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BEER tastes better in a panel of neighbours. On equilibrium exchange rates in CEE countries

Summary

We follow the behavioural equilibrium exchange rate (BEER) approach to estimate the misalignments of real exchange rates in selected Central European countries, including Poland, in a sample that covers both pre- and post-crisis period. We add to the existing literature by applying a panel approach with FM-OLS estimator of the cointegrating relationship that represents the BEER equation. In our quarterly sample, comprising the years 2000–2013 for Poland, Hungary, Czech Republic and Romania, the parameters of real interest rate disparity, risk premium, Harrod–Balassa–Samuelson effect and terms of trade take the expected sign. These estimates seem to be more precise and robust to post-crisis instability than those obtained on the basis of time series approach, at least for Poland.

Keywords: behavioral equilibrium exchange rate, FM-OLS, panel cointegration, CEE exchange rates

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1. Introduction

Increased volatility of Central and Eastern European currencies that emerged in the course of the financial and fiscal crises after 2008 gave rise to questions whether individual currencies became over- or undervalued, or just reverted to valuation based on countries' macroeconomic fundamentals. These questions are still valid and open and might resurface in the future, as all the free-floaters among EU's New Member States are obliged to adopt the euro. This requires entering the ERM II mechanism, and hence setting a reasonable central parity that would likely evolve into the conversion rate of national currencies against the euro. Failure to set a rate compatible with macroeconomic fundamentals could result in a costly and prolonged period of competitiveness adjustment.

There has been huge empirical literature on modelling equilibrium exchange rates in CEE countries and in Poland in particular. Among a few methodological approaches applied over the recent years, one should at least mention FEER[4], CHEER[5], and BEER[6]. This paper adds to the last of the abovementioned strands of literature by applying the BEER approach to a panel of CEE countries: Poland, Czech Republic, Hungary and Romania. This set of countries is relatively homogeneous in terms of the monetary regime (*de facto* free or managed floaters, for most of the sample period)[7], fundamental characteristics (post-communist economies in the catching-up process and new member states of the EU since 2004–2007), and – last but not least – data availability. Égert et al. 2005 points to a sufficient similarity within this group to form a panel. With non-stationary real exchange rate series, we estimate the cointegrating relationship defining BEER with FM-OLS method[8].

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[4] M. Rubaszek, *Economic convergence and the fundamental equilibrium exchange rate in Poland*, "Bank i Kredyt" 2009, vol. 40(1), pp. 7–22.

[5] P. Kębłowski, A. Welfe, *Estimation of the equilibrium exchange rate: the CHEER approach*, "Journal of International Money and Finance" 2010, vol. 29(7), pp. 1385–1397.

[6] J. Bęza-Bojanowska, R. MacDonald, *The Behavioural Zloty/Euro Equilibrium Exchange Rate*, National Bank of Poland Working Papers 55, National Bank of Poland, Economic Institute, 2009; R. Kelm, *Model behawioralnego kursu równowagi złotego do euro w okresie styczeń 1996–czerwiec 2009 r.*, "Bank i Kredyt" 2010, vol. 41(2), pp. 21–41.

[7] This is the criterion that excluded most of the other EU new member states from our analysis, in particular the Baltic states (ERM II participants), Slovenia, Cyprus, Malta (euro adoption in 2006–2008) and Slovakia (ERM II since 2005, euro adoption in 2009).

[8] P. Pedroni, *Fully Modified OLS for Heterogeneous Cointegrated Panels*, in: *Nonstationary Panels, Panel Cointegration and Dynamic Panels*, ed. B.H. Baltagi, Elsevier, Amsterdam 2000, pp. 93–130.

By extending our empirical basis from one country to a panel, we attempt to solve two problems. Firstly, available quarterly samples are too short to effectively apply the state-of-the-art time-series frameworks for nonstationary data, like vector equilibrium correction. Simplified approaches (like Engle-Granger or DOLS/FM-OLS on a time series bases) also seem to be deficient here. Therefore, a switch to a panel should reinforce our precision of estimates while avoiding the use of monthly data, which are often noisy and approximated. Secondly, the sample covers some post-crisis periods and this poses further challenges regarding parameter stability. A panel equation is less likely to be overfitted to specific developments in an individual economy; such developments, on the other hand, provide better guidance of the currency's misalignment.

This paper is structured as follows. In Section 2, we present our methodological framework, including the BEER approach and the FM-OLS estimator. Section 3 contains the description of our sample and our empirical findings. Section 4 concludes.

2. Methodological framework: using FM-OLS to estimate BEER equation

The behavioural equilibrium exchange rate (BEER) approach, proposed by Clark and MacDonald[9], formulates the real exchange rate of a country as a function of long-, mid- and short-term factors that determine its level and fluctuations in the respective horizon:

$$q_t = \boldsymbol{\beta}_1' \mathbf{Z}_{1t} + \boldsymbol{\beta}_2' \mathbf{Z}_{2t} + \boldsymbol{\tau}'\mathbf{T}_t + \varepsilon_t, \tag{1}$$

where: q_t – real exchange rate, $\boldsymbol{Z}_{1t}$ – vector of long-term fundamentals, $\boldsymbol{Z}_{2t}$ – vector of mid-term variables, usually linked to the business cycle, $\boldsymbol{T}_t$ – vector of short-term variables, $\boldsymbol{\beta}_1$, $\boldsymbol{\beta}_2$, $\boldsymbol{\tau}$ – respective parameter vectors, ε_t – error term. In this framework, the components $\boldsymbol{\tau}'\mathbf{T}_t + \varepsilon_t$ stand for misalignment from the equilibrium.

[9] P. Clark, R. MacDonald, *Exchange rates and economic fundamentals: a methodological comparison of BEERs and FEERs*, in: *Equilibrium Exchange Rates*, eds R. MacDonald, J. Stein, Kluwer, Amsterdam 1999, pp. 285–322.

Equation (1) is commonly estimated in a cointegration framework[10], as the real exchange rate is supposed to revert to long-term equilibrium after stochastic shocks. The use of nonstationary data requires the application of appropriate estimation techniques, both with time-series and in panel data[11]. Here, we use the fully-modified ordinary least squares estimator (FM-OLS) proposed by Perdoni[12] and, according to the author, equipped with good finite-sample properties. Consider a set of equations:

$$\begin{aligned} y_{it} &= \mathbf{X}_{it}'\boldsymbol{\beta} + \gamma_i + u_{1it} \\ \mathbf{X}_{it} &= \mathbf{X}_{it-1} + \mathbf{u}_{2it} \quad , \end{aligned} \tag{2}$$

where γ_i denotes a constant specific to i-th unit in the panel. Phillips and Moon[13] define the estimators of contemporaneous ($\boldsymbol{\Sigma}_i$) and long-term variance-covariance (forward-looking $\boldsymbol{\Lambda}_i$ and joint $\boldsymbol{\Omega}_i$) of the shocks $\mathbf{u}_{it} = (u_{1it}, \mathbf{u}_{2it}')'$:

$$\boldsymbol{\Sigma}_i = E(\mathbf{u_{it}}, \mathbf{u'_{it}}) = \begin{bmatrix} \sigma_{\mathbf{11i}} & \boldsymbol{\sigma}_{\mathbf{12i}} \\ \boldsymbol{\sigma}_{\mathbf{21i}} & \boldsymbol{\Sigma}_{\mathbf{22i}} \end{bmatrix} \tag{3}$$

$$\boldsymbol{\Lambda}_i = \sum_{j=0}^{\infty} E(\mathbf{u}_{it}, \mathbf{u}'_{it-j}) = \begin{bmatrix} \lambda_{11i} & \boldsymbol{\lambda}_{12i} \\ \boldsymbol{\lambda}_{21i} & \boldsymbol{\Lambda}_{22i} \end{bmatrix} \tag{4}$$

$$\boldsymbol{\Omega}_i = \sum_{j=-\infty}^{\infty} E(\mathbf{u}_{it}, \mathbf{u}'_{it-j}) = \begin{bmatrix} \omega_{11i} & \boldsymbol{\omega}_{12i} \\ \boldsymbol{\omega}_{21i} & \boldsymbol{\Omega}_{22i} \end{bmatrix} = \boldsymbol{\Lambda}_i + \boldsymbol{\Lambda}_i' - \textstyle\sum_i. \tag{5}$$

FM-OLS assumes a single cointegrating relationship and non-cointegrated regressors (implying non-singular Ω_{22i} matrices). The assumption of common long-term variances for all units in the panel leads to the unweighted version of the estimator. In the weighted version, applied in this analysis[14], the following computations are performed for individual units i:

[10] F. Maeso-Fernandez, C. Osbat, B. Schnatz, *Towards the estimation of equilibrium exchange rates for CEE acceding countries: methodological issues and a panel cointegration perspective*, European Central Bank Working Papers, vol. 353, 2004.

[11] A. Welfe, *Ekonometria. Metody i ich zastosowanie*, PWE, Warszawa 2009.

[12] P. Pedroni, *Fully Modified OLS*..., op.cit.

[13] P.C.B. Phillips, H.R. Moon, *Linear Regression Limit Theory for Nonstationary Panel Data*, "Econometrica" 1999, vol. 67, pp. 1057–1111.

[14] Also see: P. Pedroni, *Fully Modified OLS*..., op.cit.

$$\hat{\omega}_{1.2} = \hat{\omega}_{11} - \hat{\boldsymbol{\omega}}_{12i}\hat{\boldsymbol{\Omega}}_{22i}^{-1}\hat{\boldsymbol{\omega}}_{21i}$$

$$\hat{\boldsymbol{\lambda}}_{12i}^{+} = \hat{\boldsymbol{\lambda}}_{12i} - \hat{\boldsymbol{\omega}}_{12i}\hat{\boldsymbol{\Omega}}_{22i}^{-1}\hat{\boldsymbol{\Lambda}}_{22i} \tag{6}$$

$$\tilde{y}_{it}^{++} = \tilde{y}_{it} - \hat{\boldsymbol{\omega}}_{12}\hat{\boldsymbol{\Omega}}_{22}^{-1}\hat{\mathbf{u}}_{2} - \hat{\omega}_{1.2i}^{1/2}(\hat{\omega}_{1.2i}^{1/2}\tilde{\mathbf{X}}'_{it} - (\hat{\boldsymbol{\omega}}_{12i}^{1/2}\tilde{\mathbf{X}}_{it})')\hat{\boldsymbol{\beta}}_{0}$$

with $\hat{\boldsymbol{\beta}}_{\mathbf{0}}$ as first-step estimates of long-term coefficients using a consistent estimator (for example OLS[15]. Then, individual variables are weighted by the reciprocal of long term variances:

$$\tilde{\mathbf{X}}_{it}^{*} = \hat{\boldsymbol{\Omega}}_{\mathbf{22i}}^{\mathbf{-1/2}}\tilde{\mathbf{X}}_{\mathbf{it}}$$

$$\tilde{y}_{it}^{*} = \hat{\omega}_{1.2}^{-1/2}\tilde{y}_{it}^{++} \tag{7}$$

$$\hat{\boldsymbol{\lambda}}_{12i}^{*} = \hat{\omega}_{1.2}^{-1/2}\hat{\boldsymbol{\lambda}}_{12i}^{+}\hat{\boldsymbol{\Omega}}_{22i}^{-1/2'}$$

Weighted FM-OLS estimator takes the form:

$$\hat{\boldsymbol{\beta}}_{FM-OLS} = \Big(\sum_{i=1}^{N}\sum_{t=1}^{T}\tilde{\mathbf{X}}_{it}^{*}\tilde{\mathbf{X}}_{it}^{*'}\Big)^{-1}\sum_{i=1}^{N}\sum_{t=1}^{T}(\tilde{\mathbf{X}}_{it}^{*}\tilde{y}_{it}^{*} - \hat{\boldsymbol{\lambda}}_{12i}^{*}{'}) \tag{8}$$

with the asymptotic variance-covariance matrix[16]:

$$\hat{\mathbf{V}}_{FM-OLS} = \Big(\frac{1}{N}\sum_{i=1}^{N}\Big(\frac{1}{T^2}\sum_{t=1}^{T}\tilde{\mathbf{X}}_{it}^{*}\tilde{\mathbf{X}}_{it}^{*'}\Big)\Big)^{-1}. \tag{9}$$

3. BEER model for a panel of CEE countries

We use data covering four countries (PL, CZ, HU, RO) over the period 2000Q1-2013Q4. This panel is unbalanced, and the source of all data is the Eurostat database.

Our dependent variable is the quarterly average real exchange rate (RER) of the national currency against the euro, deflated with producer price index

[15] See: J. Bai, C. Kao, *On the Estimation and Inference of a Panel Cointegration Model with Cross-Sectional Dependence*, Center for Policy Research, Paper 89, 2005.

[16] P. Pedroni, *Fully Modified OLS*..., op.cit.

(see Bęza-Bojanowska and MacDonald[17] for motivation of this deflator) and defined so that an increase reflects depreciation of a national currency against the euro (*D* stands for the domestic economy, *EA* – for the euro area):

$$RER^D = E^D \cdot \frac{P^{EA}}{P^D} \tag{10}$$

Based on the literature[18], we use 2 standard sets of regressors: real interest rate disparity, general government deficit (or, alternatively, debt), net foreign assets (NFA), relative terms of trade, as well as a proxy for Harrod–Balassa–Samuelson effect. We also use real oil price in one of the models (its use in the other model provides statistically and economically unreliable results).

<u>Real interest rate disparity (RIR)</u> is calculated using 3M money market rates and 12-month PPI price dynamics. Once growing, it should in principle control for currency appreciation due to cyclical factors[19]. We avoid using disparity calculated on the basis of government bond yields, as the risk attributable to such instruments has been fundamentally repriced during the crisis and this variable could capture the effect of macroeconomic and fiscal risk.

The latter is measured with <u>general government (GG) debt or deficit (DEBT/DEF)</u>, expressed as a percentage of GDP and adjusted to quarterly frequency via linear interpolation. By using these two variables for robustness check, we follow the exercise of Bęza-Bojanowska and MacDonald[20] for Poland, but – like these authors – we do not find qualitative differences. Expressed as positive numbers, these variables are expected to cause depreciation when growing. This variable is not calculated in relative terms, i.e. affects only the 4 economies in question (and not the euro area). In consequence, we implicitly treat investors' reactions to fiscal „bad news" asymmetrically, i.e. we expect depreciation to occur even when deficit or debt grow comparably or less in a CEE country than in the euro area as a whole. This seems to be consistent with the stylized facts regarding the market risk perception and management over the sample period.

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[17] J. Bęza-Bojanowska, R. MacDonald, op.cit.

[18] Ibidem; J. Frait, L. Komárek, M. Melecky, *The real exchange rate misalignment in the five central European countries*, Working Paper, University of Warwick, Department of Economics, Coventry 2006.

[19] M. Rubaszek, op.cit.

[20] J. Bęza-Bojanowska, R. MacDonald, op.cit.

Net foreign assets (NFA) were calculated as suggested by Lane and Milesi--Ferretti[21]:

$$\begin{aligned} NFA_t &= NFA_{t-1} + \Delta NFA_t \\ \Delta NFA_t &\cong CA_t + KA_t \end{aligned} \tag{11}$$

with CA – current account, KA – capital account, all expressed as percentage of GDP of the country of interest. In theory, growing NFA should lead to appreciation, because[22] it increases an economy's credibility and improves current and expected CA. However, some authors also find a reverse relationship: in their opinion, the FDI inflow in a catching-up economy should boost the demand for home currency and outweigh the previously mentioned effects (a phenomenon described as „financial deepening").

The proxy for Harrod–Balassa–Samuelson effect (HBS) is calculated as a ratio between the labour productivities in the tradable and nontradable sector, in relative terms between a CEE country and the euro area. Employing quarterly data allows to use NACE accounts and treat agriculture and industry (excluding construction) as tradable, while the rest of the economy – as the nontradable sector. We compute labour productivity in the tradable sector as:

$$LP_t^T = \frac{GVA_t^T}{L_t^T} \tag{12}$$

with GVA– gross value added index, L – employment. The GVA is indexed as:

$$GVA_t^T = \prod_i (GVA_t^i)^{\lambda_i} \tag{13}$$

where GV_A^{ti} is the volume index of seasonally adjusted gross value added in i-th sector classified as tradables, over the period 2000–2013. The employment is simply cumulated over sectors as a number. We treat the nontradable sector analogously to (12)–(13). Finally, our proxy is expressed so that its growth shoul! lead to appreciation:

[21] P.R. Lane, G.M. Milesi-Ferretti, *The Transfer Problem Revisited: Net Foreign Assets and Real Exchange Rates*, "The Review of Economics and Statistics" 2004, vol. 86(4), November, pp. 841–857.

[22] M. Rubaszek, op.cit.

$$BS_t^D = \frac{LP_t^{D,T} / LP_t^{D,NT}}{LP_t^{EA,T} / LP_t^{EA,NT}} \tag{14}$$

A similar relative expression defines the terms of trade (TOT) ratio:

$$TOT_t^D = \frac{P_t^{D,Ex} / P_t^{D,I}}{P_t^{EA,Ex} / P_t^{EA,I}} \tag{15}$$

where P_t^I – (overall) imports price indices, P_t^{Ex} – (overall) exports price indices. For this variable, the expected sign is ambiguous and depends, in line with the Marshall-Lerner conditions[23], on the price elasticities of individual foreign trade streams. A similar, but not so much ambiguous meaning can be attributed to an individual good – crude oil, which is definitely an import good in the analysed CEE countries, whose real price (OIL; Brent, PPI-deflated, in domestic currency) is also included in one version of the model:

$$OIL_t^D = ER_t^D \cdot BRENT_t / P_t^D \tag{16}$$

One could expect depreciation once oil becomes more expensive, as a mechanism of correcting the deficit that could switch on in such a situation. However, this is not a prefect proxy either, as this depreciation should not necessarily be channelled against the euro (in fact, the euro area is to a large extent a net importer of oil as well).

Table 1 contains the results of unit root tests for individual variables. There is not much doubt left as regards GG debt, HBS effect and oil price being I(1). On the other hand, the real interest rate disparity and GG deficit robustly seem to be stationary (see Benassy-Quere et al.[24] for similar results). Doubtful cases (I(0) or I(1)) are the real exchange rate, NFA and TOT. However, neither of these results contradicts the necessary condition for conintegration, i.e. there are no considerable symptoms of I(2)-ness and there is more than one I(1) variable.

[23] A.P. Lerner, *The diagrammatical representation of cost condition in international trade*, "Economica" 1934, vol. 1, pp. 319–334; A.P. Lerner, *The diagrammatical representation of cost condition in international trade*, "Economica" 1934, vol. 1, pp. 319–334.

[24] A. Bénassy-Quéré, S. Béreau, V. Mignon, *How Robust are Estimated Equilibrium Exchange Rates? A Panel BEER Approach*, CEPII Working Paper 2008-01, CEPII, March 2008.

Bearing this in mind, as well as the standard reservations as regards the limited power of panel unit root tests[25], we proceed to check the sufficient condition for cointegration for the entire set of our variables with the real exchange rate indicated as the dependent variable. The set of tests proposed by Pedroni[26] indicates our set I (with deficit and oil price) as cointegrated in 4 out of 7 cases (at significance level 0.1; see Table 2), while set II is cointegrated according to 3 out of 7 tests. Benassy-Quere et al.[27] consider such mixed results as a widespread phenomenon, and – given the properties of the tests – as an indication towards rejection of the null. This is confirmed by the Kao test[28] at any significance level.

Table 1. Panel stationarity tests

D() – differencing		Common unit root	Individual unit roots		
		Levin, Lin & Chu	Im, Pesaran and Shin	ADF– Fisher Chi-square	PP – Fisher Chi-square
RER	Stat.	–1.771	–1.462	13.035	12.493
	P-val	0.038	0.072	0.111	0.131
D(RER)	Stat.	–10.362	–12.394	119.879	119.459
	P-val	0.000	0.000	0.000	0.000
RIR	Stat.	–2.421	–5.649	47.742	30.017
	P-val	0.008	0.000	0.000	0.000
DEF	Stat.	–1.515	–4.075	44.415	18.911
	P-val	0.065	0.000	0.000	0.015
DEBT	Stat.	–0.788	0.134	5.627	1.618
	P-val	0.215	0.553	0.689	0.991

[25] G.S. Maddala, *On the Use of Panel Data Methods with Cross-Country Data*, "Annales d'économie et de statistique" 1999, no. 55–56; G.S. Maddala, S. Wu, P.C. Liu, *Do panel data rescue the purchasing power parity (PPP) theory?*, in: *Panel Data Econometrics: Future Directions*, eds J. Krishnakumar, E. Ronchetti, North-Holland, Amsterdam 2000, pp. 35–51; B.H. Baltagi, *Econometric Analysis of Panel Data*, John Wiley & Sons, New York 2005.

[26] P. Pedroni, *Critical values for cointegration tests in heterogeneous panels with multiple regressors*, "Oxford Bulletin of Economics and Statistics" 1999, vol. 61 (S1), pp. 653–670; P. Pedroni, *Panel cointegration. Asymptotic and finite sample properties of pooled time series tests with an application to the PPP hypothesis*, "Econometric Theory" 2004, vol. 20, pp. 597–625.

[27] A. Bénassy-Quéré, P. Duran-Vigneron, A. Lahrèche-Revil, V. Mignon, *Burden Sharing and Exchange-Rate Misalignments within the Group of Twenty*, Working Papers 2004-13, CEPII Research Center, 2004; A. Bénassy-Quéré, A. Lahrèche-Révil, V. Mignon, *World Consistent Equilibrium Exchange Rates*, Working Papers 2006-20, CEPII Research Center, 200602; A. Bénassy-Quéré, A. Lahrèche-Révil, V. Mignon, *Is Asia Responsible For Exchange Rate Misalignments Within The G20?*, "Pacific Economic Review" 2008, vol. 13(1), pp. 46–61.

[28] C. Kao, *Spurious regression and residual-based tests for cointegration in panel data*, "Journal of Econometrics" 1999, vol. 90(1), May, pp. 1–44.

D() – differencing		Common unit root	Individual unit roots		
		Levin, Lin & Chu	Im, Pesaran and Shin	ADF– Fisher Chi-square	PP – Fisher Chi-square
D(DEBT)	Stat.	0.451	–2.458	20.159	22.344
	P-val	0.674	0.007	0.010	0.004
NFA	Stat.	–0.803	–1.357	11.817	36.936
	P-val	0.211	0.087	0.160	0.000
D(NFA)	Stat.	12.946	–3.355	26.770	94.204
	P-val	1.000	0.000	0.00	0.000
Log HBS	Stat.	–1.261	–0.395	8.473	8.191
	P-val	0.104	0.347	0.389	0.415
D(Log HBS)	Stat.	–8.994	–11.972	118.797	145.679
	P-val	0.000	0.000	0.000	0.000
Log TOT	Stat.	0.300	–2.318	29.019	56.529
	P-val	0.618	0.010	0.000	0.000
D(Log TOT)	Stat.	–14.907	–18.887	78.647	117.247
	P-val	0.000	0.000	0.000	0.000
Log OIL	Stat.	–0.181	0.191	5.081	5.484
	P-val	0.428	0.576	0.749	0.705
D(Log OIL)	Stat.	–7.951	–11.474	113.292	109.955
	P-val	0.000	0.000	0.000	0.000

Testing regressions with constant, without trend, with lag length selected automatically based on the Schwarz criterion. Estimated long-term variance using the Bartlett weights (window length: 4).

Source: own calculations.

Table 2. Panel cointegration tests

Test		Model I		Model II	
		Test statistic	p-value	Test statistic	p-value
Pedroni (1999, 2004)	Panel v-Statistic	–0,320	0,625	0,713	0,271
	Panel rho-Statistic	–0,397	0,346	–0,167	0,369
	Panel PP-Statistic	–1,893	0,029	–0,955	0,086
	Panel ADF-Statistic	–2,291	0,011	–1,088	0,030
	Group rho-Statistic	0,215	0,585	0,235	0,593
	Group PP-Statistic	–1,829	0,034	–1,256	0,105
	Group ADF-Statistic	2,209	0,014	1,644	0,050
Kao (1999) ADF		2,586	0,005	2,613	0,005

Source: own calculations.

Our cointegrating vectors, estimated with FM-OLS, are summarized in Table 3. In both models, all variables in consideration are significant at the level 0.01 and normality of the residuals is not rejected. In line with our expectations, increasing real interest rate disparity and the HBS-effect measure both lead to appreciation of a currency. At the same time, GG deficit and debt – when growing – bring about depreciation.

As regards the ambiguous cases, terms of trade improvement (i.e. relative export price growth in excess of relative import price growth) acts towards appreciation, i.e. suggests improvement in the CA after TOT growth. However, the growth of oil prices alone leads to depreciation – in line with our expectations. Also, NFA growth leads to appreciation, which is consistent with the theory[29] rather than with the stylized facts of financial deepening. Both coefficients (TOT, NFA) take opposite signs in our study than in the findings of Będza-Bojanowska and MacDonald for Poland[30].

Table 3. BEER model – cointegrating vectors estimated with FM-OLS (dependent variable: log real exchange rate)

Variable	Model I		Model II	
	Coefficient	p-value	Coefficient	p-value
RIR	–0.355	0.000	–0.365	0.000
DEF	0.195	0.000	–	
DEBT	–		0.181	0.000
NFA	–0.212	0.000	–0.138	0.000
Log HBS	–0.273	0.000	–0.254	0.000
Log TOT	–1.078	0.000	–0.986	0.000
Log OIL	0.177	0.000	–	
JB (p-value)	0.152		0.731	

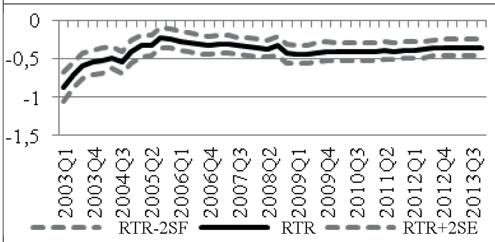
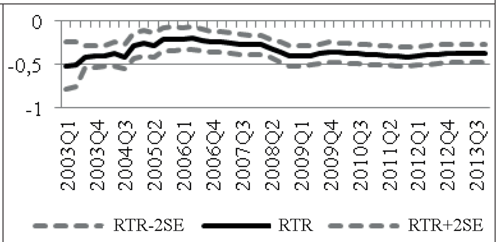
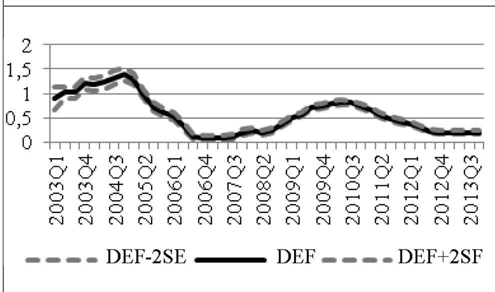
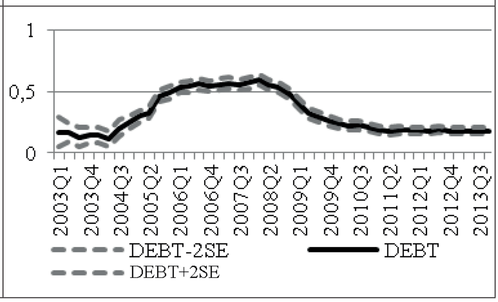
Source: own calculations.

We test for the robustness of our results in two dimensions, both in terms of time stability and of possible heterogeneity between countries. Figure 1 presents recursive estimates of the coefficients in two model versions, along with +/– two standard errors. In this analysis, none of the coefficients changes sign or evolves so that a variable drops or regains significance over the crisis years (except the initial period of unreliably short samples). In fact, in both models, the real interest rate disparity seems to exert the most stable influence. What might be seen as surprising is the fact that the NFA coefficient is gradually

[29] M. Rubaszek, op.cit.
[30] J. Bęza-Bojanowska, R. MacDonald, op.cit.

decreasing in magnitude, contrary to a possible explanation of transition from „financial deepening" mode during convergence up to the „theoretical" mode of developed economies. Similar conclusions about the HBS proxy are model-dependent and likely associated with some fluctuations in debt coefficient over the sample period. The same is the case for deficit, which implies that both proxies for macroeconomic risk are imperfect and could be perceived differently at different times.

Another dimension to test is the choice of the sample countries. For this purpose, we present the same estimates as in Table 3, but using a panel of 3 countries, i.e. excluding every country individually. In this case, we detect the following cases of inconsistence. Firstly, dropping Poland from the sample yields insignificant estimates of the real interest rate parameter. Secondly, skipping Romania inverts the sign of the terms of trade (which may explain the difference between our estimates and the time-series findings by Bęza-Bojanowska and MacDonald[31] for Poland). Additionally in both cases, this leads to magnitude's reduction or change of the sign of the risk premium parameter (depending on the model version). Thirdly, in the model without the Czech Republic, the proxy for Balassa–Samuelson effect becomes insignificant. Dropping Hungary, in turn, does not cause any qualitative change.

Model I	Model II
Sample ranges from the beginning to the indicated period; N×T from 40 to 200	
	
	

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[31] Ibidem.

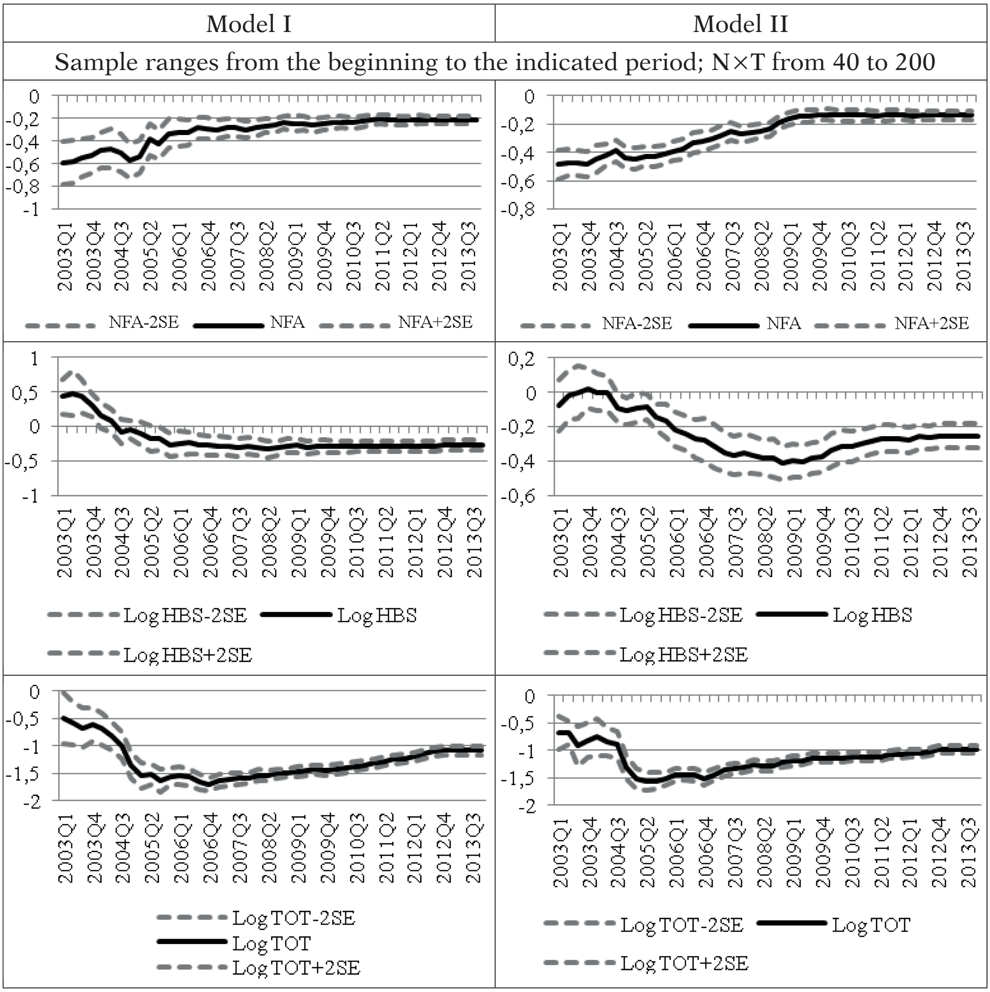


Figure 1. Recursive coefficient estimates

Source: own calculations.

Table 4. Estimates with incomplete set of countries – model I

Variable name	Full sample		Without CZ		Without HU		Without PL		Without RO	
	coef	p-val	coef	p-val	coef	p-val	coef	p-val	coef	p-val
RIR	–0.355	0.000	–0.179	0.005	–0.530	0.000	0.066	0.323	–0,670	0,000
DEF	0.195	0.000	0.197	0.000	0.846	0.000	–0.318	0.000	0.286	0.000
NFA	–0.212	0.000	–0.219	0.000	–0.187	0.000	–0.418	0.000	–0.313	0.000
Log HBS	–0.273	0.000	0.007	0.898	–0.287	0.000	–0.352	0.000	–0.257	0.000
Log TOT	–1.078	0.000	–1.224	0.000	–1.151	0.000	–1.239	0.000	0.406	0.000
Log OIL	0.177	0.000	0.072	0.120	0.213	0.000	0.120	0.004	0.160	0.000

Source: own calculations.

Table 5. Estimates with incomplete set of countries – model II

Variable name	Full sample		Without CZ		Without HU		Without PL		Without RO	
	coef	p-val	coef	p-val	coef	p-val	coef	p-val	coef	p-val
RIR	–0.365	0.000	–0.182	0.005	–0.520	0.000	0.020	0.764	–0.676	0.000
DEBT	0.181	0.000	0.238	0.000	0.359	0.000	0.196	0.000	–0.061	0.000
NFA	–0.138	0.000	–0.147	0.000	–0.086	0.000	–0.337	0.000	–0.338	0.000
Log HBS	–0.254	0.000	–0.004	0.931	–0.347	0.000	–0.446	0.000	–0.140	0.001
Log TOT	–0.986	0.000	–1.173	0.000	–0.956	0.000	–1.209	0.000	0.501	0.000

Source: own calculations.

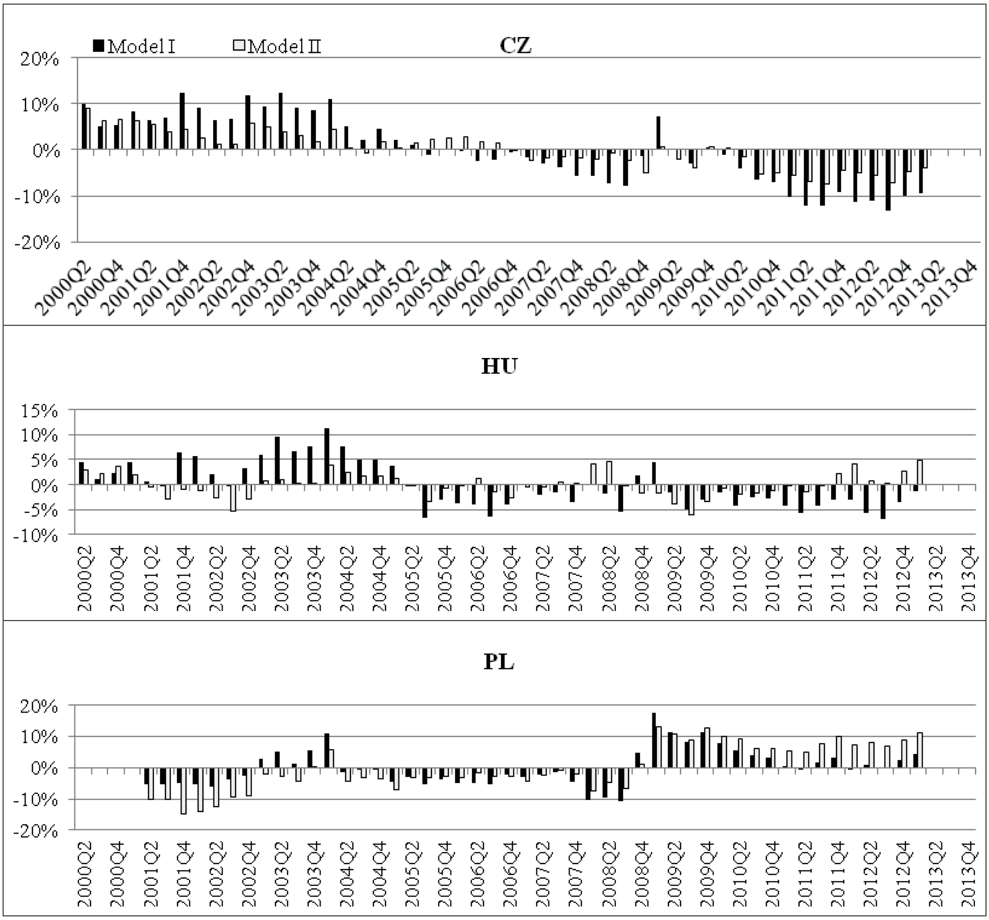
Table 6. Cointegrating vector estimated for Poland (FMOLS). Dependent variable: log RER

Variable	Model I		Model II	
	Coefficient	p-value	Coefficient	p-value
RIR	–0.513***	0.000	–0.462	0.000
DEF	0.335***	0.001	–	
DEBT	–		0.132***	0.000
NFA	–0.467***	0.000	–0.494***	0.000
Log HBS	0.238*	0.082	0.283**	0.016
Log TOT	–0.641***	0.000	–0.647***	0.000
Log OIL	0.127	0.294	–	

Source: own calculations.

Finally, we run yet another comparison, estimating our cointegrating vector only for Poland (FMOLS with time series). Cointegration is confirmed at the significance level of 0.05, but the real oil price turns out to be insignificant (see Table 6). Moreover, in both specifications the HBS proxy takes sign opposite than expected, yielding an economically unacceptable result of HBS effect significantly causing depreciation.

Using our model, we compute misalignments of the real exchange rate from the fundamentals as residuals from the cointegrating relationship (Figure 2). For Poland, this shows undervaluation of the Polish real exchange rate of 10–15% in 2009, still persisting into 5–10% towards the end of the sample (2013Q4). At the same time, the Czech rate seems to be overvalued by the same magnitude, and the Hungarian and Romanian rate – valued in line with the fundamentals.



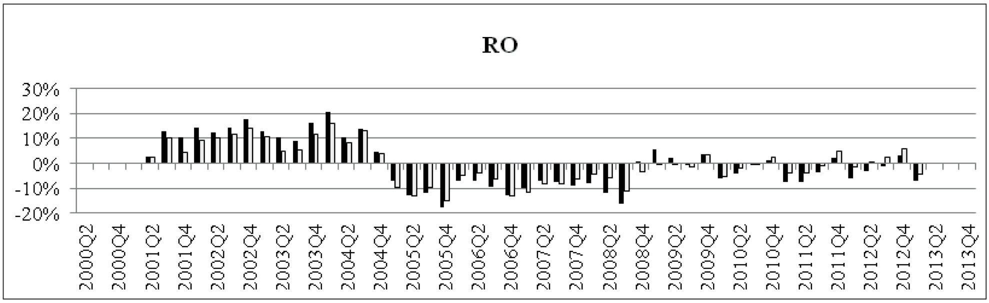


Figure 2. Currency misalignments in CEE countries, 2000–2013
Source: own calculations.

4. Conclusions

Panel-based estimates of cointegrating relationships, leading to calculation of behavioural equilibrium exchange rates and related misalignments, turn out to be more precise than in the case of time series model based on (partly) post--crisis sample. This is in particular true for the quarterly frequency, at which better (or less noisy) data is available. All standard determinants considered in the literature (real interest rate disparity, HBS effect proxy, general government deficit or debt, NFA and terms of trade) are significant at the significance level of 0.01 and relatively stable. Two things should be mentioned as regards the relations to the existing literature. Firstly, contrary to some previous findings, increase in NFA is found to cause appreciation; secondly, the same is true for increase in TOT.

One possible challenge for future research is dealing with some heterogeneity between individual countries detected in the sensitivity analysis. We find the proposed set of countries as homogenous as possible for this purpose (in terms of period, region and monetary regime), but there may still be idiosyncratic factors to control. Removing an individual country indeed poses a challenge to robustness in a "small N" situation like here; yet, a greater challenge in this situation is to convincingly increase N. All in all, however, the panel approach turn out to be promising in the light of future need for research prior to setting central parities for ERM II and conversion rates from national currencies to the euro.

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