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Long-run real exchange rate determinants: Evidence from eight new EU member states, 1993–2003

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In this paper, we estimate bilateral equilibrium real exchange rates for a group of eight new EU member states against the euro, using new and sophisticated panel-cointegration techniques. We document a stable significant positive link between productivity levels and the corresponding real exchange rate levels and a stable significant and negative link between openness and the real exchange rate. We find real exchange rate misalignments to be small and weakly mean-reverting. In the context of entry into ERM-II and EMU for most of these countries over time, the results stress the importance of allowing countries to adjust to inflation pressure and real exchange rate appreciation, either through nominal appreciation or temporarily higher domestic inflation. *Journal of Comparative Economics* **35** (1) (2007) 87–107. University of Maastricht, Department of Economics, PO Box 616, 6200 MD Maastricht, The Netherlands; Utrecht School of Economics, University of Utrecht, Vredenburg 138, 3511 BG Utrecht, The Netherlands.

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1. Introduction

As of May 1, 2004, ten countries have joined the European Union: Cyprus, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Slovenia, and Slovakia. Previously, most of these countries had centrally planned economies. From the early 1990s onward, they have gone through a severe transition process towards a market economy. In terms of nominal exchange rate regimes and actual exchange rate developments, large differences can be noted across these countries over the past ten years.

Some started with a relatively fixed exchange rate regime and switched to (managed) floating at some point, others worked the other way around with a relatively floating regime in the early nineties and a move to more fixed regimes in the second half of that decade. All of them still have their own currency and monetary autonomy. However, they share the same long-run perspective of participating in the Economic and Monetary Union (EMU) and adopting the euro as the common currency. Estonia, Lithuania and Slovenia already entered ERM-II in June 2004, while Cyprus, Malta, Latvia and Slovakia entered in 2005. Other countries will follow in due time.

A major issue in this respect is the development of the *real* exchange rate relative to the euro. Most importantly, trend-like real appreciations caused by for example the Balassa–Samuelson effect either require nominal exchange rate appreciations which are hard to reconcile with a fixed exchange rate regime or domestic inflation in excess of inflation in the euro area which is inconsistent with a strict interpretation of the Maastricht criteria. More generally, knowledge of real exchange rate determinants may be of help in assessing the readiness of each country to move to the EMU.

In this paper, we analyze the fundamentals behind the real exchange rate for eight of these ten new member states (NMS) for the period from 1993 to 2003.[1] In particular we are interested in the long-run components of the real exchange rate. Although many studies exist for particular countries, e.g. the Czech Republic, Hungary, and Poland, only few are available for the whole group of the new EU members. In addition, the econometric methodology varies across existing work, being determined by the availability of either time-series or panel data. Here, we extend the existing literature through a uniform and sophisticated panel cointegration analysis for two homogeneous "blocks": the three Baltic countries on the one hand and the five Central European countries on the other. In our analysis, we combine determinants of the external and the internal real exchange rate, respectively, and conclude that both have caused the real exchange rate to appreciate, albeit to a different extent in the two blocks.

The remainder of the paper is structured as follows. In Section 2, we develop the concept of the real exchange rate and present the theoretical model as well as its empirical specification. We present and motivate the data and the choice of the panel cointegration technique in Section 3. The corresponding empirical results are discussed in Section 4. Section 5 concludes.

2. The real exchange rate: theoretical concepts and empirical application

2.1. Theoretical concepts

In our paper, we exclusively use the bilateral real exchange rate of each country's currency against the euro rather than the country's overall effective real exchange rate. The latter measures

[1] Due to a lack of data, Malta and Cyprus had to be excluded from the analysis.

a country's competitiveness against the rest of the world. Here, however, we are especially interested in the relation between the euro area and the different new EU members, in light of their future entry into the euro area and adoption of the euro themselves.[2]

We define the nominal exchange rate for each new member country i (E_i) as the price of one unit of the country's currency in terms of the euro. A rise of the exchange rate then indicates an appreciation of the new member state's currency. The corresponding real bilateral exchange rate (Q_i) is the relative price of a standard basket of goods in country i relative to the price of the same basket in the euro area. It reflects the competitive position of each country relative to the euro area. The price levels in the euro area and the new member state are denoted by P^* and P, respectively. Then, the real exchange rate is defined as

$$Q_i = (E_i P_i / P^*). \tag{1}$$

If all goods are tradable (and homogeneous across countries) and baskets of goods are identical between countries, and if there are no market frictions and no trade barriers, the strict version of the purchasing power parity (PPP) hypothesis holds and Q_i will equal unity. Arbitrage then ensures that identical baskets are equally expensive across countries. Note that in practice, we use price indexes rather than actual prices across countries to compute the real exchange rate.

Extending the model to accommodate the existence of both tradable and non-tradable goods, we assume that a country's price index P is a geometrically weighted average of the price indexes of tradable and non-tradable goods. The weights are given by the share of the tradable goods (α) and non-tradable goods ($1-\alpha$) in the total added value of a country. Equation (1) then can be rewritten as:

$$Q_i = E_i \cdot \left(P_i^T / P^{T*}\right) \cdot \left(P_i^N / P_i^T\right)^{(1-\alpha)} \cdot \left(P^{T*} / P^{N*}\right)^{(1-\alpha)}. \tag{2}$$

Using a standard logarithmic transformation, we arrive at Eq. (3), with lower case symbols denoting logarithms

$$q_i = e_i + \left(p_i^T - p^{T*}\right) + (1-\alpha)\left[\left(p_i^N - p_i^T\right) - \left(p^{N*} - p^{T*}\right)\right]. \tag{3}$$

From Eq. (3), one immediately infers that the (logarithmic) real exchange rate is a combination of three factors:

(i) the real exchange rate for tradable goods, $e_i + (p_i^T - p^{T*})$,
(ii) the price ratio of non-tradable goods and tradable goods in the respective NMS, $(p_i^N - p_i^T)$, and
(iii) the price ratio of non-tradable goods and tradable goods in the euro area, $(p^{N*} - p^{T*})$.

We refer to the first variable as the external real exchange rate q_{1i}, while the other two terms represent the internal real exchange rates q_{2i} and q_2^* for the NMS and the euro area, respectively. Note that the overall real exchange rate now can deviate from unity even under perfect arbitrage across tradable goods due to differences between internal real exchange rates (non-tradables

2 In practice, for each NMS under consideration, the bilateral real exchange rate and the effective real exchange rate move closely together and are strongly correlated over the period 1993–2003. Two factors account for this. First, the euro area is the dominant trading partner for each of these NMS, especially for the Central European ones, giving the bilateral euro rate a large weight in the effective rate. Second, the transformation processes towards a market economy in each NMS have a strong and similar effect on both real exchange rates.

prices). Using the definitions of the internal and external real exchange rate, Eq. (3) can be reformulated as:

$$q_i = q_{1i} + (1-\alpha)(q_{2i} - q_2^*). \tag{4}$$

We now turn to a discussion of the determinants of the internal and external exchange rate, respectively. In the empirical work, we will use a reduced form that is based on Eq. (4) to explain movements in the overall real exchange rate q_i. We focus on the overall exchange rate rather than its two components, the internal and external exchange rate, as it corresponds most directly to the questions each NMS faces with respect to the appropriate domestic monetary policy and the nominal exchange rate regime. While entry in ERM-II and eventual adoption of the euro requires low domestic CPI inflation, upward pressure on the CPI-defined overall real exchange rate may induce higher inflation at home when the nominal exchange rate is relatively fixed. Obviously, whether a real exchange rate appreciation predominantly turns up in the internal or external real exchange rate is an interesting issue in its own right. However, such analysis is outside the scope of this paper.[3]

2.2. Empirical implementation

To develop the reduced-form model, we elaborate on the external and on the internal part of the real exchange rate. We assume that the external real exchange rate (q_1) takes care of external equilibrium, while the internal real exchange rate (q_2) serves to equilibrate supply and demand for domestic non-tradables, which is the internal equilibrium. For a similar approach see, for example, Alberola et al. (1999), Égert and Lahrèche-Révil (2003), and Kim and Korhonen (2005).

We note first that a country's current account position is central to the concept of external equilibrium. In the literature, external equilibrium is generally defined as a sustainable current account position (as a percentage of GDP). Typical determinants of the current account are the external real exchange rate, or for that matter the terms of trade,[4] openness and stock variables like a country's net foreign asset position.[5] Obviously, for the current account to stay close to a sustainable level, its joint determinants should not be able to drive it out of its equilibrium position too far. Put differently, non-stationary driving forces of the current account should jointly be stationary (cointegrated). In this paper, we assume that the external real exchange rate and the degree of (relative) openness of a country jointly determine the current account.[6] Consequently, the external real exchange rate (q_1) then is a function of the current account and the degree of openness. Due to the stationary character of the current account, the only remaining long-run determinant of the external real exchange rate to be included in the empirical specification is

[3] In the literature, the ratio of CPI over PPI is often used as a proxy for the internal real exchange rate, while the PPI-defined real exchange rate approximates the external real exchange rate. This allows separate analysis of the two real exchange rate components. The reliability of such an analysis depends on the precision with which the external and internal exchange rate can be identified by these proxies.

[4] Égert and Lahrèche-Révil (2003) explicitly use a terms of trade variable in the external equilibrium equation. Note, however, that the terms of trade measured as the relative price of domestic exports and domestic imports and the external real exchange rate are hardly independent.

[5] Edwards and Savastano (2000) extensively discuss the role of openness. We refer to Alberola et al. (1999) for stock-flow models of real exchange rate appreciation. See also Frait and Komárek (2002).

[6] Due to measurement problems, we abstract from using stock variables like a country's net foreign assets (or foreign debt).

the openness variable. Related to this is the argument made by De Broeck and Sløk (2001) that openness measures the responsiveness of the real exchange rate to monetary and fiscal shocks, which is larger the more closed the economy is.

To maintain a sustainable (stationary) current account, the external real exchange rate is assumed to offset trends in the current account due to changes in relative openness. Assume that the relevant starting point for a representative NMS is one of constrained trade relations. Then, the subsequent increase in openness, representing among others a cut in tariff and non-tariff protection, will trigger an increased domestic demand for foreign tradable goods and a larger current account deficit. To maintain a sustainable current account, a real exchange rate depreciation would be required. Then, rising openness implies a real depreciation. Alternatively, however, the increase in openness can work the other way as it opens the way for higher foreign (euro) demand for domestic (NMS) products and correspondingly higher exports. For NMS, closer integration with the EU not only raises its imports but its exports as well. Empirically, the sign of the openness variable is therefore ambiguous.

For a discussion of the determinants of the internal real exchange rate, we start with the Balassa–Samuelson (BS) hypothesis, as first formulated in the seminal work of Balassa (1964) and Samuelson (1964). According to the BS hypothesis, less developed countries typically experience a structural appreciation of the internal real exchange rate when they catch up with more developed countries. In the catching-up process, productivity in the domestic tradable goods sector will increase relative to that in the non-tradable goods sector. Under the assumptions that wage setting in the tradable goods sector dominates wage setting in the non-tradable goods sector and that wages in the tradable and non-tradable sector will equalize due to domestic labor mobility, prices of domestic non-tradable goods increase relative to domestic prices of tradable goods. Along the lines of BS, we therefore hypothesize a positive link between the productivity differential and the real exchange rate.

In addition, Baumol and Bowen (1966) have been the first to argue that in the catching-up process an increase in demand for non-tradables relative to tradables may occur, causing an appreciation of the internal real exchange rate as well. Here, the argument is that the composition of the standard consumption bundle will shift in the direction of non-tradable goods (and services) with an increase in income and wealth. In practice, however, the catching up process of developing countries may simultaneously be an opening up process with the rest of the world. This may result in a (temporary) increase in demand for tradables relative to non-tradables and, thus, may give rise to an opposite effect. Again, the sign of the relation is therefore ambiguous.[7]

Based on the above discussion of the determinants of the external and internal real exchange rate, we rewrite Eq. (4) as follows:

$$q_t = e_0 + e_1 \text{open}_t + e_2 \text{prod}_t + e_3 \text{demand}_t + u_t, \tag{5}$$

[7] Some empirical studies include monetary variables like interest rates and money supply as real exchange rate determinants (Randveer and Rell, 2002; Šmídková et al., 2003; Lommatzsch and Tober, 2004). Crespo-Cuaresma et al. (2005) use monetary variables as nominal exchange rate determinants. The limited interest in monetary variables is attributable to two aspects. Firstly, in the long run the real exchange rate is assumed to be independent of nominal variables such as the money supply. Secondly, the real interest rate only clears the goods market in a large, but not in a small domestic economy. In the present framework, it would only be sensible to include the real interest rate if we were to discuss short-term deviations from the long-run equilibrium. Given these properties, we also choose to exclude money and interest rates from the analysis.

where *open* stands for openness; *prod* stands for productivity and *demand* is a demand variable.[8] We expect e_2 to be positive, while the sign of e_1 and e_3 is ambiguous. We will estimate this equation using a panel-cointegration approach in Section 4. We use the level of the bilateral real exchange rate as the dependent variable. All independent variables are measured relative to the euro area.

3. Data and empirical methodology

In this section, we elaborate on the data and research methodology. Section 3.1 contains a brief description of the choice of explanatory variables and their precise definition. In Section 3.2, we briefly discuss the panel cointegration method that we use to estimate Eq. (5).

3.1. Variables and data selection

To empirically implement our exchange rate model, we first note that for most of the countries under consideration, it is impossible to find consistent data before 1993. Consequently, we have approximately 10 years of data. The use of quarterly data imposes an additional restriction as a number of variables are only available at an annual frequency. If not stated differently, the data are collected from the IMF International Financial Statistics.[9] Except for the indexes, the data are originally expressed in millions of national currency. All time series are seasonally adjusted. Nominal bilateral exchange rates are deflated by the ratio of the consumer price index in each NMS and the euro area to obtain real exchange rates.

We define the productivity variable *prod* for most countries as the logarithm of the ratio of industry production and industry employment relative to the same variable for the euro area as a whole. Only for Estonia, where the ratio of industry production and industry employment is unavailable, the logarithm of the ratio of GDP over total employment is used. In that case, GDP over total employment from the euro area is taken as benchmark. In this choice, we follow a large segment of the literature.[10] In related research, output per capita is sometimes used as a measure of overall productivity in an economy (Šmídková et al., 2003; Lommatzsch and Tober, 2004; Crespo-Cuaresma et al., 2005; Frait and Komárek, 2002; Dobrinsky, 2003; MacDonald and Wójcik, 2004). However, in our view, output per capita is at most a second-best approach for productivity.

In the recent literature, many different variables, like for example consumption, investment, government expenditures, and GDP variables, have been used to approximate demand pressures, see Égert and Lahrèche-Révil (2003), Frait and Komárek (2002), Filipozzi (2000), Kim and Korhonen (2005). In particular, overall government expenditures or government consumption are often used because these best capture the expenditure bias towards non-tradable goods, see

[8] Note that Eq. (5) represents a so-called Behavioral Equilibrium Exchange Rate Model (BEER). We refer to Maeso-Fernandez et al. (2005) for an overview and discussion of the different types of equilibrium exchange rate models used in the literature.

[9] The data are accessed via the IFS on-line service available at http://www.ifs.apdi.net/imf/logon.aspx.

[10] The existing literature employs a wide variety of other proxy variables for productivity differentials, such as (relative) wages, relative consumer versus producer prices, total factor productivity, or measures of industry structure. Data availability typically precludes tests regarding the relevance of these variables. The exception concerns the availability of data regarding the price ratio. In our view, our *prod* variable is a more direct and exogenous measure of productivity than the price ratio.

Bergstrand (1991). On the other hand, over the transition period 1993–2003, most of the governments involved downsized substantially due to privatization and the switch from a centrally planned economy to a market economy. It is unclear how to interpret the government expenditure data in this respect. Here, we propose to investigate the robustness of the demand indicator by using three different specifications. We define *demand* alternatively as

(i) the logarithm of the ratio of government consumption expenditures over GDP for each NMS relative to the euro area,
(ii) the logarithm of the ratio of private consumption expenditures over GDP for each NMS relative to the euro area, and
(iii) the logarithm of the ratio of total (government plus private) consumption expenditures over GDP for each NMS relative to the euro area.

Note that neither variable allows for the theoretically preferred exact distinction between the demand for tradable goods and non-tradable goods.

The use of the openness variable is quite standard in the literature as a proxy for increased international integration and a decline of tariff and non-tariff barriers. However, in the literature there is no consensus on the "best" definition for openness. Égert and Lahrèche-Révil (2003), for instance, use an absolute measure for each country (defined as the sum of exports and imports divided by GDP), while De Broeck and Sløk (2001) use the relative version (defined as the ratio of country i's imports plus exports as a percentage of its GDP and OECD imports plus exports over OECD GDP). Note that in a cross section or panel approach both relative and absolute measures of openness may capture differences in size among the sampled countries as smaller countries tend to have higher import and export percentages in terms of GDP than larger ones. As a result, no uniform steady state equilibrium value for relative or absolute openness should be expected across countries. Because of the interdependence with a size effect, the empirical results with respect to the *open* variable should be interpreted with some caution.

Here, we opt for a relative version, defined as the logarithm of the sum of exports and imports as a ratio of GDP in each country relative to the same ratio in the euro area. The main argument is that it is not only the degree of trade liberalization and tariff reduction of an NMS that determines its openness but also the degree of trade liberalization and tariff reduction of its trading partners, in this case the euro area.[11]

3.2. *Panel cointegration*

We next turn to the econometric methodology. The analysis focuses on the determinants of the long-run real exchange rate level across countries. In the literature, several econometric methods have been used in order to estimate such BEER. Basically, exchange rate developments have been analyzed via time-series (see for example Lommatzsch and Tober, 2004; Coudert and Couharde, 2003; Frait and Komárek, 2002; Filipozzi, 2000; and Égert and Lahrèche-Révil, 2003) or panel-

[11] We also applied an absolute version of the openness measure defined as the logarithm of the sum of each NMS imports and exports with respect to the EU as a percentage of its GDP. This introduces a measurement error since some of the NMS have strong trade links with the non-euro area EU members and potentially biases the results. Unreported results that are available from the authors show that the parameter estimates for openness are quite sensitive to the definition used. However, the sensitivity of the productivity and demand coefficients for changes in the openness variable is quite low.

data methods (Kim and Korhonen, 2005; Šmídková et al., 2003; Rahn, 2003; and De Broeck and Sløk, 2001).

Although a time-series analysis best allows for the estimation of country-specific exchange rate developments, it requires a sufficient number of observations per country. Unfortunately, for the NMS, the period over which adequate data are available is quite short and typically starts in 1993 at the earliest. Therefore panel-data analysis is used more often in studies comparable to ours. This method increases the number of observations leading to consistent estimators. However, the estimation results can only be interpreted as a 'common' estimator for the group of countries. Thus, the benefit of an extended database comes at the cost of a loss in country-specific estimates. Note that in our specific case the cross-sectional dimension is quite small with a maximum of eight NMS in one panel. Clearly, this limits the increase in the number of observations. Moreover, the use of the panel methodology introduces the potential problem of cross-sectional dependence between countries.

In the context of long-run exchange rate determination, we have to take into account that real exchange rates as well as their determinants are typically found to be non-stationary, that is, they lack a fixed mean value to which they tend to return over time, so that a cointegration analysis is required. In a cross-country analysis, this leads to a panel-cointegration method. The use of normal OLS techniques will lead to spurious regression and specific panel-cointegration techniques have to be used. Kao and Chiang (2000) have shown that OLS in panel-cointegrated models is asymptotically normal but biased. Even the bias-corrected OLS estimator does not improve the OLS estimator in general (Chen et al., 1999). Alternative methods are then necessary.

Phillips and Moon (2000) show that in the case of homogeneous and near-homogeneous panels,[12] the long-run coefficient can be obtained by a pooled fully modified (FM) estimator (see also Pedroni, 2000). This method is non-parametric as it employs kernel estimators of the nuisance parameters that affect the asymptotic distribution of the OLS estimator. It tackles the possible problem of endogeneity of the regressors as well as the autocorrelation of residuals. Kao and Chiang (2000) prefer to extend the work of Stock and Watson (1993) and Saikkonen (1991) and propose a pure panel dynamic least square estimator (DOLS). This estimation procedure is parametric and has the advantage of computing convenience. In finite samples, it is shown by Kao and Chiang (2000) that the FM estimator does not improve over the OLS in general and the DOLS appears to outperform the other methods especially if fixed effects are included. In our empirical section, we will focus on DOLS. The result of the DOLS estimation is a set of long-run coefficients that relate the selected fundamentals to the real exchange rate.

4. Empirical results

In the empirical analysis, we first test for the non-stationarity of the (logarithmic) real exchange rate series and the various explanatory variables as well as for the presence of one single homogeneous cointegration relationship among them. Panel unit root tests proposed by Levin, Lin and Chu (2002) (LLC hereafter) and Im, Pesaran and Shin (2003) (IPS hereafter) are the most popular tests in empirical studies. Nevertheless, they are shown to be inconsistent in the presence of cross-sectional dependence, as well as when N (the cross-sectional dimension) is small with respect to T (the time dimension). Several alternatives and modifications have been recently proposed in the literature, see Gutierrez (2003), Hurlin and Mignon (2005) and

[12] In heterogeneous panel cointegration, each country has its own specific cointegration relationship, as opposed to homogeneous panel cointegration where all countries share the same cointegration relationship.

Table 1
Panel-unit root tests

	LLC			IPS			IN		
	Test value	5% value	I	Test value	5% value	I	Test value	5% value	I
Baltics									
Real exchange rate	−10.12	−7.22	0	−5.36	−4.35	0	−2.50	−4.28	1
Gov't. consumption	−6.55	−7.77	1	−4.16	−4.27	1	−1.25	−3.15	1
Private consumption	−13.64	−12.37	0	−6.96	−4.75	0	−2.62	−4.20	1
Total consumption	−13.93	−10.91	0	−8.20	−5.13	0	−2.30	−4.84	1
Openness	−4.10	−8.01	1	−3.22	−3.79	1	0.16	−2.66	1
Productivity	−5.08	−7.03	1	−2.73	−3.89	1	−0.95	−3.78	1
CE5									
Real exchange rate	−4.55	−7.02	1	−3.03	−2.74	0	−1.87	−2.17	1
Gov't. consumption	−5.31	−8.10	1	−2.26	−6.03	1	0.40	−3.04	1
Private consumption	−7.77	−9.66	1	−3.36	−4.16	1	−3.10	−4.66	1
Total consumption	−6.99	−10.01	1	−3.57	−3.85	1	−2.49	−4.13	1
Openness	−4.40	−14.71	1	−1.82	−5.00	1	0.26	−5.48	1
Productivity	−4.04	−8.13	1	−2.50	−3.59	1	−0.16	−3.71	1

Notes. All variables in logarithms, an intercept has been considered in all the experiments and block size selection follows the minimum volatility rule. The 5% value columns give the non-centered critical value, the columns labeled "I" indicate the order of integration according to the test.

Breitung and Pesaran (2006) for recent surveys. In this paper, we use a subsampling approach proposed by Choi and Chue (2005). This method is applied to LLC and IPS and simulation results indicate that it behaves correctly. In addition, we also present results for an alternative test, labeled as "inverse normal panel unit root test" (IN) (Choi, 2001), which corresponds to a generalized least squares (GLS) version of the ADF test.

In Table 1 we report the test results of the three panel unit root tests together with the 5 percent critical values of the tests. For each NMS, all variables are in logarithms and relative to the euro area. In a number of cases, the tests yield conflicting results. Specifically, for private consumption and total consumption in the Baltics compared to the euro area as well as for the Baltics' real exchange rates, two out of three tests indicate stationarity. Nevertheless, in an overall perspective, we conclude that the evidence typically points to non-stationarity of the real exchange rate, productivity, openness, and the three demand variables. Therefore, continuing with the panel cointegration methods is warranted in our view.

4.1. *Estimated equilibrium real exchange rates*

We now use the DOLS method to estimate the long-run relation between bilateral real exchange rates, productivity, openness and demand, as formulated in Eq. (5). First, a joint panel for all eight NMS was estimated. However, a Hausmann test rejected the homogeneity of the cointegration relation for these eight countries.[13] Subsequently, we decided to split the total group of countries into two subgroups, based on both the difference in geographical position and in nominal and real exchange rate developments. On the one hand, we took the three Baltic countries,

[13] Estimation results for the panel of eight countries are not reported, but are available on request from the authors.

Table 2
Regression results: Baltics

	1993–2003			1995–2003		
Gov't. consumption	−0.03	–	–	−0.05	–	–
	(0.32)			(0.79)		
Private consumption	–	0.84	–	–	−0.22	–
		(7.70)			(1.98)	
Total consumption	–	–	0.39	–	–	−0.16
			(2.98)			(1.24)
Openness	−0.82	−0.67	−0.63	−0.50	−0.48	−0.46
	(11.66)	(11.55)	(10.99)	(10.26)	(11.30)	(10.43)
Productivity	0.85	1.11	0.96	0.76	0.78	0.76
	(8.98)	(13.77)	(12.04)	(17.09)	(20.09)	(18.46)
R^2	0.992	0.996	0.996	0.999	0.999	0.999
Cointegration tests (p-values)						
Full system						
Kao (1999)						
DF(ρ)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
DF(T)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
DF*(ρ)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
DF*(T)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
Fachin (2006)						
$G-t$ test bootstrap	57.80	3.60	15.00	71.80	20.50	62.50
$G-t$ test fast bootst. 1	52.70	3.00	11.20	67.10	14.40	67.20
$M-t$ test fast bootst. 1	74.10	3.40	12.10	12.90	80.00	34.50
Right-hand variables						
DF(ρ)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
DF(T)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
DF*(ρ)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
DF*(T)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
$G-t$ test bootstrap	12.90	1.80	17.50	66.40	58.50	71.40
$G-t$ test fast bootst. 1	14.00	1.80	16.20	68.60	62.40	71.30
$M-t$ test fast bootst. 1	50.60	1.90	2.50	77.10	70.30	84.70
Hausmann	20.50	18.30	19.70	50.40	61.60	64.30
Unit root residuals						
LLC	I(0)	I(0)	I(1)	I(1)	I(1)	I(1)
IPS	I(0)	I(0)	I(1)	I(0)	I(1)	I(0)
IN	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)

Notes. t-values are listed between parentheses. Note that a negative sign for a variable means that the variable has a depreciating impact on the real exchange rate. All tests are performed considering a constant and using 1000 replications. Block size selection for the cointegration test is based on $0.2T$. Results appear to be robust for other block size and are available from the authors upon request.

labeled BALTICS, and on the other hand the remaining five Central European countries, labeled CE5. Equation (5) then was estimated for each subgroup separately.

In Table 2 we first present the results for the Baltics. T-values are in parentheses below the estimated coefficients. The estimated intercept and country-specific dummies are not reported in the table. For *demand* we alternately use government consumption, private consumption and total consumption. We estimate the model both for the full sample 1993:1–2003:1 and for a shorter sample starting in 1995:1. The latter estimation is done to check the robustness of the results with respect to the inclusion of the first years after the start of the transition to a market economy. In the literature, it has been forcefully argued that typically each NMS started from a severely undervalued exchange rate in the early 1990s, which may unduly influence the estimations.[14] Removing the first two years from our sample should at least greatly diminish the severity of this problem, see Égert and Lahrèche-Révil (2003) for a similar approach. The results of the Hausmann test at the bottom of the table indicate that cross-country homogeneity for the three Baltic countries cannot be rejected.

The signs of the coefficients on productivity and openness are consistent with our hypotheses and are quite robust, both in magnitude and in significance, across specifications and sample periods. For productivity, we find an estimated elasticity of about 1 percent for the full sample. For the 1995–2003 sample, the coefficient is somewhat lower. Our results are in line with the literature, see Égert (2003). The negative (and significant) coefficient on openness confirms the hypothesis that openness through increased demand for tradables induces a real depreciation. For the demand indicators less robust results are observed. Especially in the shorter period, coefficients are relatively small in magnitude and only significant in the case of private consumption. Moreover, for private and total consumption the sign switches from positive to negative when the sample is reduced by two years. For government consumption, the coefficient is insignificant in both specifications. Note, however, that the coefficients on the other two explanatory variables are relatively insensitive to the inclusion of a specific demand indicator.

Table 2 also contains two sets of cointegration tests on the estimated long-run relations. First, we report four test statistics (p-values) based on Kao (1999). The hypothesis of no cointegration is rejected each and every time. However, like the unit root tests, standard panel-cointegration tests like in Kao (1999) suffer from possible biases due to cross-sectional dependence and finite sample size. Several tests have been proposed to tackle the problem of cross-sectional dependence (Phillips and Sul, 2003; Groen and Kleinbergen, 2003), but they all exhibit severe limitations in case of small cross-sectional and time dimensions. As an alternative, we therefore use the recent method proposed by Fachin (2006), which consists of using a block-bootstrap version of the group and mean t-statistic (Pedroni, 1999). Although Fachin's tests are more appropriate than Kao's as they take into account all model specificities, as yet to our knowledge no theoretical paper is available that shows the validity of the block-bootstrap procedures. In Table 2, we report p-values for Fachin's group (G) and mean (M) t-tests for cointegration. In each case, under the null hypothesis there is no cointegration. Since Fachin's method can only be applied in case of a single cointegrating vector, we also need to test for cointegration among the right-hand side explanatory variables and report both Kao (1999) and Fachin's (2006) statistics.[15] As opposed to Kao's tests, the Fachin tests consistently fail to reject the null hypothesis of no cointegration at the 5 percent level, apart from the case of the full sample using private consumption as demand indicator. Cointegration tests among the right-hand side (explanatory) variables follow the same pattern. Kao's tests typically reject the hypothesis of no cointegration, while the Fachin tests do not reject (apart from the full sample, private consumption case).

[14] Maeso-Fernandez et al. (2005) provide an elaborate discussion and alternatively suggest to estimate Eq. (5) for an independent group of countries not suffering from such undervaluation and then use the obtained coefficients to compute equilibrium real exchange rates for the NMS. While this approach does indeed solve the problem of potentially biased coefficients when using NMS data, it relies on the equally strong assumption that the estimated coefficients for a different group of countries apply to each NMS.

[15] We thank an anonymous referee for this suggestion.

The fragile cointegration results are puzzling and somewhat worrisome in our opinion. On the one hand, the regression results are quite stable across specifications and sample periods, on the other hand the two types of cointegration tests completely disagree. To shed more light on the issue of cointegration, we subsequently apply a 2-step approach. For this purpose, we use the DOLS residuals from the regressions in Table 2 and perform the same panel unit root tests that we used for the original series in Table 1. The results, for a size of 5 percent, are reported in the bottom part of Table 2, where the entry I(1) indicates that the unit root test does not reject the null hypothesis of a unit root, while the entry I(0) indicates rejection of the null hypothesis.[16] The evidence appears mixed. It neither (almost) uniformly indicates stationarity of the residuals nor uniformly signals non-stationarity. Some specifications end up with stationary residuals, corresponding to a cointegrating relation, while other specifications yield non-stationary residuals, consistent with a rejection of cointegration.

Overall, the unit root test results seem somewhat more supportive of the Kao cointegration test results than the Fachin results. Possibly, the difference between the Fachin (2006) cointegration results and the panel unit root results is due to the fact that the former approach is 1-step, whereas the latter is 2-step. In the method proposed by Fachin (2006), the block-bootstrap is performed on the initial series, whereas in the panel unit root test the residuals are directly block-bootstrapped. This may result in different power for the two tests. We leave the issue for future research.

In Table 3, we report corresponding results for the five Central European countries (CE5) in the sample. Broadly speaking, the results are quite similar. Again, the Hausmann tests show that cross-sectional homogeneity cannot be rejected. Deleting the first two years from the sample only marginally influences the estimation results. For productivity, the coefficients typically are positive and significant. In the full sample, the estimated productivity coefficient (elasticity) is somewhat smaller than for the Baltics, but in the period 1995–2003 the estimates are quite close. For openness, we again find a negative coefficient in most specifications. Only when government consumption is used as demand indicator, the sign on openness becomes positive, significantly so only in the period 1995–2003, however. For the CE5, the estimated coefficients are uniformly positive and in most cases significant. The pattern of cointegration tests is the same as in Table 2. Tests based on Kao (1999) uniformly indicate cointegration, while tests that account for cross-sectional dependence based on Fachin (2006) indicate absence of cointegration. The panel unit root tests on the residuals again yield mixed results. For two out of six specifications, those with government consumption over the full sample and with total consumption over the smaller sample, stationary residuals are found at the 5 percent level.

Overall, the tests are ambiguous both for the Baltics and for the CE5. No clear conclusion can be drawn on the existence of a panel cointegration relationship in our systems. Notice that Rahn (2003) also fails to find a systematic cointegration relation between the real exchange rate and its determinants in a number of cases. We conclude that the ambiguity of our cointegration results suggests strong sensitivity of the tests to minor changes in specification and sample periods. A more in-depth investigation of this issue is outside the scope of the current paper and is left for future research. Note though that we have tried to account for the uncertainty about the presence of a cointegrating relation through the estimation of different specifications and sam-

[16] Detailed results are available on request from the authors. The evidence suggests that in most cases non-stationarity would be rejected at the 10 percent level.

Table 3
Regression results: CE5

	1993–2003			1995–2003		
Gov't. consumption	0.58	–	–	0.87	–	–
	(8.87)			(13.41)		
Private consumption	–	1.93	–	–	1.53	–
		(14.67)			(11.59)	
Total consumption	–	–	2.47	–	–	2.67
			(17.78)			(21.28)
Openness	0.01	−0.59	−0.35	0.22	−0.34	−0.21
	(0.16)	(10.73)	(6.11)	(3.61)	(6.59)	(4.46)
Productivity	0.64	0.56	0.46	0.72	0.81	0.67
	(12.51)	(12.78)	(9.85)	(12.54)	(16.56)	(15.06)
R^2	0.998	0.998	0.998	0.999	0.999	0.999
Cointegration tests (p-values)						
Full system						
Kao (1999)						
DF(ρ)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
DF(T)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
DF*(ρ)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
DF*(T)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
Fachin (2006)						
$G - t$ test bootstrap	65.40	94.40	82.10	26.10	73.50	50.20
$G - t$ test fast bootst. 1	66.90	96.60	84.10	21.20	72.00	42.60
$M - t$ test fast bootst. 1	20.80	99.20	60.90	10.30	32.80	19.30
Right-hand variables						
DF(ρ)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
DF(T)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
DF*(ρ)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
DF*(T)	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
$G - t$ test bootstrap	50.20	33.40	37.50	28.90	90.00	84.60
$G - t$ test fast bootst. 1	42.60	33.10	39.40	28.70	89.20	84.40
$M - t$ test fast bootst. 1	19.30	26.10	55.70	20.00	93.40	89.00
Hausmann	12.30	11.10	14.60	16.40	25.80	20.40
Unit root residuals						
LLC	I(0)	I(1)	I(1)	I(1)	I(1)	I(0)
IPS	I(0)	I(1)	I(1)	I(1)	I(1)	I(0)
IN	I(1)	I(1)	I(1)	I(1)	I(1)	I(0)

Notes. t-values are listed between parentheses. Note that a negative sign for a variable means that the variable has a depreciating impact on the real exchange rate. All tests are performed considering a constant and using 1000 replications. Block size selection for the cointegration test is based on $0.2T$. Results appear to be robust for other block size and are available from the authors upon request.

ple periods. We interpret the fact that the variables in the different specifications are strongly related and the estimated coefficients are quite stable across specifications and sample periods as a robustness check that provides suggestive evidence of the existence of a stable cointegrating relation.

In the next subsection, we approach the issue from a different angle, using the concept of misalignments.

4.2. Misalignment

To illustrate our results in more detail, we use the estimated coefficients from Tables 2 and 3 to construct time paths of the fundamental real exchange rates for each country and graphically compare these with observed real exchange rates. To save space, only the estimations with private consumption as demand indicator are used. In Figs. 1 (Baltics) and 2 (CE5) the line labeled REAL RATE is the actual real exchange rate, while the lines labeled 'FIT 93' and 'FIT 95' represent the estimation results for the periods 1993–2003 and 1995–2003, respectively.

A number of points stand out. First, especially in the first two years of the sample sizable gaps between the actual real exchange rate and the estimated rates emerge, with generally strong undervaluation of the actual real rate. Poland and Hungary show some overvaluation at the start. Second, from 1995 onward, the estimated (fundamental) real exchange rate paths (FIT 93 AND FIT 95 respectively) move closely together. Third, the estimated under- or overvaluation at the end of the sample period is not excessively large. Fourth, sizable swings have occurred in the past, but the graphs suggest that differences between the actual and the estimated fundamental

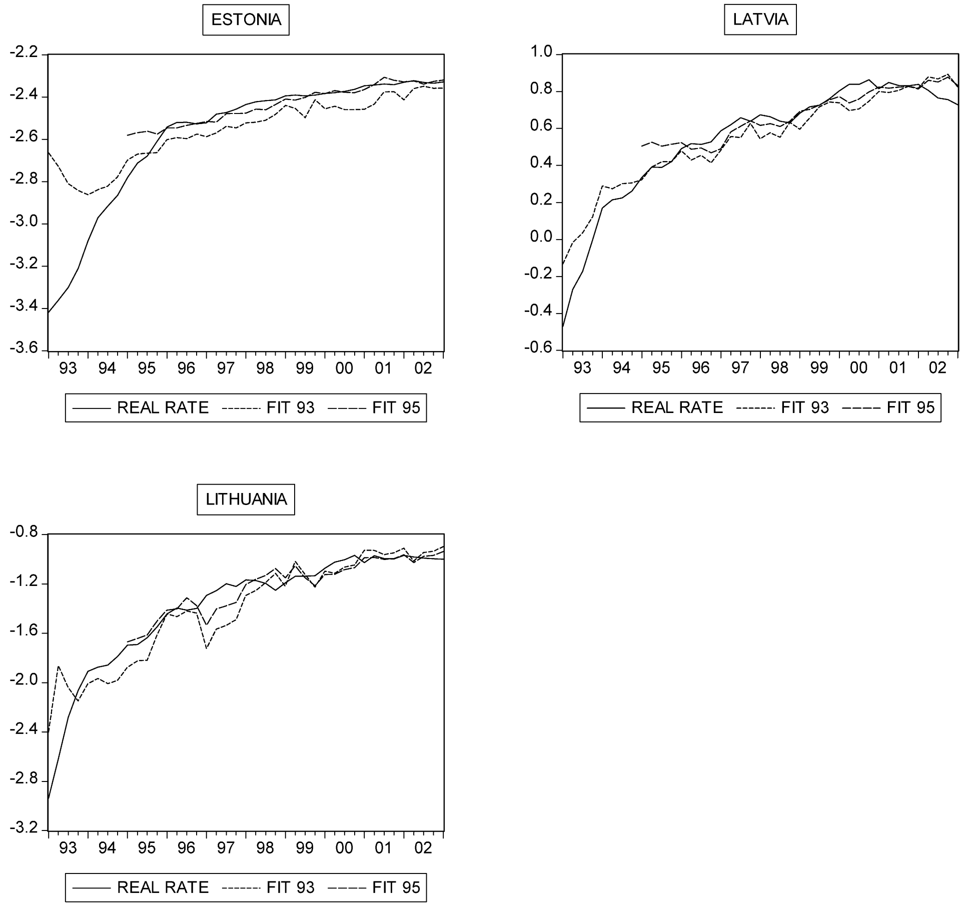

Fig. 1. Actual and estimated real exchange rates: Baltics.

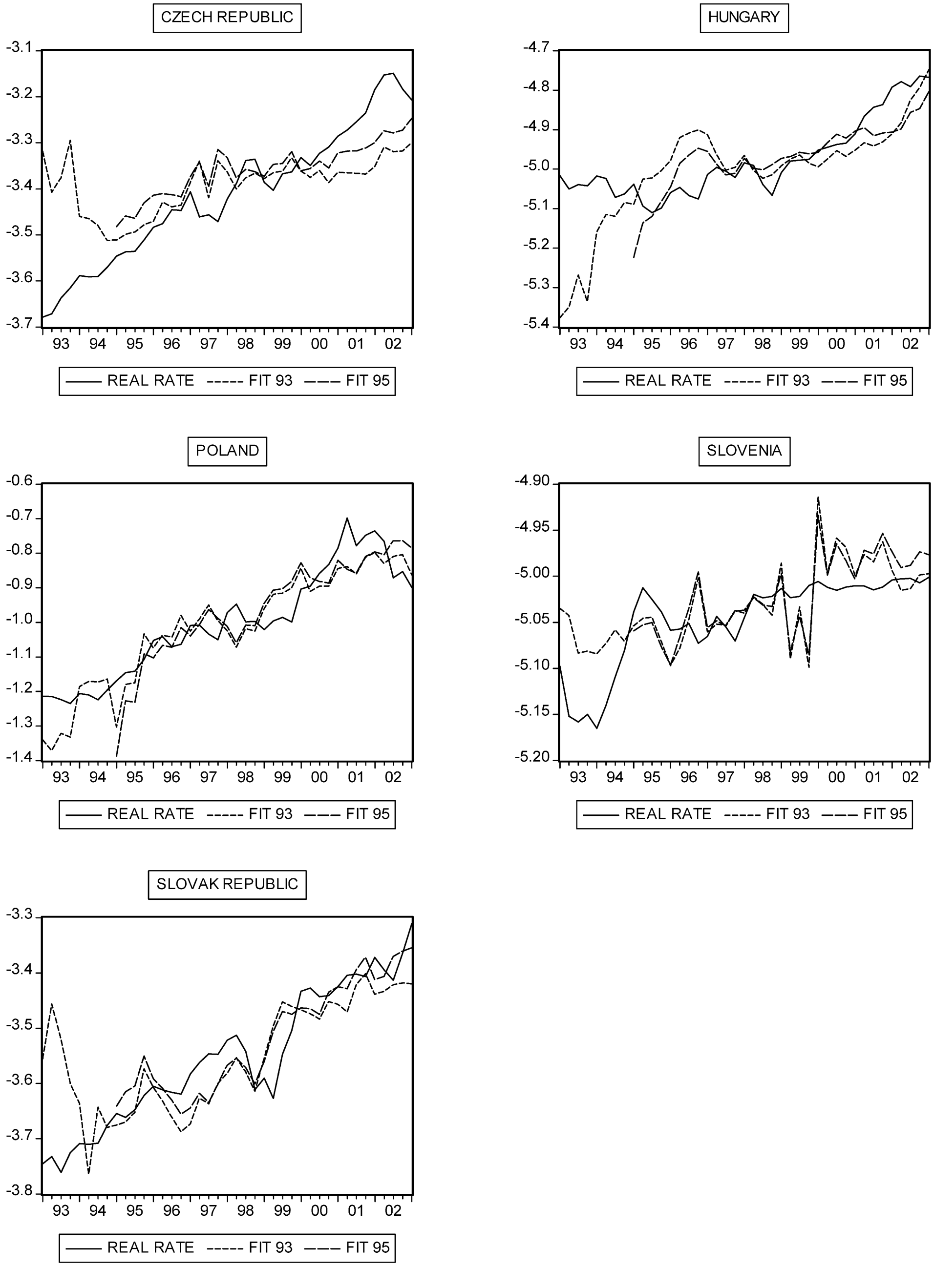

Fig. 2. Actual and estimated real exchange rates: CE5.

exchange rate tend to disappear over time. Overall, we conclude that our results do not display excessive sensitivity to model specification and choice of estimation period.

So far we have looked at the estimated fundamental exchange rate level directly and compared it to the actual real exchange rate. Now we turn to the difference between the actual and fundamental real exchange rate, which may be cautiously interpreted as a measure of under- or overvaluation, or misalignment as it is loosely called in the literature. Caution is required for several reasons. First, our estimated equilibrium exchange rate is not based on PPP and underlying actual prices, but on a BEER concept using price indexes. Second, it is unclear whether the individual countries' exchange rates were on average in equilibrium over the estimation period. Especially for the longer sample that includes the first years of strong undervaluation, this seems to be a heroic assumption. Third, the concept of misalignment implicitly assumes that actual exchange rates return to their equilibrium values over time. However, our cointegration results throw some doubt on this property. Finally, even if the above held up, point estimates of misalignments are surrounded by a confidence interval and, thus, only provide imprecise measures of the actual misalignment. Nevertheless, measuring misalignments is important because it provides information on possible strains on the exchange rate regime and on the direction of future changes in the exchange rate.

In Table 4, we consider the evidence on overvaluation or undervaluation for each country using the most recent period in our sample, the first quarter of 2003. For each country we present the estimated misalignment in 2003:1 as a percentage of the underlying fundamental real exchange rate at that time. Using all models, we have six estimates per country. A negative number implies an undervaluation, while a positive number indicates an overvaluation. In our opinion, the results derived from the estimation of the shorter sample 1995–2003 are more reliable than those from the longer sample due to the fact that the longer sample includes the strong undervaluation years 1993–1994. These may unduly result in point estimates of undervaluation later in the sample. We, therefore, focus on the short sample results.

According to Table 4, the different model specifications tend to give the same answer to the question of whether a specific currency was overvalued or undervalued in 2003:1. The exceptions are the Czech Republic and Hungary. In the Baltics, all countries were undervalued in early 2003, although especially Estonia was quite close to equilibrium. The misalignments for Latvia and

Table 4
Estimated misalignments in 2003:1

	EST	LAT	LIT	CZE	HUN	POL	SLV	SVK
Misalignment (%, 2003:1)								
Full sample								
Gov't. consumption	2.87	−7.14	−4.02	4.56	10.84	−8.01	−5.70	14.94
Private consumption	3.00	−9.30	−9.74	9.63	−1.96	−3.56	−0.38	11.49
Total consumption	6.96	−6.32	−4.24	4.76	−4.36	−3.90	−0.43	15.38
Short sample								
Gov't. consumption	−0.97	−9.51	−5.92	−1.87	8.86	−13.07	−4.70	10.78
Private consumption	−0.85	−9.82	−6.00	4.08	3.68	−10.71	−2.40	4.45
Total consumption	−0.64	−9.53	−6.12	−1.05	−5.03	−9.84	−1.46	7.78
Standard deviation of the misalignment (%, 1995:1–2003:1)								
Full sample								
Gov't. consumption	5.06	8.55	12.63	7.74	5.95	6.74	4.47	8.28
Private consumption	4.33	7.13	15.38	8.13	7.16	6.84	3.65	5.06
Total consumption	4.98	7.10	12.23	7.09	5.91	7.13	3.37	5.76
Short sample								
Gov't. consumption	4.75	6.50	8.93	7.11	5.72	7.89	4.42	8.70
Private consumption	4.83	6.70	8.88	7.08	6.59	8.22	3.51	4.50
Total consumption	4.77	6.59	8.78	6.16	5.30	9.36	3.48	5.17

Lithuania were between 6 and 10 percentage points. For the Central European countries, we find considerable overvaluation for Slovakia and undervaluation for Poland. Note that the Polish zloty strongly depreciated in nominal terms in the two previous years. Prior to that, the zloty was typically overvalued. For Slovenia, some small undervaluation is reported, while the results for the Czech Republic and Hungary are mixed. Especially for the Czech Republic the absolute size of the misalignments is limited.

Our results are largely comparable to similar results of research on bilateral real exchange rates. Rahn (2003) calculates misalignments for the first quarter of 2002, based on estimations over the period 1990–2002 for Poland and Hungary, 1993–2002 for the Czech Republic and Estonia and 1992–2002 for Slovenia. His conclusions are close to our full sample results from 1993–2003, even though the timing of the misalignment computations differs by a year. Like us, Rahn reports an overvaluation for Poland, the Czech Republic and Estonia, while the results for Hungary are mixed. Contrary to our results, Rahn finds an overvaluation for Slovenia. Note that for some countries, we find different results compared to Rahn when using the sample period 1995–2003.

Kim and Korhonen (2005) document an overvaluation of the real effective exchange rate for the Czech Republic, Hungary, Poland and Slovakia in 2002, based on estimates over the period 1975–1999. Bulĩr and Šmídková (2005) calculate an overvaluation of the Czech koruna, the Polish zloty and the Hungarian forint in 2003, while the Slovenian tolar is close to equilibrium. Their calculations concern bilateral exchange rates and are based on data that cover the period 1995–2003. Overall, the literature shows that both the sample period over which the equations are estimated as well as the point in time for which misalignments are computed have a significant effect on point estimates of misalignments. Caution is therefore necessary when comparing the results across studies.

The bottom panel of Table 4 reports the standard deviations of the misalignments over the sample period 1995–2003. Note that the size of the standard deviation of the misalignment, focusing on the shorter sample estimates, differs from around 3.5 percent for Slovenia to almost 9 percent for Lithuania. In most countries, the reported misalignment in 2003 is less than one standard deviation away from zero. Only for Latvia, the misalignment is about 1.5 times the standard deviation. Given the uncertainty about the estimated level of the fundamental exchange rate, the estimated misalignments are small enough to be insignificantly different from zero.

Despite the fact that the Fachin test results do not formally support cointegration, we nevertheless investigate whether and how fast misalignments are corrected. To this purpose, we estimate the extent to which the actual real exchange rate moves back in the direction of the fundamental exchange rate in the next quarter. In Table 5, we present the panel estimation results of the corresponding error-correction model. Theoretically, the model could be extended to include other determinants of short-run real exchange rate dynamics such as real interest rate differentials. Here, we confine the analysis to a simple regression of the change in the real exchange rate on the lagged value of the misalignment.

The coefficients on the lagged misalignment variable are consistently negative, but often only marginally significant. The size of the mean reversion effects is generally small. For the Baltics, a stronger result is found over the full sample (panel A). However, that is probably due to the strong undervaluation in the first years that is corrected afterwards. Over the shorter sample (panel B), the results for the Baltics and the CE5 are surprisingly close. They suggest that in a given quarter around 8 percent of the prevailing misalignment is eliminated. The observed slow speed of adjustment is often found in empirical real exchange rate analyses.

Table 5
Panel error-correction results

	Baltics, 1993–2003			CE5, 1993–2003		
	Gov't. cons.	Priv. cons.	Total cons.	Gov't. cons.	Priv. cons.	Total cons.
Panel A						
Intercept	0.034*	0.034*	0.034*	0.007*	0.007*	0.007*
	(7.32)	(7.38)	(7.91)	(3.78)	(3.76)	(3.80)
Residual_{t-1}	−0.173*	−0.190*	−0.206*	−0.045*	−0.055**	−0.040**
	(7.02)	(7.11)	(8.83)	(2.60)	(2.59)	(2.20)
N	123	123	123	205	205	205
Adj. R^2	0.283	0.289	0.387	0.027	0.027	0.018
F-Stat.	49.225*	50.483*	78.006*	6.767*	6.722*	4.838**
Half-life (quarters)	2.9	2.6	2.4	11.2	9.1	12.4
	Baltics, 1995–2003			CE5, 1995–2003		
	Gov't. cons.	Priv. cons.	Total cons.	Gov't. cons.	Priv. cons.	Total cons.
Panel B						
Intercept	0.015*	0.015*	0.015*	0.007*	0.007*	0.007*
	(4.78)	(4.75)	(4.79)	(3.25)	(3.30)	(3.26)
Residual_{t-1}	−0.084	−0.072	−0.090**	−0.080**	−0.127*	−0.089**
	(1.80)	(1.55)	(1.91)	(2.42)	(3.43)	(2.34)
N	99	99	99	165	165	165
Adj. R^2	0.022	0.014	0.026	0.029	0.062	0.027
F-Stat.	3.232	2.389	3.639**	5.876**	11.768*	5.470**
Half-life (quarters)	5.9	7.0	5.6	6.3	3.9	5.6

Note. The t-statistics are in parentheses.
* Significance at the 1% level. ** Idem, 5%.

5. Conclusion

In this paper, we have estimated fundamental bilateral real exchange rates for a set of eight new EU member states (NMS) using a panel-cointegration approach over the period 1993–2003. Given the differences between the Baltic States and the Central European countries, we have estimated the model for the three Baltic States separately and for the five Central European countries separately.

Overall, the results consistently demonstrate that an increase in productivity in the tradables sector of an NMS relative to the euro area does indeed lead to an appreciation of the NMS real exchange rate. This offers indirect support for the BS effect. Also, we generally find significantly negative effects from openness to the real exchange rate. For demand indicators, less convincing and consistent evidence is found. Cointegration tests yield mixed results and throw doubt on the convergence of actual exchange rates to fundamental levels. More and longer data series may help to shed more light on the issue.

Concerning the estimated misalignments, the results are robust and generally in line with the literature. In the early 1990s, considerable misalignments are observed. Casual evidence suggests that these have disappeared to a large extent. While the actual and estimated real exchange rate deviate from one another through time, the gaps remain relatively small, especially when compared to the uncertainty surrounding the estimates.

Our research shows steadily appreciating real exchange rates for all countries under consideration and documents a significant positive link between productivity levels and the corresponding real exchange rate levels. Future rises in real exchange rates and relative productivity levels cannot be excluded on the basis of either our own analysis or the literature as a whole.

Nevertheless, real exchange rate appreciations solely due to differences in productivity growth have been limited over the past years. Between 1998 and 2003, average appreciation due to productivity effects varied between zero and 0.66 percent. Given our estimated productivity coefficient of about 0.75 both for the Baltics and the CEE countries, this would yield a maximum real appreciation of about 0.5 percent per year. Assuming relative productivity rises in any of the NMS do not exceed 0.66 percent per year in the future, the required real exchange rate appreciation, in a fixed exchange rate regime, can be brought about by additional CPI inflation of at most 0.5 percent per year. From this we conclude that real exchange rate appreciations purely brought about by productivity effects are probably too limited to exert extra tensions between the Maastricht criteria of nominal exchange rate stability and inflation stability. Note that according to the Maastricht inflation criterion an individual country's inflation should not be more than 1.5 percent over average inflation in the three lowest inflation countries. Typically, average inflation in the three lowest inflation countries is between 0.5 and 1 percent below average euro area inflation.

However, between 1998 and 2003, the overall average real appreciation across NMS countries varied between 0.12 and 1.09 percent per year. Future real exchange rate appreciations in the same order of magnitude and limited to at most 1.1 percent per year for an individual NMS can under nominally fixed exchange rates be brought about by CPI inflation in that NMS of at most 1.1 percent in excess of average euro area inflation. An excess inflation around one percent, however, could violate the Maastricht inflation stability criterion.

Consequently, inflation pressure and real exchange rate appreciation in the new member states probably remain a fact of life in the near future. In some instances, real exchange rate appreciations may be strong enough to cause tensions between the Maastricht criteria of nominal exchange rate stability and low inflation. In the context of entry into ERM-II and EMU for most of these countries over time, it stresses the importance of allowing countries to flexibly adjust to these developments, either through nominal appreciation or temporarily higher domestic inflation without jeopardizing their fast entry into the euro area.

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