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Have real exchange rates and competitiveness in Central and Eastern Europe fundamentally changed?

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ABSTRACT

In this paper we analyse the evolution of the RER and its main fundamentals for a group of Central and Eastern European countries. We focus particularly on the possibility of structural breaks in the relationship. We find clear evidence of breaks, mainly caused by the great recession. We find that a combination of both supply-side and demand-side effects are behind the variations in the real exchange rate, and overall, capital inflows have contributed to improve competitiveness in these countries.

1. Introduction

In this paper we analyse how the relationship between the real exchange rate (RER) and its fundamentals may have changed in the past two decades. The RER can be understood as a measure of the competitiveness of a country, since it measures the ratio of national prices to foreign prices in a common currency. Analysis of how the RER evolves is relevant for policy since it is related to the purchasing power parity (PPP) hypothesis, which states that the RER should be equal to one. If it is, then it implies that prices in different countries should be such that the same amount of products could be bought for the same amount of money. The PPP theory has been used as a measure of economic integration, and has been empirically demonstrated increasingly in empirical analyses.

In this paper we analyse the long-run determinants of the RER for a group of Central and Eastern European countries (CEECs), namely Bulgaria, Czechia, Estonia, Hungary, Latvia, Lithuania, Poland, Romania, Slovakia and Slovenia. This group of countries are among the latest to join the EU, and some of them have joined the euro area. Their common communist background and their different degrees of integration with the EU make them of key interest. Many of these countries liberalised their capital markets and eliminated exchange rate limitations before becoming member states. Many of them also experienced large inflows of capital in the years around the time that they joined the EU, but suffered significant reversals in the global financial crisis. These inflows of capital were accompanied by high inflation rates, which forced them to intervene in the markets to avoid large currency appreciations. Each country set up a different exchange rate regime, ranging from currency boards to fully flexible, with the aim of helping the catching up process while suiting the idiosyncrasies of each country. Whether or not the changes that they suffered in the global financial crisis affected the relationship between the RER and its fundamentals, remains unsolved. This paper focuses on how the relationship between the RER and its fundamentals has changed over time. The reason is that, as established in Cuestas (2009) among many others, the PPP

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hypothesis does not seem to hold in Central and Eastern Europe, and a nonlinear deterministic trend is needed to find stationarity in the RERs in these countries. This implies that the RER is not stationary by itself, and we should look at other variables whose stochastic trends could cancel out those of the RER. It may also be noted that the RER is not only a measure of economic integration but also a measure of competitiveness against other countries.

The analysis of the determinants of the RER in CEECs is important for policy reasons. Since competitiveness is on the agenda of most European central banks and ministries of finance and economics because competitive devaluations have become impossible, understanding which key variables determine the RER over the long run can have some relevance for policy (Cuestas, 2023). Most of the countries analysed are already members of the euro area while some others maintain a fixed exchange rate against the euro, and so it is impossible for them to alter the nominal exchange rate. In addition, the Euro Plus Pact and the Macroeconomic Imbalance Procedure of the Six Pack introduced by the EU target several measures of competitiveness, making the analysis of the evolution of competitiveness more important (see Gabrisch & Staehr, 2015).

Three recent contributions that analyse the determinants of the RER are Comunale (2017, 2019 and 2018) and Cuestas et al. (2019). The first two of these focus on the whole of the EU, while the third paper focuses on CEECs. All three papers analyse how the RER affects economic growth, and although Comunale (2017, 2019) and Comunale (2018) only use linear models, Cuestas et al. (2019) introduce nonlinear models to explain how RER deviations impact economic growth.[1] In this paper we focus on the existence of nonlinearities in the long-run equation of the RER, introducing endogenously determined structural breaks. The idea is to study how the relationship has changed over time at discrete points in time. In a similar contribution, Cuestas and Coleman (2023) analyses how the speed of mean reversion has changed over time for the same group of CEECs, applying the Bai and Perron (1998, 2003a and 2003b) method. However, that analysis is univariate and does not include any RER determinants. Here we show that there is a clear change in the speed of mean reversion around the years when the great recession ignited.

In this paper we go a step further by estimating long-run changes in the relationships between the RER and its main fundamentals by using Bayesian methods and the Bai and Perron approach. With the Bayesian method we use Bayesian structural vector autoregressive (BSVAR) models to analyse how the posterior distribution of the RER response functions to an innovation in the fundamentals has shifted, and the differences before and after the beginning of the crisis.

Our results show that there are clear structural breaks, as the estimated parameters differ significantly across the different subperiods. The subperiods are essentially the crisis and the recovery afterwards. This means that policy analysis should focus on the latest period estimated in order to identify policies that could enhance competitiveness.

The remainder of the paper is organised as follows. The next section gives the background to the specification of the model and the expected signs of the parameters. Section 3 summarises briefly the statistical methods used for the empirical analysis. In section 4 we present the data and the main results and, finally, the last section concludes.

2. Background

Our analysis builds upon Comunale (2017a, 2017b), 2018) and Cuestas et al. (2019). In Comunale (2017a, 2017b) and Cuestas et al. (2019), the authors estimate long-run equilibrium equations for the RER, then incorporate misalignments in their growth equations. Comunale estimates the equilibrium equations in a panel set up using determinants derived from Lane & Milesi-Ferretti (2004). She includes inflows, terms of trade and productivity differentials. Cuestas et al. follow a similar approach, using the fundamentals proposed by Berg and Miao (2010) and Vieira and MacDonald (2012) (see also Coleman, 2008). The determinants they use include openness, terms of trade, government consumption, investment and income.[2]

However, to the best of our knowledge there has not been any attempt to update the data to 2019 including or accounting for structural breaks in models that estimate changing parameters.

Our long-run equilibrium RER specification is as follows:

$$q_t = c + \beta_1 ca_t + \beta_2 gco_t + \beta_3 gfcf_t + \beta_4 tot_t + \beta_5 y_t + \varepsilon_t \tag{1}$$

where q is the RER in logs, ca the current account as a percentage of GDP divided by 100, gco the log of real government consumption, $gfcf$ the log of real gross fixed capital formation, tot the log of the terms of trade measured as the ratio between export prices and import prices, and y the log of real GDP. The ca acts as a good proxy for capital flows and hence controls for openness and interest rates. In addition, since the capital markets of these countries are not as developed as those in the West are, the ca should be enough to catch the effect of this type of investment from abroad.

Although β_4 is expected to have a positive sign, the expected sign of the remainder of the variables is ambiguous, and depends on which sector the money is spent in. If the money is spent mainly in the non-tradeable sector then β_1 should be a negative demand shock of y and β_2, β_3 and β_5 should be positive shocks. However, if the spending happens proportionally more in the more productive tradeable sector, then β_1 should be a positive supply shock for y and β_2, β_3 and β_5 should be negative shocks (Benigno & Thoenissen, 2003). We should remember that spending in the more productive sectors helps to improve the competitiveness of the country, and so the real exchange rate would tend to depreciate.

[1] See also Christidou and Panagiotidis (2010) and Cuestas and José Regis (2013) for nonlinear models and the PPP hypothesis among many others.

[2] Comprehensive literature reviews can be found in both Comunale (2018) and Cuestas et al. (2019).

Table 1
DOLS long-run estimates.

	ca	gco	gfcf	tot	y	C
Bulgaria	0.65**	−0.47	0.50**	0.88**	−0.40	7.68**
Czechia	−0.90	1.00	0.16	−1.72*	0.22	−7.99
Estonia	0.64**	0.30	0.09	0.48**	−0.03	2.31*
Hungary	−0.94*	0.34	−0.65**	−0.63	1.14**	−4.28
Latvia	0.26	−0.58**	0.02	1.40**	0.28	5.81**
Lithuania	−0.21	−0.61	−0.33**	−0.57**	0.86**	3.73
Poland	−1.49**	0.19	0.25**	−0.82*	−0.22	2.77
Romania	−0.59	0.00	0.09	0.52**	−0.12	4.85**
Slovakia	1.41**	0.52*	−0.17	1.72	0.36	−1.39
Slovenia	0.38	0.26**	0.07	0.00	−0.08	2.73**

Note: ** significant at 5%. * significant at 10%. Estimations obtained with one lead and one lag. The estimated coefficients for leads and lags of the first differences are not shown. The standard errors have been obtained by HAC.

As explained in the next section, our contribution is in the estimation of different equilibrium relationships for different subperiods. This lets us derive policy implications using the estimations for the latest period found.

3. Methodology

In this paper we estimate cointegrating relationships using the DOLS method proposed by Stock and Watson (1993). However, we aim here to identify structural breaks and estimate broken-type equations as suggested by Bai and Perron (1998, 2003a and 2003b). With this approach we estimate the following DOLS with breaks in the long run regressors:

$$q_t = \sum_{p=1}^{m+1}\left(c_p + \sum_{i=1}^{k}\beta_{p,i}x_{i,t\in p}\right) + \sum_{i=1}^{k}\sum_{j=-h}^{q}\delta_{i,j}\Delta x_{i,t+h} + \varepsilon_t \quad (2)$$

where p refers to a subperiod within the sample. The model allows us to estimate the coefficients for different subperiods, with m breaks. In our case, we have chosen only the long-run parameters β and the constant to be time dependent, while the coefficients of the variables in first differences are non-breaking parameters.

In order to choose the number of breaks m, we use the sequential procedure proposed by the authors from a maximum of breaks.

As robustness analysis we also estimate some of our relationships using BSVARs. We estimate a BSVAR for the full sample and for the subperiods before and after the beginning of the crisis. These models are estimated like structural vector autoregressive (SVAR) models, such as:

$$\delta_0 Y_t = \delta(L)Y_t + \varepsilon_t \quad (3)$$

where δ_0 is the matrix of contemporaneous parameters, δ is a matrix of coefficients for the lagged variables, and L is the lag operator in polynomial form. As δ_0 cannot be fully identified, we use Cholesky decomposition to identify the shocks. Using Bayesian methods has several advantages over traditional frequentist SVARs. First, the Bayes theorem means the additional information that is used as the priors lets us obtain posterior probability distributions of the coefficients and impulse-response functions, as follows:

$$\pi(\partial|Y) = \frac{f(Y|\partial)\pi(\partial)}{f(Y)} \quad (4)$$

where ∂ is a vector of coefficients, $\pi(\partial|Y)$ is the posterior distribution conditional on the sample Y, $f(Y|\partial)$ is the likelihood function, $\pi(\partial)$ is the prior distribution about the parameters, and $f(Y)$ is simply the density function of the data in the sample. We use the independent normal-Wishart prior, which is based on the Litterman (1986) Minnesota prior and, apart from assuming that the parameters are normally distributed and that the series contain a unit root, relaxes the assumption that the residual variance-covariance is unknown. Hence the variance of the coefficients is obtained as follows:

$$\sigma^2_{\delta_{ii}} = \left(\frac{\lambda_1}{l^{\lambda_3}}\right)^2 \quad (5)$$

$$\sigma^2_{\delta_{ij}} = \left(\frac{\sigma_i^2}{\sigma_j^2}\right)\left(\frac{\lambda_1\lambda_2}{l^{\lambda_3}}\right) \quad (6)$$

Following the literature, we choose $\lambda_1 = 0.1$, $\lambda_2 = 0.5$ and $\lambda_3 = 2$. The second advantage comes because unit roots or a lack of cointegration between I (1) processes do not affect the inference since t or F statistics are not used (Sims, 1988).

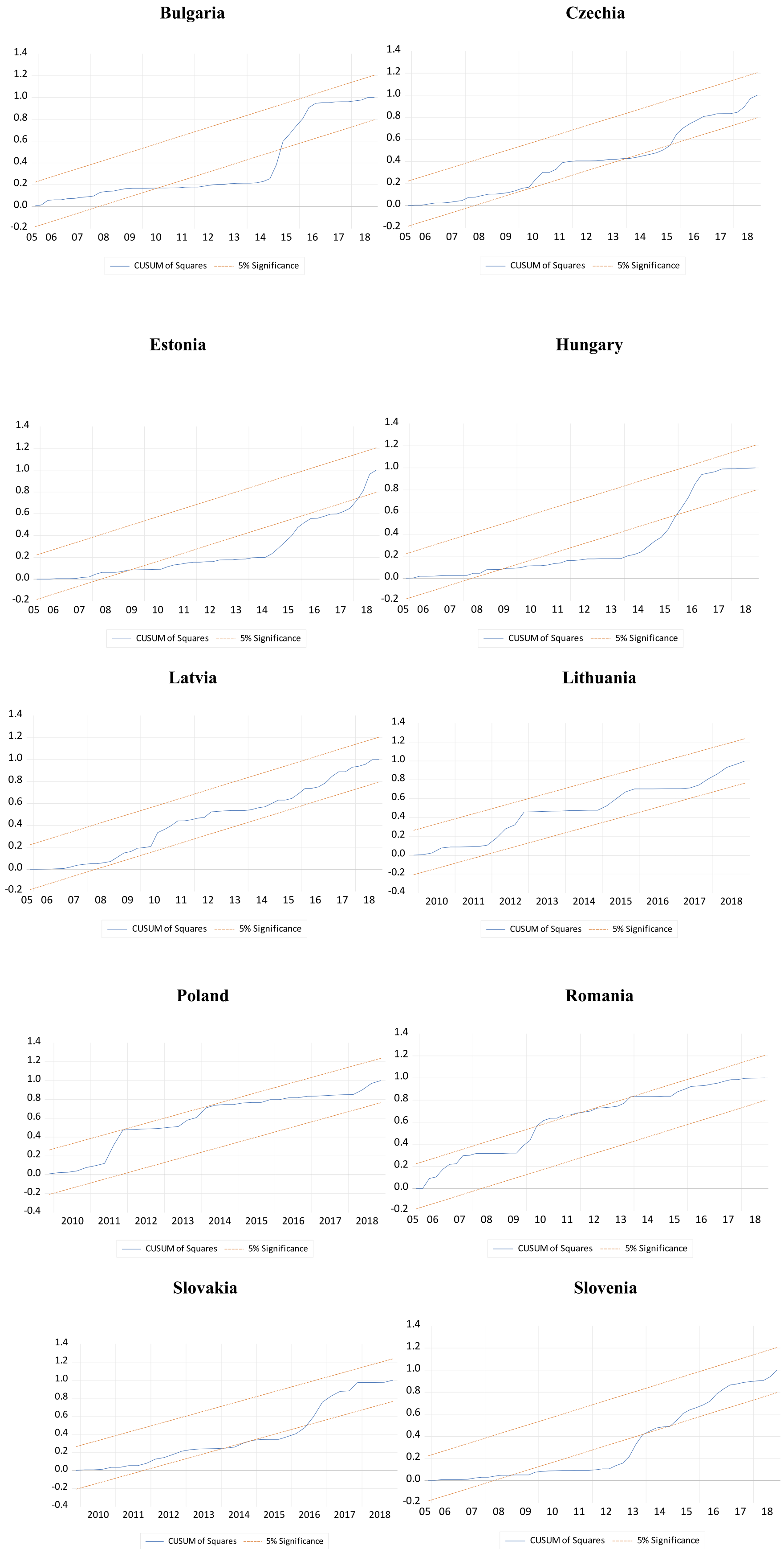

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Fig. 1. CUSUM of squares.

4. Empirical analysis

The data consist of quarterly series for the log of the real effective exchange rate (deflator: consumer price index – 37 trading partners – industrial countries), *q*; the current account as a percentage of GDP divided by 100, *ca*; log of real government consumption, *gco*; log of real gross fixed capital formation, *gfcf*; log of the terms of trade measured as the ratio between export prices and import prices, *tot*; and log of real GDP, *y* for our target CEECs. The series run from 2000Q1 until 2019Q1, except for those of Poland, Lithuania and Slovenia, which start in 2004Q1. The data have been downloaded from *Eurostat* and whenever available, the series have been downloaded as seasonally adjusted. The series were not available seasonally adjusted for Slovakia, so they have been adjusted using the *Census X13* procedure.[3]

Now we present the results of our estimations. First, we have tested for unit roots in the data and found that all the series are I (1) processes in levels. To make sure that our regressions are not spurious, we have tested for cointegrating relationships between the variables and we have found evidence of cointegration in all cases.[4]

The estimations have been carried out using DOLS, which lets us estimate long-run relationships controlling for endogeneity by including leads and lags of the first differences of the regressors. In our estimations we have used one lead and one lag throughout the paper. The standard errors have been obtained by the Newey-West heteroscedastic and autocorrelation correction (HAC) method.

In Table 1 we present the results for the long-run coefficients for a model with no breaks. The results fail to show clear evidence that the variables explain the long run of the real exchange rate. Only in the cases of Bulgaria, Hungary, Lithuania and Poland did three of the five regressors seem to be significant at conventional significance levels. The current account is only significant and with the expected sign in Hungary and Poland. Government consumption is only significant in Latvia, Slovakia and Slovenia. Gross fixed capital formation is significant in Bulgaria, Hungary, Lithuania and Poland. The terms of trade have the expected sign and are significant for Bulgaria, Estonia, Latvia and Romania. Finally, real income is only significant for Hungary and Lithuania. In Fig. 1, we present the CUSUM of squares instability test, which shows that there are clear signs of instabilities in many of the estimations.

As a preliminary and complementary analysis we estimate BVARs for all the countries. Fig. 2 presents the impulse response functions for the effect of different innovations on the real exchange rate, using the following ordering: *ca, gco, gfcf, tot, y* and *q*. This means that the real exchange rate is allowed to react contemporaneously to innovations in the other variables. The results show a similar picture to that in Table 1. However, in some of the cases where there was no significant effect in Table 1, the impulse-response function shows that the posterior distribution shifts significantly.

In order to see how the parameters may have changed before and after the crisis, we estimate the BVAR for the periods before and after 2009. The results are displayed in Fig. 3. It is apparent that the innovations affect the RER differently before and after 2009.

It is particularly notable that shocks to all the variables tend to have a positive impact on the RER before the crisis, except in the Baltic states, but are ambiguous after the crisis ignites. This proves yet again that positive shocks during the great moderation before the crisis caused the RER to appreciate in most of our target countries. The RERs in the Baltic states appreciated substantially during the transition process from the USSR to market economies, which is shown in our analysis.

Finally, we complement this analysis by estimating the DOLS equation as a panel by pooled weights and HAC for both subperiods. Again, the results from Tables 2 and 3 show important changes in the values and signs of the parameters before and after the crisis.

To account for structural breaks endogenously determined, we use the Bai and Perron (1998, 2003a and 2003b) method to estimate the DOLS equations. We assume for this that the leads and lags of the first difference of the regressors do not change over the sample. We first obtain the number of breaks using the sequential method of Bai and Perron, with a maximum of two breaks and one lead and one lag. The standard errors have again been corrected using the Newey-West HAC method. We find in nearly all cases that we should estimate the models with two breaks, with the exception of Poland, where no break is found. That Poland is an exception is not surprising since that country hardly felt the crisis. The results are displayed in Table 4. With these estimations we find more evidence of significance for the variables. We will explain the results country by country.

- Bulgaria: joined the EU in 2007. The results show that the current account has a positive and significant impact on the real exchange rate in the three subperiods, whereas government consumption appears to have a negative effect only from 2003, which is a sign of high spending by the government on tradeables. Investment proxied by *gfcf* also had a positive effect on the real exchange rate. However, income does not seem to play a role for this country.
- Czechia: joined the EU in 2004. The estimations show that the first break coincides with the beginning of the crisis, and the second with the beginning of the full recovery. The current account only seems to have a role before the crisis, which shows how capital inflows were feeding this country but then caused a depreciation. Public sector consumption seems to have a positive effect until 2015 but a negative impact afterwards, which may show a shift in the type of government consumption. The terms of trade seem to

[3] The charts of the series can be found in a working paper version of this paper published by the authors: https://ideas.repec.org/p/jau/wpaper/2019-12.html.

[4] Results available on request.

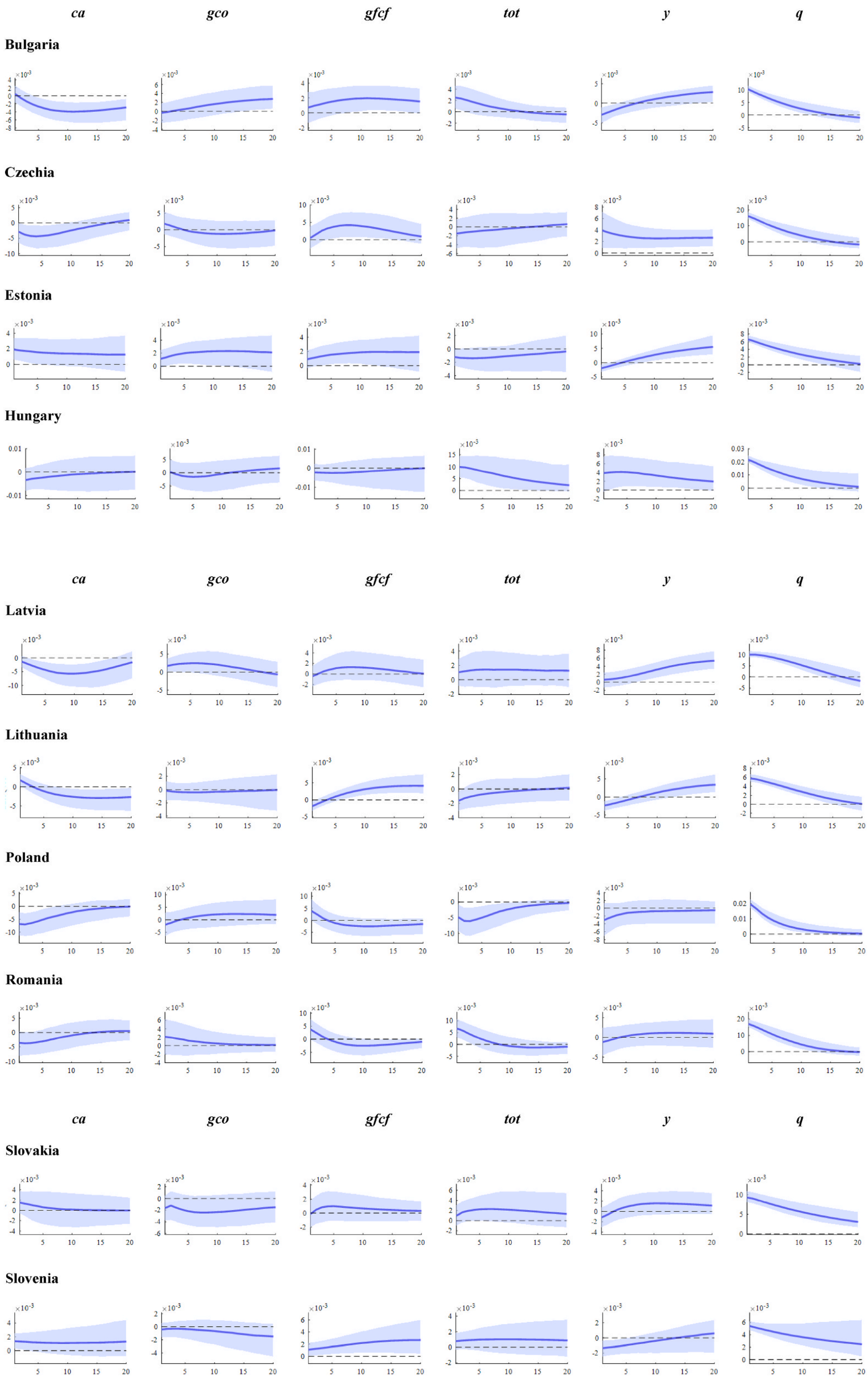

Fig. 2. Response of real exchange rates to innovations.

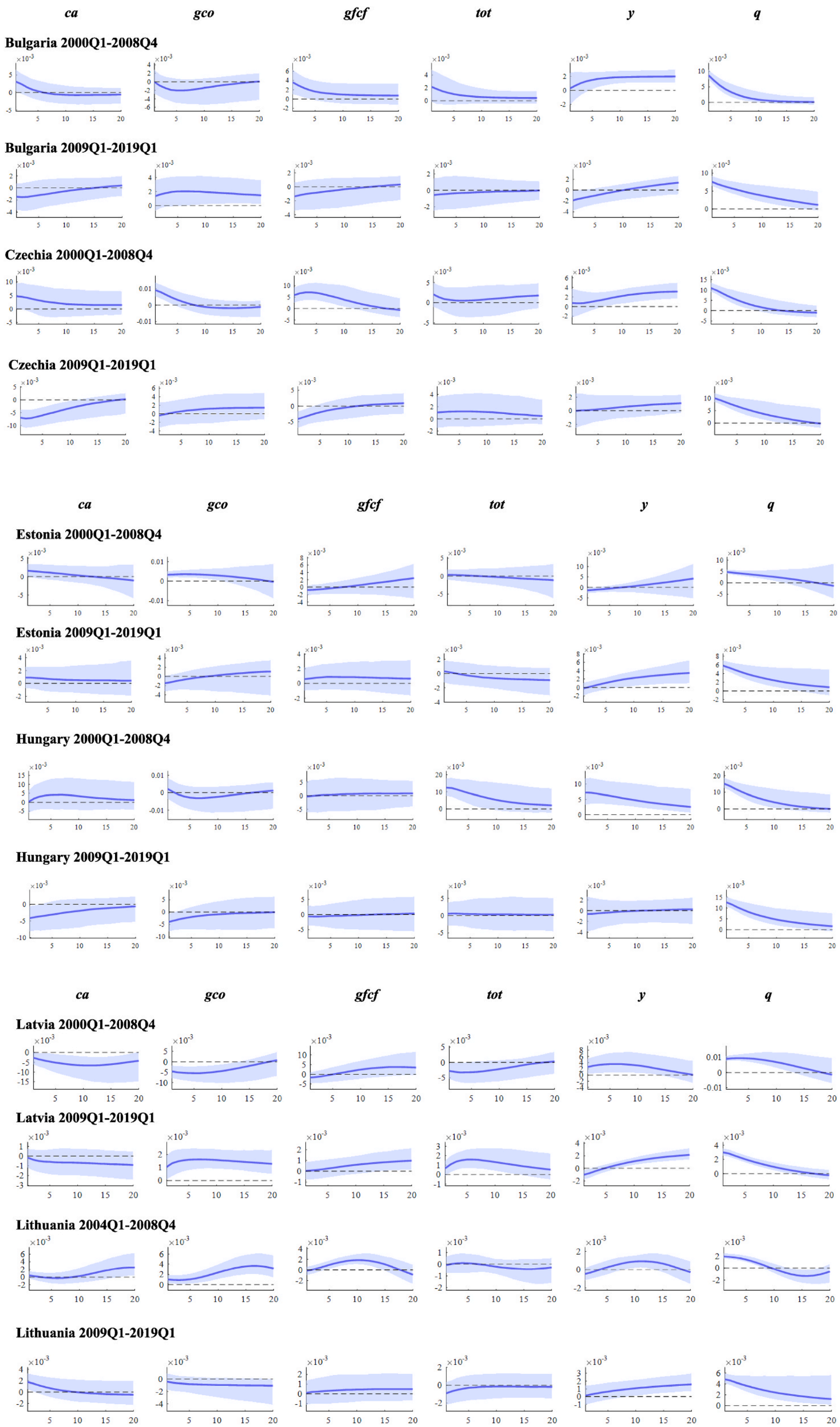

Fig. 3. Response of real exchange rate to innovations, subperiods.

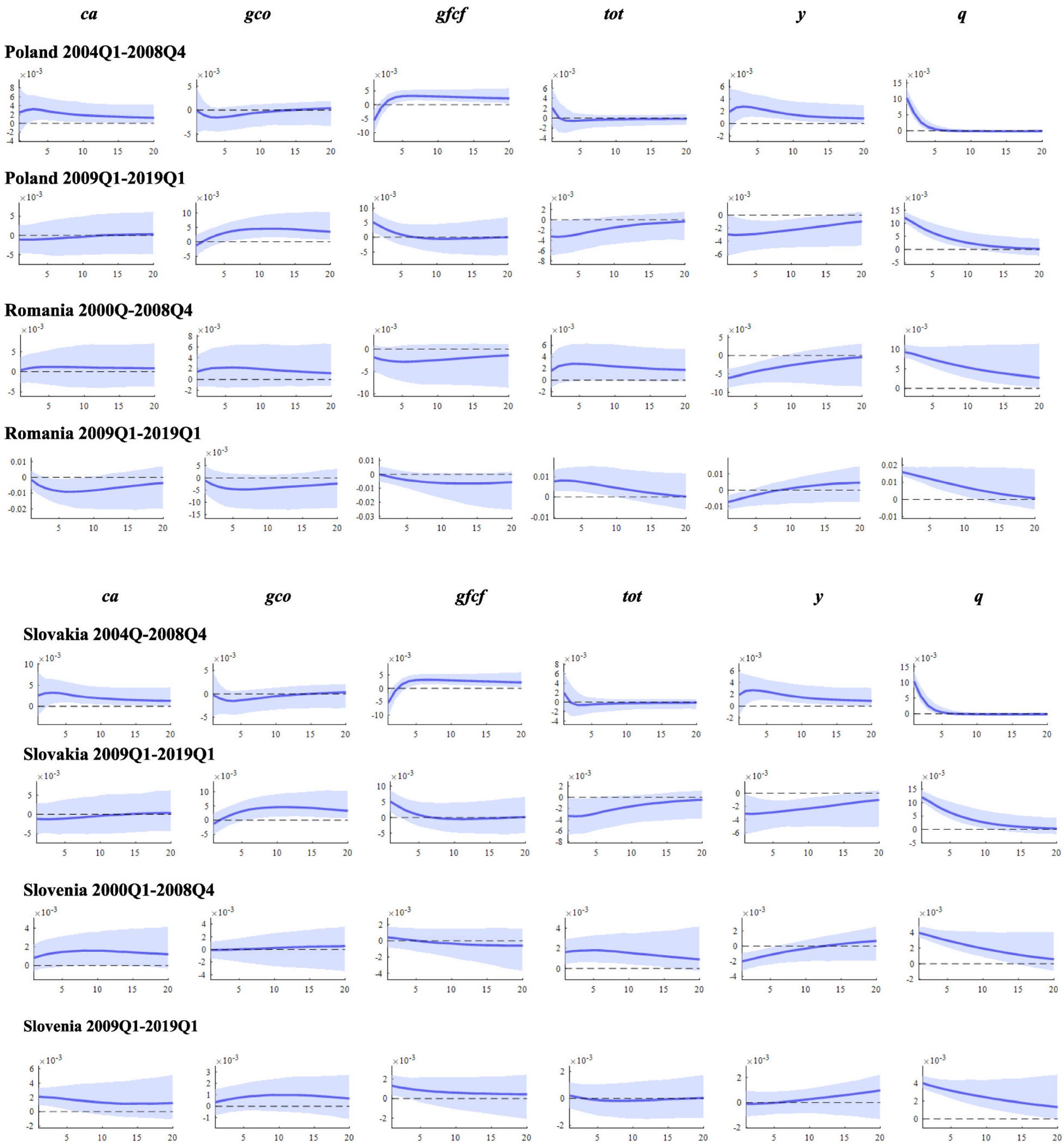

Fig. 3. (*continued*).

have had a negative impact across the full period, whereas income only has a positive effect after 2015, highlighting that supply-side growth may have been exhausted.

- Estonia: the years 2005–2014 are marked by an abnormal boom and bust period. The current account only had an impact on the RER after 2015, when it pushed the currency up in real terms. Public consumption had a positive impact before 2015. Gross fixed capital formation seems to have made a positive contribution until 2015, perhaps because of the catching up process and the effect on the housing market. The terms of trade show a negative sign, which may reflect how rises in the prices of Estonian products imply that international market position was not necessarily lost for Estonian products. Finally, as in Czechia, income seems to have been supply driven before 2015 and demand driven afterwards. Estonia joined the EU in 2004.
- Hungary: the second break coincides with the eruption of the crisis, but the magnitude and in some cases the significant of the variables may have changed in the different subperiods, though the signs remain the same in three of them. The current account

Table 2
DOLS long-run estimates in panel 2000–2008.

Variable	Coefficient	Std. Error	t-Statistic	Prob.
ca	0.31	0.25	1.22	0.22
gco	−0.04	0.13	−0.32	0.75
gfcf	0.05	0.10	0.47	0.63
tot	0.36	0.17	2.08	0.04
y	0.28	0.16	1.76	0.07

Note: The estimated coefficients for leads and lags of the first differences are not shown. Estimations obtained with one lead and one lag. The estimated coefficients for leads and lags of the first differences are not shown. The standard errors have been obtained by HAC.

Table 3
DOLS long-run estimates in panel 2009–2019.

Variable	Coefficient	Std. Error	t-Statistic	Prob.
ca	0.38	0.10	3.68	0.00
gco	0.07	0.06	1.05	0.29
gfcf	0.11	0.02	4.04	0.00
tot	0.13	0.09	1.36	0.17
y	−0.15	0.05	−2.69	0.01

Note: The estimated coefficients for leads and lags of the first differences are not shown. Estimations obtained with one lead and one lag. The estimated coefficients for leads and lags of the first differences are not shown. The standard errors have been obtained by HAC.

only had a negative impact on the evolution of the real exchange rate after 2005. The negative sign for government consumption shows that government consumption was most probably mainly in tradeables. The effect of the terms of trade is always positive and significant, as is the effect of income. Hungary joined the EU in 2004.

- Latvia: the current account is only significant in the middle years of 2005–2009. Government spending had a positive impact in the full period as did investment. The terms of trade do not appear to have had a significant effect on the real exchange rate in the sample. Although income is only significant in the first and last subperiods, it had a negative impact on the evolution of the real exchange rate. Again, this may be due to economic growth being of a more supply-side type in Latvia. Latvia joined the EU in 2004.
- Lithuania: the current account only had a positive and significant impact after the beginning of the crisis, whereas government consumption had a negative effect on the real exchange rate before the crisis and after the full recovery in 2015. Investment had a positive impact after the beginning of the crisis. The terms of trade vanish from the equation from 2015, since their negative sign is only significant before 2015. Lithuania joined the EU in 2004.
- Poland: this country did not experience a drop in output growth at any point in these years, which corroborates the finding that there is no evidence of changing parameters. The current account had a negative impact on the evolution of the real exchange rate, as did the terms of trade. However, the effect of investment was positive throughout the sample. Government spending and income did not have any significant effect on the real exchange rate. Poland joined the EU in 2004.
- Romania: the central years of 2005–2011 coincide with a period of relatively fast boom and bust. Romania joined the EU in 2007. Only government consumption, investment and output seem to have had an impact on the real exchange rate in the sample. The negative sign of the first two of these implies a higher proportion of spending went on tradeable goods, whereas the positive sign in output again shows evidence of demand driven growth.
- Slovakia: the central period covers the beginning of the crisis and the beginning of the recovery. We find that the current account is positive and significant in the whole sample, whereas government consumption is only significant before the crisis. The terms of trade are negative in the extreme periods but positive in the central period. Slovakia also joined the EU in 2004.
- Slovenia: the first break occurs in 2007Q3 and may be a prelude to the beginning of the crisis, while the second is in 2011 with the beginning of the end of the crisis. The current account did not have any impact, while government consumption only has an impact before the crisis. The terms of trade change in sign after the crisis and output seems to have a negative effect, again meaning that economic growth was based more on supply. Slovenia joined the EU in 2004, like most of our target countries.

Some commonalities are worth highlighting. There are two breaks in all cases except Poland. The first occurs towards the mid-2000s and coincides with membership of the EU, and in some cases with the beginning of the global financial crisis. The second break coincides with the beginning of the global financial crisis for some and for others with the recovery from that crisis. We observe that most variables carry a positive sign or are not significant in the first subperiod. This shows that most countries faced the same types of shock during this transition period. It is interesting though that the sign of government consumption is negative for the first subperiod for Hungary and Romania. This pattern repeats for the same countries in the second subperiod, when government spending is focused in more productive sectors. Investment, proxied by the gross fixed capital formation, has a consistent positive sign across countries and periods in most cases, which implies that it has worked in support of appreciation of the currency during the period analysed. It is important to note the effect of income on the RER. It is difficult to extrapolate a general pattern for this variable in the first and second subperiods, but where it is significant in the second subperiod it is also negative, except in Estonia. As mentioned earlier, the second subperiod coincides with EU membership and the recovery after the global financial crisis. This shows the efforts of

Table 4
DOLS long run estimates with breaks.

Country	Bulgaria	Czechia	Estonia	Hungary	Latvia	Lithuania	Poland	Romania	Slovakia	Slovenia
Variable/period	00:3–03:2	00:3–09:2	00:3–05:1	00:3–05:3	00:3–05:1	04:3–08:1	04:3–18:4	00:3–04:4	04:3–08:2	00:3–07:2
ca	1.04**	1.74**	1.61	−0.27	−0.22	0.06	−1.49**	−0.18	1.36**	−0.16
gco	0.17	0.82*	0.18*	−0.91**	0.82**	−1.14**	0.19	−0.25**	0.29**	0.55**
gfcf	0.59**	1.16**	0.14	−0.79**	0.23**	−0.01	0.25**	−0.90**	−0.12	0.12
tot	0.85**	−0.52	−0.01	4.47**	0.24	−0.50**	−0.82*	0.03	−4.30**	0.62**
y	−0.59	−0.82**	0.36*	2.9**	−1.16**	0.50*	−0.22	1.09**	0.03	−0.36*
C	4.29**	−4.56**	0.26**	−10.0**	6.91**	8.07**	2.77	3.21	3.03**	2.85**
	03:3–14:3	09:3–15:1	05:2–14:3	05:3–08:1	05:2–09:1	08:2–14:4		05:1–11:1	08:3–12:1	07:3–11:1
ca	0.47**	0.82	0.38	−1.84**	0.41**	0.53**		0.22	0.90**	0.06
gco	−0.38**	1.04**	0.42**	−1.71**	1.91**	−0.24		−0.28**	−0.04	0.15
gfcf	0.35**	1.34**	0.62**	−0.06	0.32*	0.24**		0.23**	0.04	0.18**
tot	1.03**	−3.31**	0.01	1.62**	−0.04	−0.34**		0.01	0.60**	−0.34*
y	−0.17	−1.95**	0.43**	0.21	−0.70	−0.52**		−0.10	0.21	−0.75**
C	6.03	3.43**	0.02	17.33**	−4.27**	9.11**		5.91	2.58**	8.93**
	14:4–18:4	15:2–18:4	14:4–18:4	08:2–18:4	09:2–18:4	15:1–18:4		11:2–18:4	12:2–18:4	11:2–18:4
ca	0.61**	0.18	2.02**	−0.36*	−0.21	0.46**		0.06	0.91**	0.21
gco	−0.53*	−1.49*	0.04	−1.45**	0.64**	−1.58**		−0.36**	−0.02	−0.07
gfcf	0.45**	0.49	−0.53**	−0.04	0.10**	0.29**		−0.37**	0.09	0.26**
tot	0.38	−0.73*	−0.08**	1.79**	0.17	0.24		0.05	−1.45*	0.65**
y	0.13	1.79**	0.31**	0.35**	−0.47**	0.10		0.16**	−0.01	−0.32**
C	3.67**	−5.68**	0.81**	13.80**	3.72**	12.39**		9.26**	4.10**	6.10**

Note: Model estimates with non-breaking one lead and one lag of the regressors in first differences. The estimated coefficients for leads and lags of the first differences are not shown. ** significant at 5%. * significant at 10%. The standard errors have been obtained by HAC.

these countries to gain competitiveness through productivity and supply-side policies when they were catching up with the rest of the EU members and after the crisis ignited. Focusing on the latest of the subperiods lets us say in general that capital inflows have had a positive impact on reducing prices and so depreciating the currency. Similar results are found by Comunale (2017a, 2017b) and proposed by the theoretical model of Benigno and Thoenissen (2003). Government consumption, investment and the terms of trade do not seem to have a homogenous pattern amongst the countries analysed, whereas real output seems to have generally had a more demand-side effect, although with some exceptions.

5. Conclusion

In this paper we have analysed the behaviour of the RER fundamentals in a long-run relationship with the RER. We have used different methods to estimate the long-run elasticities for different subperiods both before and after the beginning of the crisis, and with breaks endogenously determined.

The results show that it is necessary to account for changing parameters, and that both supply-side effects and demand-side effects can be observed, depending on the countries analysed.

The findings show that most of the countries (but Poland) exhibit two breaks that differentiate three periods with similar patterns but marked for different years. This fact would indicate unrelated economic cycles and the distinct convergence rate among them, and relative to the EU.

Of our target countries, the three Baltic states, Slovakia and Slovenia are members of the euro area. The remaining countries should eventually join the euro according to their membership agreements. The Maastricht criteria cannot really be assessed from our results, but it should be remembered that these criteria, which were designed to ensure nominal convergence, assumed implicitly that real convergence would naturally follow. From the experience of the euro, we know that this assumption was wrong and that what matters for real convergence are the drivers of competitiveness. This result justifies the analysis of RER in the context of countries that wish to adopt the common currency.

Nevertheless, we must also bear in mind that EU countries that want to apply to adopt the euro need to meet the Maastricht criteria. One of the Maastricht criteria is that the currency of the country needs to have been in the Exchange Rate Mechanism II for at least two years without any devaluation against the euro. So far only Bulgaria has taken this first step towards adopting the euro.

In order to enhance competitiveness, governments should enhance those factors that depreciate the currency and reduce those that tend to appreciate it. Capital inflows for example seem to have a positive effect on competitiveness in most cases.

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