaps, too little credit is granted to productive and innovative firms and SMEs in general.

Second, *Markus Eller* (OeNB) presented the main findings of his paper, coauthored with *Helene Schuberth* and *Florian Huber* (both OeNB), investigating the global drivers of capital flows. In light of the high volatility of capital flows seen in recent years, the authors propose a dynamic factor model with stochastic volatility that incorporates several stylized features of the time series. Assuming that capital flows across the globe can be explained through different unobservable factors, Eller and his coauthors used a straightforward decomposition to shed some light on the importance of macroeconomic, financial and trade factors in explaining variation in capital flows and whether the importance of the different determinants changes over time. The main findings of the paper are that capital flows are strongly driven by global financial factors that reflect international monetary policy, long-term interest rates, equity prices and credit, and there is ample evidence that the explanatory power of global financial factors has increased since the 2008–09 global financial crisis. The discussant raised several important points, most notably that capital flows could also be driven by commodity prices, which have not been included in the proposed model, however.

The general discussion highlighted how difficult it is to improve international policy coordination, especially in situations where business cycles diverge across countries. The pivotal role of the U.S. economy for global macroeconomic factors was also highlighted.

Session II: Monetary policy and capital flows (part II)

In the first talk of the session, *Ignacio Hernando* (Banco de España) discussed the impact of unconventional monetary policy measures adopted by the Fed and their impact on Latin America. His presentation was based on joint work with *Fructuoso Borrallo* and *Javier Vallés* (both Banco de España). The three economists used an event study approach to provide some insights into the dynamic effects of the U.S. Federal Reserve System's unconventional monetary policy on sovereign bond yields, exchange rates, capital flows and equity prices. The main findings were that the announcement of the first large-scale asset purchase program significantly reduced government bond yields by around 22 basis points. Moreover, Latin American exchange rates depreciated significantly vis-à-vis the U.S. dollar around the tapering period. *Martin Suster* (Národná banka Slovenska), the discussant, emphasized the relevance of the research question and mentioned that the findings were not driven by trade or geographical location but by domestic economic conditions of the Latin American economies.

Julio Ramos-Tallada (Banque de France) talked about the relationship between banking risk, monetary policy shocks and the credit channel in Brazil. He used a large dataset of financial statements of banks available for the period from 1995 to 2012 to investigate the main determinants of the bank lending channel in Brazil. A micro-based identification approach for monetary policy shocks is used to disentangle demand from supply shifts in credit. Under this identification scheme, lending supply and short-term interest rates as a proxy for monetary policy tend to be negatively correlated over the full sample. Interestingly, another key finding suggests that the relationship between credit supply and monetary policy is not related to bank characteristics. The discussant emphasized the need to evaluate the relationship between the bank lending channel and other monetary policy transmission channels.

In the general discussion it was mentioned that the Fed was increasingly adopting a forward-looking stance with respect to global economic developments. Moreover, in recent years the effectiveness of monetary policy transmission across the globe has been challenged by issues related to the zero lower bound of nominal interest rates.

Session III: Fiscal and financial stability (part I)

In the first afternoon session, the focus shifted from monetary policy issues to fiscal and financial stability. *Markus Jorra* (Deutsche Bundesbank) addressed the question of how Argentina's (partial) sovereign default in 2014 influenced GDP growth, FDI inflows and bilateral trade. He proposed a novel approach that uses a synthetic control estimator, where a hypothetical no-default scenario is constructed and compared with the actual outcome of different macroeconomic quantities. Jorra's empirical analysis presents little evidence for causal effects of sovereign defaults. More specifically, the effects on GDP and industrial production appear to be rather muted, suggesting that the 2014 sovereign default was more or less cost-free. The discussant, *Mariya Hake* (OeNB), mentioned that the muted effects on GDP and industrial production stem from the fact that the selective default could be seen as a special case.

Flavia Corneli (Banca d'Italia) investigated the medium- and long-run implications of financial integration without financial development. Within a neoclassical

growth model, Corneli showed that in the medium term, financial integration reduces capital accumulation in developing economies due to a high risk premium of production activities. By contrast, the long-run effect revealed that within emerging economies, integration brings higher capital than a so-called autarky steady state. For advanced economies, the model shows that the effect is reversed. In the medium term, integration yields higher growth rates of capital while the long-run effect points toward a reduction in capital growth relative to the autarky solution. The discussant mentioned that the paper should be better integrated with existing literature that draws different conclusions regarding the relationship between financial integration and financial development.

In this session, the general discussion centered on the measures adopted to investigate the economic consequences of a sovereign default. For instance, financial variables like refinancing costs could have changed significantly in response to a sovereign default. In addition, the specific definition of financial market integration, i.e. that sovereigns are able to issue risk-free bonds, was also questioned.

Session IV: Fiscal and financial stability (part II)

The final session of the first day was mainly concerned with the determinants of bank lending activities in emerging market economies, most notably the Central and Eastern European (CEE) region. *Ádám Banai* (Magyar Nemzeti Bank) presented a paper, coauthored with *Judit Temesvary* (Hamilton College and Cornell University), which investigates the main driving forces behind foreign bank lending activities in CEE. In light of the particularly pronounced heterogeneity of lending activities across the region during the crisis, a better understanding of cross-country differences is helpful for policymakers. Within a panel model, the authors analyze the lending activities of 25 banking groups in the CEE region, operating in 11 different economies. They find several important determinants of foreign lending activities. For instance, nonperforming loan ratios exhibit a significantly negative effect on lending activity. In addition, the attitude toward risk-taking on the part of the parent group and the risk level of a given host country prove to be important for explaining lending activities. Moreover, external indebtedness, funding capabilities of the parent group and the capitalization of subsidiaries all tend to influence lending in a statistical significant fashion. *Katharina Steiner* (OeNB) discussed this session. She mentioned that the seemingly ad-hoc selection of various indicators should be better rationalized.

The second paper of the session, presented by *Arben Mustafa* (Central Bank of the Republic of Kosovo), dealt with the relationship between banking sector competition and banks' net interest margins in the CEE countries. In his paper, Mustafa analyzes the impact of three popular measures of banking sector competition on interest margins. The main conclusions are that banking sector competition exerts a statistically significant negative effect on interest margins, with all three measures of competition yielding similar results. In addition, the analysis suggests that the ability of less efficient banks to translate inefficiencies into higher interest margins vanishes in more competitive markets. The discussant suggested looking at the role of funding costs and loan pricing as a possible extension of the analysis.

Ramona Jimborean (Banque de France) shed some light on the possible drivers of private nonfinancial sector (NFS) borrowing in emerging market economies. The relevance of the research question is underlined by the fact that many crises in

emerging markets have been preceded by rapid leverage growth. To gain a better understanding of the driving forces behind NFS borrowing, Jimborean investigated external and domestic determinants on three different measures of private NFS borrowing. Her analysis suggests that credit demand is positively related to borrowing from domestic banks and from all sectors and negatively related to the nominal exchange rate. Moreover, contractionary domestic monetary policy reduces private NFS borrowing from domestic banks. In addition, global funding conditions, global GDP growth and U.S. monetary policy also influence some of the measures considered. The discussant provided several suggestions: most importantly, endogeneity issues between interest margins, indicators of competition and non-interest income might be solved by lagging the explanatory variables.

Keynote lecture II: "Structural external imbalances as a constraint on convergence in the European economy"

In his keynote speech, *Michael Landesmann,* Director of Research at the Vienna Institute for International Economic Studies (wiiw), focused mainly on the CESEE region and explained that external imbalances are not per se a problem in the context of economic convergence – which is an important element of the EU integration process – but that they may threaten convergence when they become unsustainable. He referred to these structural imbalances as the Achilles' heel of catching-up. Persistently low exporting capacities, de-industrialization trends in the aftermath of the transformation, lacking focus on the tradeable sector and agglomeration tendencies have contributed to what has now become manifest in a long-run structural problem. He emphasized that the role of the real exchange rate in this context is not always well understood. Often, real appreciation is the consequence of previous good trade performance rather than a factor driving future trade performance. Taking a forward-looking stance, Landesmann identified the following areas as important determinants of competitiveness: moving toward products with higher income elasticity, building sufficient export capacity and diversifying the export structure to reduce vulnerabilities.

In the discussion, he highlighted the Central European countries – Poland, the Czech Republic, Slovakia, Hungary and Slovenia – as good examples of re-industrialization. He also emphasized that tradeable services need manufacturing exports as a carrier and mentioned the conflict between industrial support and EU competition policy, which implies an insufficient implementation of industrial policy, too much focus on incumbent players and too little effort directed at nurturing the entry of new players in each country.

Session V: Economic growth and international economics

In the following session, the question raised by the previous keynote lecture – whether the exchange rate can be used as a policy variable or whether it is rather an outcome – was taken up by *Maurizio Michael Habib* (European Central Bank, ECB) based on joint work with *Elitza Mileva* (Fordham University and World Bank) and *Livio Stracca* (ECB). Using a difference-in-difference model and information on global capital flows, official exchange rate regimes and foreign exchange reserves as instruments for the real exchange rate, Habib found evidence that real exchange rate depreciation may influence economic growth positively. He also found evidence for an asymmetric effect between appreciation and depreciation, a

larger effect for developing countries and currency peggers, and non-linearities depending on the foreign currency position. The discussant, *Paul Pichler* (OeNB), stressed the unclear differentiation between depreciation and undervaluation.

The second speaker, *Simona Manu* (ECB) presented her paper, coauthored with *David Lodge* and *Ioannis Grintzalis* (both also ECB), on the influence of the global financial cycle and domestic credit gap on output fluctuations. According to the authors' results, the global financial cycle is an important factor in shaping the cyclical stance of emerging economies, exerting a negative influence up until 2006 and then again since 2012. The discussant stressed the value of this approach as it seems to be comparatively robust against end-of-sample instability and data revisions and he advocated using different output gap measures based on the policy objective (monetary, fiscal, macroprudential). His criticism that the measurement of the credit gap as a simple deviation from long-term growth is too simplistic was rejected by Manu, who argued that this simplicity is useful given the high volatility of credit growth in emerging economies.

Risto Herrala (Bank of Finland) presented an empirical analysis coauthored with *Rima Turk Ariss* (Lebanese American University) investigating the effects of political instability on credit availability and financing conditions. Based on survey data from 860 firms in the Middle East and North Africa (MENA) region the authors found weak direct effects from political instability on capital accumulation, but evidence for indirect effects through credit limits. In the discussion, participants debated the role of international banks and other external sources of financing, as well as whether constraints arise from the supply side or rather the demand side.

The general discussion centered on possible threats to further capital liberalization and the often too one-sided lessons that were drawn from previous liberalization episodes (i.e. the accumulation of too many reserves by emerging economies following the crises in Latin America in the 1990s). The importance of distinguishing "good" from "bad" credit booms was emphasized and the extent to which policymakers should lean against the wind was discussed.

Keynote lecture III: "If China decelerates, rebalances and liberalizes, what happens to the rest of us?"

Iikka Korhonen, Director of the Bank of Finland Institute for Economies in Transition (BOFIT), delivered the third keynote lecture. According to Korhonen, the Chinese economy is currently characterized by three broad trends: Economic growth is trending downward (2015: around 7%, 2016–17: 6% expected), the structure of growth is changing (moving from investment- to consumption-driven growth), and the liberalization of financial markets is progressing (both in terms of domestic financial intermediation and cross-border capital movements), regardless of recent market turbulence. The implications of these developments are (among others): Producers of raw materials as well as of machinery and equipment will continue to suffer from less demand and/or lower prices, the renminbi will be increasingly widely used in international transactions, and capital flows from and to China will rise.

In the ensuing discussion, Korhonen pointed out that he expected Chinese reserves (currently amounting to USD 3.4 trillion) to continue to decline in a more liberalized environment. Looking at the medium term, economic growth in

China may settle at 4% to 5% per annum, influenced by demographic factors among others.

Session VI: Further selected topics related to emerging market economies

The first speaker of the final session of the workshop, *Peter Tóth* (Národná banka Slovenska), talked about nowcasting GDP growth in selected Central, Eastern and Southeastern European countries. In his presentation, which was based on a paper coauthored with Martin Feldkircher, Florian Huber, Julia Wörz, Josef Schreiner (all OeNB) and Marcel Tirpák (ECB), he again emphasized the importance of taking into account the heterogeneity of the CESEE countries when specifying appropriate nowcasting models. His findings suggest that small-scale dynamic factor models and models based on a bridge equation tend to outperform simple benchmark models for all countries considered. In the final session, *Christian Buelens* (National Bank of Belgium) acted as a discussant. Buelens emphasized the sensitivity of the findings with respect to the estimation and verification period.

Mehmet Fatih Ulu (Central Bank of the Republic of Turkey) finally tackled the question how access to new foreign intermediate goods is related to product innovation within a structural modeling framework. His main results, based on data for India, show that the import of intermediate goods increases the revenues for each product created and, through knowledge spillovers, the probability that domestic firms introduce new products is increased significantly. The discussant highlighted the relevance of the calibration exercise but stressed that calibration should not be interpreted as a test for causality.

IMF presents its fall report on CESEE Regional Economic Issues (REI): "Reconciling Fiscal Consolidation and Growth"

Summarized by Christina Lerner[1]

On November 13, 2015, the IMF's Regional Economic Issues (REI) report for Central, Eastern and Southeastern Europe (CESEE)[2] was presented at the Oesterreichische Nationalbank (OeNB). This event also marked the official launch of the IMF's fall 2015 REI, as the report was published on the IMF's website at the same time and was then presented in a series of seminars in several European cities.

Jörg Decressin, Deputy Director of the IMF's European Department, analyzed the risks and policy priorities for CESEE countries, while *Johannes Wiegand*, Deputy Chief of the Emerging Economies Division in the IMF's European Department, focused on reconciling fiscal consolidation and growth in CESEE. *Anna Ilyina*, Chief of the Emerging Economies Division in the IMF's European Department, then provided further insights into the economic performance of CESEE countries.

The presentations were followed by a discussion during which journalists, OeNB economists and experts from various economic institutions and commercial banks raised additional topics and exchanged their views.

Opening remarks

The event was opened by *Ewald Nowotny*, Governor of the OeNB, and chaired by *Franz Nauschnigg*, Head of the OeNB's European Affairs and International Financial Organizations Division.

In his opening remarks, Governor Nowotny underlined the importance of CESEE as a region of strategic interest for the OeNB and stressed the importance of Vienna as a hub for CESEE know-how, as regards economic, administrative, commercial and legal linkages to the CESEE region. He furthermore mentioned that CESEE cannot be considered as a homogenous region, as, for instance, the Southeastern European (SEE) countries are facing a bigger fiscal consolidation challenge than the Central and Eastern European (CEE) countries.

Positive growth in 2016

At the beginning of his presentation, Jörg Decressin explained that the 2015 GDP growth forecast for the CESEE region as a whole remained broadly unchanged compared to the REI publication of May 2015, but that there were notable shifts in the contribution of countries: Growth was revised up for CEE and SEE, kept unchanged for Russia and Turkey, and revised down in the Baltic states and other CIS countries. In other words, most CESEE countries are growing at a healthy pace, whereas Russia and other CIS economies are facing significant economic challenges and are in recession.

However, stabilization is expected for 2016, as the Russian economy adjusts to low oil prices and sanctions: The CESEE region as a whole is set to contract by

[1] *Oesterreichische Nationalbank, European Affairs and International Financial Organizations Division, christina.lerner@oenb.at.*

[2] *The IMF's semiannual CESEE REI report assesses the macroeconomic outlook for CESEE in light of the latest global economic and financial developments and covers more than twenty countries, including Central and Eastern Europe (CEE), Southeastern Europe (SEE), the Baltic region, the Commonwealth of Independent States (CIS), and Turkey.*

0.6% in 2015, but to expand by 1.3% in 2016. The growth divergence among CESEE countries is mainly driven by different levels of domestic demand, and partly spurred by payouts from EU structural and cohesion funds. Also the rates of inflation differ among the countries, with very low inflation prevailing in many economies.

New risks to the outlook

The balance of risks has shifted to the downside, as China's slowdown and the ongoing refugee crisis represent new risks. Jörg Decressin highlighted that it now depends on European politics whether the refugee crisis will translate into a downside or rather an upside risk, as in the medium term the associated spending increases might boost growth.

According to Decressin, policy priorities depend on how far along the economies are in their postcrisis adjustment and on their exposure to external risks: Where the recovery is well advanced, the focus should be on implementing structural reforms and rebuilding fiscal buffers, while in case of economies in recession, fiscal policy should aim at supporting domestic demand, and monetary policy should address inflation where it is high. In addition, countries vulnerable to external shocks (as for instance Turkey) need to be prepared to deal with market pressures by using exchange rate flexibility as a shock absorber alongside macroprudential policies designed to contain the buildup of financial sector risks.

Growth-friendly consolidation and budget reform necessary

Johannes Wiegand then elaborated on fiscal consolidation in CESEE and its long-term impact on growth: As large fiscal challenges remain in CESEE, lifting potential growth is a key medium-term goal for CESEE countries.

Wiegand highlighted that the growth friendliness of government budgets remains a key policy challenge: Most CESEE economies had entered the global financial crisis with growth-unfriendly budgetary structures: on the spending side, budgets were characterized by high public consumption and large unproductive transfers, whereas on the revenue side, they showed a disproportionate reliance on labor taxes, notably social security contributions.

Hence, in the aftermath of the global financial crisis, economies in CESEE came under pressure to correct external and internal imbalances as the global financial crisis forced sharp, procyclical fiscal adjustments in many CESEE countries.

As regards budget structures, Wiegand stated that intra-CESEE differences are modest and that these structures resemble those of advanced Europe more closely than those evident in other emerging economies. However, since the global financial crisis, the budget structure has changed markedly and budget balances have generally improved: governments came under pressure to consolidate public finances, which tended to have a positive impact on the quality of budgets. Growth-friendly consolidation and budget reform is critical to strengthen long-term growth prospects. However, fiscal consolidation has not yet run its course, with sizeable adjustment needs remaining, especially in SEE, where the budgetary situation has deteriorated. In contrast, the expenditure structure of the Baltic and CEE countries has improved, as they have cut current spending.

According to Wiegand, key policy priorities should include reducing unproductive transfers while protecting productive spending on health, education and

public infrastructure, reforming public employment, leveraging access to EU funds and shifting taxation from income to consumption: Growth friendly consolidation and budget reform is critical to strengthen long-term growth prospects in CESEE.

The IMF's CESEE REI report can be downloaded at:
http://www.imf.org/external/pubs/ft/reo/2015/eur/eng/pdf/rei1115.pdf

Kaushik Basu on "Globalization and the art of designing policy"

Summary of the 20th Global Economy Lecture

Compiled by
Julia Wörz

The 2015 Global Economy Lecture[1] was delivered by *Kaushik Basu*, Chief Economist and Senior Vice President of the World Bank. In his presentation at the OeNB on November 16, 2015, he illustrated how globalization influences national economic policies and how economic events in one country can have repercussions in other countries, which may then feed back into the country of origin again. As an example he mentioned the downgrade of the United States' credit rating by Standard & Poor's on August 5, 2011. Contrary to expectations and to textbook economic theory, this led to a strong demand for U.S. T-bills as Chinese investors were aiming at preventing losses on their sizable U.S. dollar holdings, which in turn created depreciation pressure on the Indian rupee. More generally, he emphasized the discrepancy between national policy needs and the potential global effects of national policies and argued in favor of so-called "reasoned intuition." In his view, policymakers need to take into account the effects of their national policies abroad and to factor in foreign policy and market reactions when designing their policies. He also pointed toward multiple equilibria which emerge in case of self-fulfilling prophecies and herding behavior. Using the example of credit markets, Basu illustrated that a small change in price signals (i.e. an interest rate rise) may induce a large and sudden fall in credit supply if the economy moves from a high-credit equilibrium to a low-credit equilibrium.

Turning to the euro area, which Basu called the biggest globalization experiment in history, he observed that the flaws in the construction of the euro area were not apparent initially and became evident only through the global economic crisis, when borrowing costs within the euro area suddenly started to diverge. In his view, the impossibility of sharing the public debt burden in the euro area is one of the factors hampering the smooth functioning of credit markets in the euro area and is thus an issue that needs to be revisited. In the remainder of his talk, he focused on emerging economies. He identified the following policy needs in a globalized world: better global economic governance through institutions like the G-20, the IMF and the BIS; strategic thinking in constructing domestic policies that takes into account foreign policy reactions; openness to new ideas.

In the discussion he recalled that trust is an important element for the functioning of societies. However, in a globalized world, we need regulation to guide economies toward the desired behavior. Regulation can also help to rule out multiple (and undesired) equilibria. Finally, he compared the euro are to India, where many different states share a common currency and emphasized that one economy needs one currency. Although today's world is becoming a global economy, he does not consider all countries, and especially not all emerging economies, to be ready for sharing a world currency.

[1] *The Global Economy Lecture is an annual event organized jointly by the Oesterreichische Nationalbank (OeNB) and The Vienna Institute for International Economic Studies (wiiw).*

Notes

Studies published in Focus on European Economic Integration in 2015

For more information, see *www.oenb.at*.

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Periodical publications

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Annual Report (Sustainability Report) — English | annually

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http://www.oenb.at/en/Publications/Oesterreichische-Nationalbank/Annual-Report.html

Konjunktur aktuell — German | seven times a year

This online publication provides a concise assessment of current cyclical and financial developments in the global economy, the euro area, Central, Eastern and Southeastern European countries, and in Austria. The quarterly releases (March, June, September and December) also include short analyses of economic and monetary policy issues.

http://www.oenb.at/Publikationen/Volkswirtschaft/Konjunktur-aktuell.html

Monetary Policy & the Economy — English | quarterly

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http://www.oenb.at/en/Publications/Economics/Monetary-Policy-and-the-Economy.html

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http://www.oenb.at/Publikationen/Statistik/Statistiken---Daten-und-Analysen.html

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http://www.oenb.at/en/Publications/Economics/Conference-on-European-Economic-Integration-CEEI.html
The proceedings have been published with Edward Elgar Publishers, Cheltenham/UK, Northampton/MA, since the CEEI 2001.
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