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The Maastricht criteria in the context of the Central European transition economies:

A potential conflict between real and nominal convergence?

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Abstract

This paper examines the extent to which productivity developments in the Central European transition economies have contributed to the sustained rise in their real effective exchange rates vis-à-vis the Eurozone over the past 11 years. The countries examined (the Czech Republic, Hungary, Poland and Slovakia) are all members of the European Union and are set to join the European Monetary Union in the future. This requires nominal convergence in accordance with Maastricht criteria. In the context of monetary union, movements in the real exchange rate reflect inflation differentials expressed in a common currency. Such differentials with the Eurozone will become an important issue once EMU membership is discussed.

Existing studies have so far been unable to focus their analysis explicitly on Central Europe without applying some markedly restrictive assumptions. This study succeeds in relaxing some of these assumptions, and illustrates that these have led to biased estimates. Estimates suggest that the impact of productivity differentials on inflation differentials is indeed significant, albeit to a lesser extent than previously suggested. Finally, this study questions the suitability of a more flexible convergence policy within the framework of the Maastricht Treaty.

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I. Introduction

On May 1, 2004 ten new members joined the European Union. Union membership should lead to membership in the European Monetary Union (EMU), and therefore requires nominal convergence in accordance with Maastricht criteria. In particular, exchange rate convergence and price stability are required.[1] Convergence criteria are the same they were for existing members. The question arises whether these can be justifiably imposed on the newly acceded countries. In particular this may result in a conflict between nominal and real convergence.

The EU has in its history expanded several times through waves of enlargement. The latest expansion is arguably different in many respects, and warrants putting the suitability of existing convergence criteria into question. Most newly acceded countries are small, open economies with large ties to other EU states. As such they should reap substantial benefits from euro adoption. What sets them apart from other EU states, however, is their transition status. Their transition from centrally-planned to market-based systems since 1990 has been characterised by rapid productivity growth. Indeed, their relative backwardness has made "catch up" with the EU a key economic objective.

The Balassa-Samuelson hypothesis (Balassa, 1964; Samuelson, 1964) implies that high rates of productivity growth may lead to increased inflation and continuous real exchange rate appreciation, both of which have been evident features in the accession economies. There may therefore arise a conflict between exchange rate and price stability within ERM II for accession countries.[2] Inflation differentials associated with the Balassa-Samuelson effect reflect an equilibrium phenomenon and do not require a policy response. Thus, rather than calling for exchange rate flexibility, the presence of a strong Balassa-Samuelson effect may highlight the drawbacks of the Maastricht criterion on inflation for accession countries.

The existing body of literature, which has tested the Balassa-Samuelson hypothesis (B-S) for different regions throughout the world, is vast and expanding. The focus on Central Europe, however, has been limited – mainly due to data restrictions. In particular, a series of papers by Egert (2002a, 2002b) and Egert et al. (2003) analyses B-S in the context of Central Europe under markedly restrictive assumptions. The purpose of this paper is to show that their resulting

[1] The Maastricht criterion for inflation convergence states that a country joining EMU should have an inflation rate no more than 1.5% above the average of the three lowest rates of inflation in the Eurozone.
[2] During ERM II, nominal exchange rates need to be kept within a fixed band, limiting the central bank's ability to combat rising domestic inflation through nominal currency appreciation.

estimates are in fact biased and inconsistent. This paper will present a more thorough study of B-S in the context of Central Europe, from which conclusions about the suitability of existing convergence criteria are drawn.

II. Related Research

Theoretical Literature

The Balassa-Samuelson hypothesis (B-S) is best understood, considering a two-country world, where a single input, labour, is used to produce both traded and non-traded goods.[3] It assumes purchasing power parity to hold for traded but not non-traded goods, the latter being sheltered from foreign competition. Finally, domestic labour mobility is assumed to cause cross-sector wage equalisation.

The B-S effect argues that a country experiencing high productivity growth relative to another country in the tradable versus the non-tradables sector will experience an equilibrium real exchange rate appreciation. The central tenet is that if productivity growth in tradables outpaces that of non-tradables, the non-tradable sector will experience wage growth in excess of its productivity growth due to cross-sector wage equalisation. The result is higher non-tradables prices, an increase in the aggregate price level and an equilibrium appreciation of the real exchange rate.[4]

Bergstrand (1991) added a demand side explanation for the relative changes in non-tradables inflation. Assuming non-homothetic tastes,[5] he showed that demand may affect the relative price of non-tradables through altering the composition of output. Hereby countries with higher per capita GDP have higher price levels because non-traded services are luxuries in consumption, whereas traded commodities are necessities. GDP growth thus causes consumers to spend a relatively higher proportion of their additional income on non-tradables causing an increase in the relative price level of non-tradables.

[3] The Balassa-Samuelson effect does not rely on the assumption that labour is the only factor of production. Identical results are obtained in a model including capital.

[4] Rising productivity induces increases in income and wealth, hence rising consumption. If demand growth were to be biased towards traded goods, then the supply side effect could be offset, partly or even undone. If, as considered by Bergstrand (1991), demand is biased towards services, the bulk of non-traded goods, the demand side effect reinforces the supply side effect.

[5] Homothetic tastes imply that individuals do not change their relative consumption between goods as income levels rise. Non-homothetic tastes thus imply that the composition of the consumption basket shifts with rising income levels. For example, as countries become richer the proportion of income spent on food tends to drop, whereas luxury goods account for an increasingly greater share of the consumption basket.

Empirical Literature

An extensive body of empirical analysis has tried to validate B-S in the context of industrialised countries with varying success. In general, studies focusing on the US and Japan provide evidence in favour of B-S, whereas in the case of OECD countries results are more varied (see Appendix 1 for details).

In the context of the European transition economies, only short time series are available. Studies thus usually rely on pooled observations in panel studies. Most focus on the link between sectoral productivity growth gaps and relative non-tradables inflation. Following the extension of Bergstrand (1991) the general equation tested looks as follows:

$$\log\left(P_{N,i,t}\right) = \beta_{0,i} + \beta_1\theta_{i,t} + \beta_2 g_{i,t} + \beta_3 \log y_{i,t} + \beta_4 \Delta\pi_{i,t} + \beta_5 Misc_{i,t} \qquad (1.1)$$

This empirical approach goes back to De Gregorio et al. (1994) – please refer to the footnote for variable definitions.[6] The coefficient on $\theta_{i,t}$ measures the impact of productivity growth as suggested by B-S with an expected positive sign. Variables g and y proxy demand shifts. With a positively sloped supply curve, both are expected to enter positively. Relaxing the assumption of instantaneous price adjustment in the non-tradables sector, De Gregorio et al. (1994) also include the first difference of inflation to capture the possibility of transitory dynamics. *Misc* are miscellaneous other explanatory variables, which in part intend to capture movements off an equilibrium path that is mainly determined by productivity developments. Such variables include different exchange rate regimes (Halpern-Wyplosz, 2001); openness, terms of trade and money-to-GDP ratio (Broeck-Slok, 2001); oil price shocks (De Gregorio et al., 2004); or structural variables (Coricelli-Jazbec 2004).[7]

Amongst the most thorough studies, Halpern and Wyplosz (2001) investigate a panel of eleven European transition economies over the period 1991-9. They characterise the real exchange rate path during the early transition period as one with an initial steep depreciation[8] followed by a continued appreciation[9]. Having initially overshot their equilibrium path, real exchange rates

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[6] $P_{N,i,t}$ corresponds to the relative price of non-tradable goods in country i at time t. $\theta_{i,t}$ is the difference of total factor productivity across sectors corrected by the labour shares (θ=(δ_N/ δ_T)log θ_T – log θ_N), g is government expenditure over GDP (in real terms), y is per capita income, and $\Delta\pi$ is the first difference of the rate of inflation.

[7] A theoretical justification for the inclusion of these variables is provided in Frenkel and Mussa (1985).

[8] The initial depreciation is attributed to a sudden excess demand for foreign assets, flight from the domestic currencies because of a burst of inflation, and/or loose exchange rate policies.

[9] The subsequent appreciation is said to reflect a variety of factors including productivity gains associated with the introduction of market forces (i.e. over-employed industrial sector releases labour into the underdeveloped service sector) and technological catch-up; price liberalisation leading to increased inflation and real appreciation; changes in the composition of public spending, redirecting from the industrial to the non-traded sector; and high potential returns on capital leading to large capital accumulation (Begg et al. (2003) estimate that foreign direct investment

appreciated due to a combination of a return to the equilibrium path and as a result of the transformation process. The rate of equilibrium appreciation is higher the more complete the market system is and the faster capital is accumulated.[10]

Testing for the presence of the individual mechanisms underlying B-S, the link from labour productivity to wages is found to be highly significant. Allowing for labour market characteristics to affect real wages, they find that unemployment reduces wages, but only in industry. This again supports B-S which sees wages in the service sector driven not by labour market conditions in this sector but by a tendency towards cross-sector wage equalisation. Following the formulation in (1.1), β_1 enters with a positive significant coefficient, confirming the B-S effect. A 10% rise in relative industry productivity is found to increase the relative price of non-tradables by 2.4%/4.4% in the short run/long run (similar results are reported by Begg et al, 2003). The coefficients on the demand side factors support the possibility of a bias towards non-tradables. When controlling for different exchange rate regimes they plausibly conclude that floating regimes strengthen the B-S effect.[11]

Coricelli and Jazbec (2004), using an unbalanced panel for 19 transition countries over the period 1990-1998, find that initial adverse conditions and structural reforms affect the real exchange rate only in the first five years of transition. Subsequently, the B-S effect seems to dominate in Central and Eastern Europe, whereas demand factors overweigh in the Baltic states. Following the approach by De Gregorio et al. (1994), equation (1.1) is augmented to include a structural misalignment variable, $lab_{i,t}$, proxied by the ratio between tradable and non-tradable sector employment.[12] The labour ratio enters the regression with an expected positive sign, suggesting that delays in structural reforms tend to act as a restraining force on the real exchange rate. This force seems to diminish after five years. The coefficients on time dummies, accounting for unobservable factors during the transition process, also diminish after the fifth year, reinforcing the claim that structural reforms did not affect real exchange rate determination at later stages of transition.

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increases industrial labour productivity six times more than in the service sector). Note that the first and last effect lead to an increase in the real exchange rate through the B-S effect.

[10] These stylised facts are supported by papers from Krajnyák and Zettelmeyer (1998), Begg et al. (1999) and Coricelli and Jazbec (2001, 2004)

[11] If the exchange rate is free to absorb some of the equilibrium real appreciation in the form of nominal appreciation rather than forcing adjustment through absolute price changes, the effect is bound to appear faster, which makes a difference given the short time series.

[12] During the pre-transition period labour was inefficiently allocated between the two sectors, with a heavy emphasis on the former. The variable thus intends to capture the rigidity of the labour market to structural changes in the economy.

Finally another noteworthy study by De Broeck and Sløk (2001), use panel data comprising three groups of countries[13] over the period 1993-8. Rather than relative prices they choose the trade-weighted real effective exchange rate (REER) as their dependent variable. The B-S effect seems to be a robust explanation for the development of the REER in Eastern Europe, but fairs less favourably (yet significantly) in other transition economies, which are at a comparatively early transitional stage. Notably, in the case of these "other" transition countries, the significant relationship reflects a fall in tradables productivity which is associated with a depreciation of the real effective exchange rate.[14] *Misc* variables, which have significant coefficients, are openness and government balance. The coefficients on other variables such as the terms of trade or money-to-GDP ratio appear insignificant.

Since no reliable measure of capital is available for most transition countries, it is common practice for total factor productivity to be proxied by average labour productivity. How to differentiate between traded and non-traded sectors, however, differs across studies and to some extent reflects the availability of sectoral data. Canzoneri et al. (1999), using the OECD sectoral data base, capture traded goods through the manufacturing and agricultural sectors, and non-traded goods through a whole range of service sectors.[15] De Broeck and Sløk (2001), keep the agricultural sector separate as it typically comprises a mixture of tradable and non-tradable activities. Similarly, Fischer (2002) treats agriculture as a residual on the grounds that its prices are often publicly regulated. Many other studies fail to include precise specifications.

Most sectoral data is only available on an annual basis, resulting in panels with a relatively small time-series dimension, and moreover, preventing the focus on narrower ranges of countries. A series of papers by Égert (2001, 2002) and Égert et al. (2003) has therefore attempted a more disaggregated approach, using quarterly or even monthly data. In particular, given the apparent absence of monthly or quarterly data for the non-tradable sector, non-tradable productivity growth is set to zero, and the traded sector is proxied by industrial production alone. Égert (2001) argues that progress in non-tradable productivity across countries should be about the same, which in turn means no-biased estimates of the productivity differential.

[13] Eastern Europe (9 countries), other transition countries (16 countries), OECD (17 countries)

[14] This is in line with B-S, which predicts that, irrespective of their direction, changes in productivity in the tradable sector will affect the real effective exchange rate.

[15] In particular, the non-traded sector is made up of the "wholesale and retail trade, restaurants and hotels" sector, the "transport, storage and commtication" sector, the "finance, insurance, real estate and business setices" sector, the "community social and personal services" sector, and the "non-market services" sector.

Limited data availability has also led most studies (including Égert's) to include the early 1990s into their regressions. In failing to control for the structural changes which affected real exchange rates during this period, estimated coefficients are likely to provide biased estimates. Also, in constructing panels of countries for different regions, it is clear from the literature that countries at different stages in the transition process exhibit different characteristics.[16] Large panels, necessitating restrictive assumptions concerning the homogeneity of effects across countries, may thus provide unsatisfactory estimates of true "population effects". The increased availability of data today should warrant studies with a narrower focus.

Finally, the choice of the objective variable in existing studies makes their results inadequate for drawing concise conclusions about the presence of B-S and more importantly, in the context of EU transition economies, about the feasibility of adhering to Maastricht convergence criteria. Most studies choose relative "non-tradable inflation" as their dependent variable and treat this as a proxy for the real exchange rate. However, it is the inflation differential *between* trading regions measured in a common currency, which is to be explained. Real exchange rate appreciations that reflect large productivity gains in one country *relative to another* in the tradable versus the non-tradable sector are an equilibrium phenomenon and do not require a policy response. In order to determine the magnitude of this equilibrium appreciation, observations for the EU transition economies thus need to be scaled relative to those of its trading partners in the Eurozone.

In choosing the REER, De Broeck and Sløk (2001) and Fischer (2002), do indeed appreciate the importance of this approach. However, given that the REER includes exchange rate and inflation differentials with countries outside the euro-area, their approach will not give us precise estimates with which to draw conclusions about the potential conflict between nominal and real convergence for EU accession countries in the run-up to EMU membership. Égert et al. (2003) avoid this problem by using Germany as their proxy for the Eurozone, making the real deutschmark exchange rate their objective variable.

This study proposes an improved proxy for the Eurozone, which includes France, Germany, Italy and the Netherlands. These countries account for about 80 percent of euro area GDP and just under 80 percent of the trade share with Central Europe. Moreover, having been able to

[16] In general, transition economies are likely to display trend appreciation through B-S at least as long as they catch up with more advanced economies. Such supply-side factors will tend to diminish the more these countries have caught up. This is arguably the case when considering the results by Coricelli-Jazbec (2004), that B-S seems to be a dominant force in driving the real exchange appreciation in Central and Eastern Europe, whereas in the more advanced Baltic states demand factors have become more pronounced.

obtain detailed quarterly sectoral output and employment data for both the traded and non-traded sectors for the Czech Republic, Hungary, Poland and Slovakia, as well as for France, Germany, Italy and the Netherlands, the estimates of this study provide more accurate estimates of both traded and non-traded labour productivity developments in these countries. Finally in analysing the data for 1995Q1-2005Q3, we omit the early 1990s from our data set. We thereby avoid the difficulties involved in controlling the multitude of effects present at the time.

III. Analytical Framework

In the absence of price level data, analysis needs to be focussed on the changes in productivity and the real exchange rate. Thus all variables need to be expressed in terms of logs. To simplify notation in presenting the analytical framework, variables here are expressed in growth rates, identified by "^", and in a common currency.

B-S is best understood, considering a two-country world, where a single input, labour, is used to produce both traded and non-traded goods.[17]

$$\begin{aligned} Y^{NT} &= A^{NT} L^{NT} \quad & Y^{NT*} &= A^{NT*} L^{NT*} \\ Y^{T} &= A^{T} L^{T} \quad & Y^{T*} &= A^{T*} L^{T*} \end{aligned} \tag{3.1}$$

where superscripts *T* and *NT* indicate the traded and non-traded sectors; *Y*, *A* and *L* refer to output, average labour productivity and labour, respectively, and * denotes the Eurozone. Consumers are assumed to spend a share α (1-α) on tradable (non-tradable) goods, which using a Cobb-Douglas utility function gives rise to the following price indices:

$$\begin{aligned} P &= \left(P^{T}\right)^{\alpha}\left(P^{NT}\right)^{1-\alpha} \Rightarrow \hat{P} = \alpha \hat{P}^{T} + (1-\alpha)\hat{P}^{NT} \\ P^{*} &= \left(P^{T*}\right)^{\alpha*}\left(P^{NT*}\right)^{1-\alpha*} \Rightarrow \hat{P}^{*} = \alpha \hat{P}^{T*} + \left(1-\alpha^{*}\right)\hat{P}^{NT*} \end{aligned} \tag{3.2}$$

Purchasing power parity is assumed to hold for traded but not non-traded goods, the latter being sheltered from foreign competition:

$$\begin{aligned} \hat{P}^{T} &= \hat{P}^{T*} \\ \hat{P}^{NT} &\neq \hat{P}^{NT*} \end{aligned} \tag{3.3}$$

Competition for labour leads workers to be paid the marginal product of what they produce:

$$\begin{aligned} \hat{P}^{T} &= \hat{W}^{T} - \hat{A}^{T} \quad & \hat{P}^{NT} &= \hat{W}^{NT} - \hat{A}^{NT} \\ \hat{P}^{T*} &= \hat{W}^{T*} - \hat{A}^{T*} \quad & \hat{P}^{NT*} &= \hat{W}^{NT*} - \hat{A}^{NT*} \end{aligned} \tag{3.4}$$

[17] The Balassa-Samuelson effect does not rely on the assumption that labour is the only factor of production. Identical results are obtained in a model including capital.

where W refers to the wage rate. Domestic labour mobility encourages consistency in wages across sectors:

$$\hat{W}^{T} = \hat{W}^{NT}$$
$$\hat{W}^{T*} = \hat{W}^{NT*} \tag{3.5}$$

Finally, the real exchange rate represents the ratio of aggregate price indices, expressed in common currency:

$$R\hat{E}R = \hat{P} - \hat{P}^{*} = \left[(1-\alpha)\hat{P}^{T} + \alpha\hat{P}^{NT}\right] - \left[\left(1-\alpha^{*}\right)\hat{P}^{T*} + \alpha^{*}\hat{P}^{NT*}\right], \tag{3.6}$$

Substituting in from equation (3.4) we obtain:

$$R\hat{E}R = \left[(1-\alpha)\left(\hat{W}^{T} - \hat{A}^{T}\right) + \alpha\left(\hat{W}^{NT} - \hat{A}^{NT}\right)\right] - \left[\left(1-\alpha^{*}\right)\left(\hat{W}^{T*} - \hat{A}^{T*}\right) + \alpha^{*}\left(\hat{W}^{NT*} - \hat{A}^{NT*}\right)\right] \tag{3.7}$$

Substituting in from equation (3.5) we obtain:

$$RER = \left[(1-\alpha)\left(\hat{W}^{T} - \hat{A}^{T}\right) + \alpha\left(\hat{W}^{T} - \hat{A}^{NT}\right)\right] - \left[\left(1-\alpha^{*}\right)\left(\hat{W}^{T*} - \hat{A}^{T*}\right) + \alpha^{*}\left(\hat{W}^{T*} - \hat{A}^{NT*}\right)\right]$$

$$\Rightarrow R\hat{E}R = \left[\hat{W}^{T} - \hat{A}^{T} + \alpha\hat{A}^{T} - \alpha\hat{A}^{NT}\right] - \left[\hat{W}^{T*} - \hat{A}^{T*} + \alpha^{*}\hat{A}^{T*} - \alpha^{*}\hat{A}^{NT*}\right] \tag{3.8}$$

Substituting back in from equation (3.4) we are left with:

$$R\hat{E}R = \left[\hat{P}^{T} + \alpha\left(\hat{A}^{T} - \hat{A}^{NT}\right)\right] - \left[\hat{P}^{T*} + \alpha^{*}\left(\hat{A}^{T*} - \hat{A}^{NT*}\right)\right], \tag{3.9}$$

which given equation (3.3) simplifies down to:

$$R\hat{E}R = \alpha\left(\hat{A}^{T} - \hat{A}^{NT}\right) - \alpha^{*}\left(\hat{A}^{T*} - \hat{A}^{NT*}\right) \tag{3.10}$$

Hence by combining principles (3.1-3.5) with the definition of the real exchange rate, we construct the Balassa-Samuelson measure as the percentage change in the real exchange rate caused by the difference in sectoral productivity growth gaps ($\hat{A}^{T} - \hat{A}^{NT}$) and ($\hat{A}^{T*} - \hat{A}^{NT*}$). Note from (3.10), that differences in consumer preferences (as reflected by α) may induce deviations from relative PPP in the absence of cross-country differences in sectoral productivity growth gaps. For a theoretical justification of using relative non-tradable prices as a proxy for the real exchange rate, please refer to Appendix 2.

IV. Data Description and Preliminary Analysis

Data Description

Our objective variable is the CPI-based, trade-weighted real effective exchange rate vis-à-vis the euro-area (*REER*, see Appendix 3). Our proxy for the Eurozone is comprised of France, Germany, Italy and the Netherlands (marked by subscripts *j*, in equation (4.1)). Indices were constructed using EU harmonised CPI series and euro exchange rates (synthetic indices prior to 1999). Country weights, w_j, were assigned according to the country's respective share in the overall trade with Central European country, *i*:

$$REER_{it} = \sum_{j=1}^{4} w_j \frac{P_{it} e_{it}}{P_{jt}} \tag{4.1}$$

The nominal exchange rate, *e*, is defined as the unit price of Central European currency expressed in euros. Similarly, the *REER* is expressed as the price of Central European goods in terms of Eurozone goods. Therefore an increase in the $REER_i$ corresponds to a real currency appreciation of Central European country, *i*. In the absence of price-level data, all *REER* indices were constructed using quarter-on-quarter growth rates. The resulting indices enter future equations in log terms.

Our main explanatory variable measures the sectoral productivity growth gap differential between the respective CE country and the Eurozone (*BS*, see Appendix 4). Due to a lack of data on capital stocks for CE economies, total factor productivity is proxied by average labour productivity (*A*):

$$\log(BS)_{it} = \alpha_{it} \log\left(\frac{A^T}{A^{NT}}\right)_{it} - \alpha_t^* \log\left(\frac{A^{T*}}{A^{NT*}}\right)_t \tag{4.2}$$

The classification of sectors follows De Gregorio et al (1994), whereby a sector is considered tradable if more than 10% of total production is exported (see Appendix 4 for the resulting sector breakdown). Not considered on the traded side are agriculture, forestry and fishing due to distortions caused by regional arrangements such as the Common Agricultural Policy. On the non-traded side public administration, defence and compulsory social security are excluded, due to the difficulty in controlling for large shifts in public sector employees.

To control for other broad Macroeconomic developments that affect the real exchange rate a number of "control variables" are considered. These, in part, intend to capture real exchange rate movements off an equilibrium path that is mainly determined by productivity developments (*BS*). Variables considered are the terms of trade, government expenditure-to-

GDP ratio, GDP per capita, openness, and money-to-GDP ratio. As we are examining the *REER*, i.e. relative prices, all variables are computed relative to the Eurozone. (see Appendix 7 for a discussion on data construction and the economic rationale of including these variables).

Preliminary Analysis

Visual inspection of the *REER* indices in Appendix 3, reveals that over the past eleven years exchange rates in Central Europe have appreciated between 23% (Poland) to 28% (Slovakia) in real terms. This upward trend has been stable in all countries, with the exception of Poland, which experienced a prolonged period of real depreciation during 2001-2003. Meanwhile, the sectoral productivity indices in Appendix 4 reveal that productivity growth in the traded sector has outpaced non-tradable productivity over the sample period (with the exception of Slovakia, 1996Q3-1997Q3). The average productivity growth gap ranged from 2.4% in Slovakia to 4.9% in Poland. According to B-S, sectoral productivity growth differentials between the traded and non-traded sector, should induce wage growth to outpace productivity growth in the non-tradable sector, due to cross sector wage equalisation. This in turn produces non-tradable inflation in excess of tradable inflation. Visual inspection of the sectoral inflation series in Appendix 6 generally confirms this picture, providing further preliminary evidence of the presence of B-S.

Yet, as previously highlighted, domestic productivity gaps alone are not sufficient to bring about real appreciation in line with B-S. Instead, we need to take into consideration the foreign sectoral productivity growth differential. The corresponding graphs in Appendix 5 indicate that this differential has been substantial in the euro-area (the average productivity growth gap ranged from 1% in Italy to 2.4% in France). Yet, studies using relative non-tradable prices as their objective variable fail to account for this when constructing their coefficient estimates.

Another striking feature in the data is that, while tradable productivity has grown slower in the Eurozone compared to Central Europe, non-tradable productivity has also been considerably lower. Hence, by assuming non-traded productivity to be equal across countries, Égert et al (2003) obtain biased estimates of the B-S effect in Central Europe. In particular, their B-S coefficient is likely to have a positive bias, overstating the Balassa-Samuelson effect. This is illustrated in Figure 1 below, which captures the basic productivity developments in Central Europe and the Eurozone. The angles on the respective graphs (A to D) can be thought of as a rough proxy for the average productivity growth rate.[18] Note that Eurozone non-tradable productivity growth is set to zero here, in order to simplify analysis:

Figure 1: Productivity growth gaps in Central Europe and the Eurozone

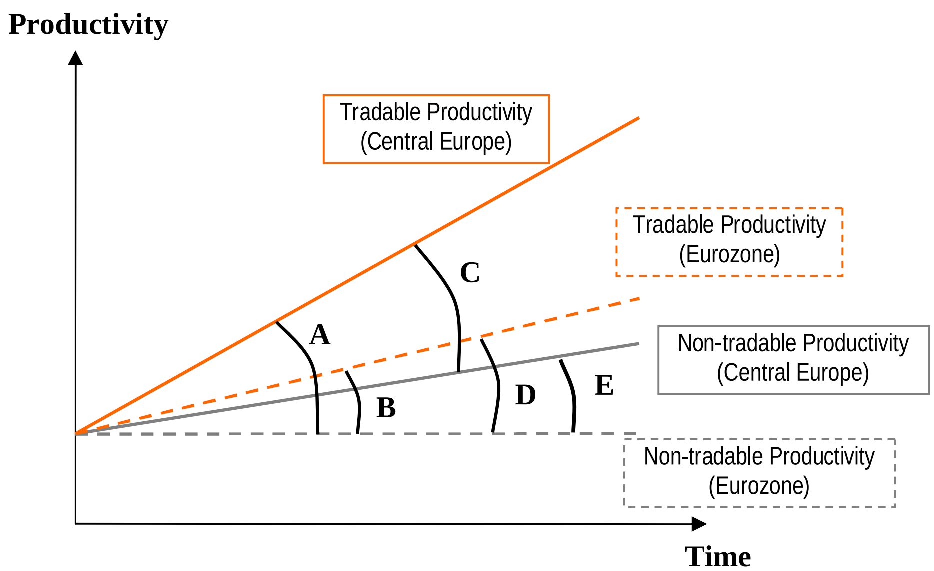


Figure 1 illustrates that, in comparing the differences in tradable productivity growth between Central Europe and Germany (A-B), Égert et al (2003) overstate the true sectoral productivity growth gap differential (C-D) by the amount E, which corresponds to the difference in non-tradable productivity growth.

Finally, Appendix 7 depicts the control variables mentioned earlier, plotted against the *REER* of the respective Central European country. Visual inspection suggests that the relative terms of trade and government expenditure-to-GDP ratio are most closely correlated with the respective *REER*s in Central Europe.

V. Empirical Analysis

To determine an equilibrium relationship between our variables, we start by inspecting their unit root behaviour. We are hereby able to take advantage of the increased data and power that the cross-section dimension of our data brings by implicitly treating variable observations in different countries as repeated draws from the same distribution. Thus, while Dickey-Fuller-type tests notoriously have low power in distinguishing the null of a unit root from stationary

[18] Note that a constant slope would correspond to a declining growth rate. This, however, does not alter the basic point the graph is trying to illustrate.

alternatives, using the cross-section dimension of our panel increases the power of tests compared to tests based on a single draw from the respective "population". Moreover, in contrast to standard DF-type tests, panel test statistics converge to normally distributed random variables as N and T become large, given that panel estimators average across individuals.

In testing for unit roots we adopt the methodology proposed by Im, Pesaran, and Shin (1995, 1997), which takes the following form:

$$\Delta y_{it} = \rho_i y_{i,t-1} + \sum_{L=1}^{p_i} \theta_{iL} \Delta y_{i,t-L} + z'_{it}\gamma + u_{it} \tag{5.1}$$

The null hypothesis is that each series in the panel contains a unit root for all i $(H_0 : \rho_i = 0)$, the alternative posits that at least one of the series in the panel is stationary. Essentially, the test averages the individual ADF test statistics obtained from estimating equation (5.1) for each i:

$$\bar{t} = \frac{1}{N}\sum_{i=1}^{N} t_\rho \tag{5.2}$$

Furthermore, IPS standardise their test statistics based on Monte Carlo simulations of the mean and variance of the t_ρ series underlying the $\bar{t}$ statistic, with different values obtained depending on N and the lag length used in the ADF tests. These simulated values are used along with equation (5.2) to obtain the z-statistic, which follows a standard Normal distribution:

$$z_{\bar{t}} = \frac{\sqrt{N}\left[\bar{t}_{NT} - N^{-1}\sum_{i=1}^{N} E(t_\rho)\right]}{\sqrt{N^{-1}\sum_{i=1}^{N} Var(t_\rho)}} \sim N(0,1) \tag{5.3}$$

(For further discussion on the IPS test, please refer to Appendix 8.)

To run the various tests the individual number of lags were chosen to minimise the Schwarz Information criterion.[19] The deterministic components were determined based on the sample information provided by the data.[20] The results, reported in Appendix 8, do not lead us to reject the null of non-stationarity for any of the variables at the 5% or 10% level, except *openness*. By rerunning the tests in first differences, we can infer that all variables are integrated of order one, apart from *openness*, which appears to be I(0).

Given that we are dealing with non-stationary data, standard inference based on t-statistics is no longer valid. Rather, finding an equilibrium relationship requires a combination of our variables

[19] Both the Schwarz (BIC) and Akaike Information Criterion (AIC) add a penalty which increases with the number of regressors. Because the penalty is larger for BIC, however, the latter tends to favour more parsimonious models.
[20] Thus, by not including a trend in a given test, it is implicitly assumed that the average rate of change for the respective variable is zero over the given sample period.

to be stationary to ensure that deviations from equilibrium are of temporary nature. In other words, we need to test for cointegration. In doing so the Engle and Granger approach to cointegration was adopted. This is a two-step approach, whereby a long-run equilibrium equation is estimated as a first step:

$$\log(REER)_{it} = \beta_0 + \beta_1 \log(BS)_{it} + varlist_{it} + u_{it} \tag{5.4}$$

where $varlist_{it}$ corresponds to the list of control variables and the residuals, u_{it}, are taken as a measure of disequilibrium, on which our test of cointegration is based. The second step involves estimating the corresponding Error Correction Model (ECM): a short-run model, in which all variables are made stationary through differencing; where the appropriate dynamics of each variable are included in order for the error term to be well-behaved (white noise); and in which we include the lagged residual of our long-run equation ($\hat{u}_{it-1}$):

$$\begin{aligned}\Delta \log(REER)_{it} = \phi_0 + \sum_{j=1} \phi_j \Delta \log(REER)_{it-j} \\ + \sum_{h=0} \delta_1 \Delta \log(BS)_{it-h} + \sum_{k=0} \delta_k \Delta varlist_{it-k} + \lambda \hat{u}_{it-1} + \varepsilon_{it}\end{aligned} \tag{5.5}$$

The lagged residual of our long-run equation hereby acts as our "error correction term", which captures the speed at which short-run deviations converge to their long-run equilibrium values. A negative and significant coefficient, λ, indicates that errors are self-correcting, and confirms that the relationship between the variables is indeed stationary. Note that testing for cointegration in panel data has the same advantageous effects in terms of power that were outlined in the context of the unit root tests conducted earlier.

Step 1: Finding a Long-run equation

The panel cointegration test procedure followed here is the method developed by Pedroni (1995, 1999), who uses the following model:

$$y_{it} = \alpha_i + \delta_t t + \beta_{1i} x_{1i,t} + \beta_{2i} x_{2i,t} + \ldots + \beta_{Ki} x_{Ki,t} + e_{it} \tag{5.6}$$

with tests for the null of no cointegration based on the residuals, e_{it}. Given that intercept and slope coefficients are allowed to vary across individuals, this approach allows for considerable short- and long-run heterogeneity. Pedroni (1999) developed several different tests with which to account for possible dynamics of the error term. The one chosen here is a parametric test similar to the ADF-type test, which allows the number of lags to be estimated directly in the model. This test, in essence, is analogous to the IPS (1997) unit root test outlined earlier: the ADF tests on the residuals of the long run equation are estimated separately across countries; by averaging the resulting test statistics and applying the mean and variance adjustments as outlined in Pedroni (1999) we obtain a test statistic, which follows a standard normal distribution.

Given that we are focusing on a specific set of countries that are not randomly selected from a large population, we would intuitively expect there to be a high correlation between the "fixed effects" of these countries and the regressors. This is confirmed by a formal Hausman test, which suggests that the fixed effects method of estimation is superior to a random effects method. Please refer to Appendix 9 for a discussion of different panel estimation methods and the Hausman test.

In trying to find a cointegrated long-run equation several combinations of variables were tested. For none of the combinations was it possible to reject the null of no cointegration at the 5% level (see Appendix 10). At the 10% level, however, the following relationship was found to be cointegrated:

$$\log(REER)_{it} = \underset{(0.00)}{2.03} + \underset{(0.11)}{0.41}\log(BS)_{it} + \underset{(0.18)}{0.91}\log\left(\frac{tot}{tot*}\right)_{it} + \underset{(0.12)}{0.24}\log\left(\frac{govt}{govt*}\right)_{it} + \varepsilon_{it} \tag{5.7}$$

where standard errors are given in parenthesis. Note that adding any additional variables to the long-run equation, did not make the error "sufficiently more stationary" to warrant their inclusion. Given that all our variables are expressed in logs, coefficient estimates can be interpreted as elasticities. Hence, according to equation (5.7), a ten percent increase in the sectoral productivity gap differential with the Eurozone is, on average, associated with a 4.1% percent increase in the *REER*, ceteris paribus. Similarly a ten percent increase in the relative terms of trade and the relative share of government expenditure in GDP are, on average, associated with a 9.1% and 2.4% increase in the *REER*, respectively.

Note that the long-run equation does not include any dynamics. Hence it is unsurprising that our error term is serially correlated. Similarly, the t-statistics in equation (5.7) are not normally distributed, preventing us from doing any inference based on their values. Yet, given that our variables are cointegrated, the OLS estimators can be proven to be super-consistent.[21]

Step 2: Constructing an Error Correction Model

In order to double check whether specification (5.7) is indeed cointegrated, we construct an error correction model, which takes into account the short-run dynamics of our variables. Given

[21] A formal proof of this is beyond the scope of this study but can be found in Banerjee et al. (1993), chapter 5, section 5.6.

that we are dealing with quarterly data, a first step in trying to capture the appropriate dynamics was to include four lags for each variable ($j = h = k = 4$) and then to iteratively drop the most insignificant variable.[22] Also, care was taken to ensure that the error remained well-behaved when dropping variables from the model. In doing so, the following ECM was obtained:

Table 1: Error Correction Model

log(*REER*)		**ECM**	
Regressors	*Coeff.*	*s.e.*	*p-value*
cons	0.001	0.001	0.236
log(*BS*)_diff	0.124	0.056	0.028
log(*BS*)_diff(-4)	-0.080	0.044	0.067
log(*tot*)_diff	0.444	0.121	0.000
log(*tot*)_diff(-1)	0.187	0.086	0.031
log(*open*)_diff	-0.073	0.040	0.071
log(*open*)_diff(-3)	0.034	0.016	0.030
log(*pcGDP*)_diff	0.438	0.066	0.000
log(*govt*)_diff(-3)	-0.099	0.041	0.018
$\hat{u}$(-1)	-0.144	0.046	0.002
R^2 (within)		0.617	

Note that the ECM includes variables, which were omitted in the long-run equation. Despite the fact that these variables were unable to explain a significant amount of long-run variation in the objective variable they may nonetheless be able to account for a significant amount of short-run variation. Also, *openness* was included into the regression as all variables in our short-run equation are stationary.

Given that variables enter the ECM in first differences, the interpretation of our coefficient estimates differ from the long-run equation (5.7). In particular, differences in logs represent changes in growth rates. Also, the ECM does not include any lagged explanatory variables, which simplifies coefficient interpretation. Interpreting the impact coefficient on log(*BS*)_diff, a one-percentage point increase in the growth of the sectoral productivity gap differential versus the Eurozone is, on average, associated with a 0.12 percentage point increase in the rate of change in the *REER*.

Most importantly, however, the above specification includes an error correction term, which is negative and significant (i.e. error correcting), whereby the size of the coefficient measures the speed of adjustment: short-run deviations from the long-run equilibrium, on average, “correct”

[22] Note that we are not able to drop more than one variable at a time without doing an F-test on the joint-insignificance of these particular variables

by 14.4% quarter-on-quarter. This result reaffirms our earlier conclusion that our variables are indeed cointegrated.

Diagnostic Testing

In order to test whether the residuals of our short-run model are white noise a number of robustness checks were performed. Conducting a Skewness and Kurtosis test, we obtain a test-statistic of 1.94, which follows an asymptotic chi^2 distribution under the null of normally distributed errors with two degrees of freedom. The corresponding p-value of 0.38 indicates that there is insufficient evidence to suggest that our residual distribution is non-normal:

Skewness/Kurtosis test for Normality			Joint Test	
Variable	Pr(Skewness)	Pr(Kurtosis)	adj. chi2(2)	Prob>chi2
ecmresid	0.277	0.392	1.940	0.379

This result is reinforced by inspecting the residual histogram in Appendix 12, which does not seem to contain any outliers or excess Kurtosis.

In conducting a serial correlation test of order 4 we are unable to reject the null of no serial correlation (p-value = 0.81). Similarly, testing for different orders of serial correlation does not yield alternative conclusions: testing for 1st, 2nd, 8th and 12th order serial correlation we obtained p-values for the joint insignificance of all lagged residuals of 0.96, 0.96, 0.61 and 0.38 respectively (see Appendix 12).

Finally, in order to avert any potential problems with heteroscedasticity robust standard errors were used in the estimation process.

VI. Evaluation of Results

Testing for the Balassa-Samuelson effect is a data intensive task and therefore results are likely to vary between studies, depending on the sources and types of data used, and the methods of data manipulation. Indeed, coefficient estimates appear to be very sensitive to the method of data construction. In particular, it matters what kind of weights are used in constructing the *REER*s and the proxies for the traded and non-traded productivity indices of the Eurozone. Also, in constructing the *BS* variable, it matters whether or not one scales the sectoral productivity growth gaps in each country by the share of the non-tradable sector, α, in overall GDP. To illustrate, table 3 presents different versions of long-run equation (5.7), whereby the dependent variable (*REER*) and the independent variable (*BS*) are constructed using country specific weights[23] as well as average weights, and where α is omitted in some specifications but not in others (please see footnote 25 for a discussion on different weights; for a more detailed table of the regression outputs in table 3 please refer to Appendix 13):

Table 3: Alternative *BS* and *REER* measures

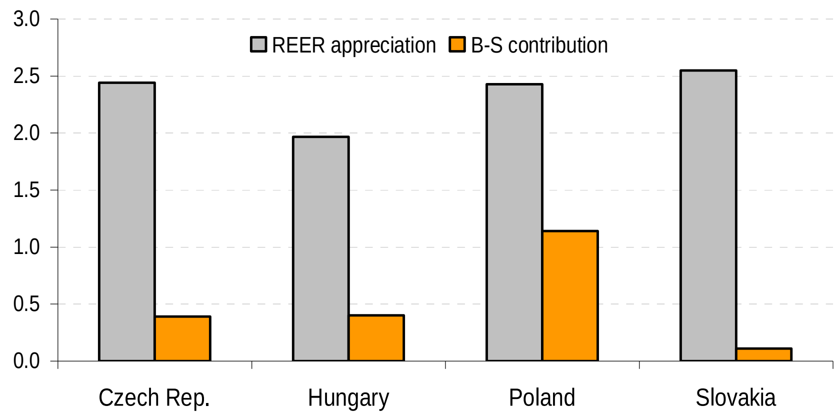
		BS1	***BS2***	***BS3***	***BS4***
		average weights		specific weights	
		excl. α	incl. α	excl. α	incl. α
REER1 average weights	$\beta 1$	0.58	0.59	0.60	0.73
REER2 specific weights	$\beta 1$	0.32	0.34	0.33	**0.41**

Table 3 illustrates that, depending on the method of data construction, estimates of the Balassa-Samuelson coefficient can vary by over 100%. Yet, in the existing literature testing for B-S these statistical nuances receive little if no attention. In particular, α is often omitted, implicitly assuming that the relative sizes of non-traded sectors are equal across regions. In the case of Central Europe and the Eurozone, however, this is an incorrect assumption (α=[0.55, 0.63], α*=[0.73]). Also, studies seem to have an inconsistency in the use of their trade weights: given that studies using the *REER* as their objective variable do not construct the respective indices themselves, weights are likely to be country-specific; in constructing the *BS* variable, however, it appears as though only one proxy for the foreign region (usually OECD) is constructed, pointing towards the use of average trade weights.

Using country-specific weights in constructing both variables, while controlling for the difference in the relative size of non-tradable sectors, equation (5.7) should produce the most accurate estimates of *BS* in Central Europe. In multiplying the coefficient on log(*BS*) by the average annual productivity growth gap differential of the respective Central European country, we attain the following estimates of the contribution of productivity differentials to inflation differentials versus the Eurozone during 1995Q1-2005Q3 (B-S):

[23] The use of country specific weights in constructing the *REER* implies that the *REER* in each Central European country is constructed using the respective country-specific trade weights with respect to the Eurozone countries. In the context of the *BS* variable, the use of country-specific weights entails constructing different Eurozone proxies for the traded and non-traded productivity indices for each of the Central European countries. In other words, the Eurozone productivity growth gap proxy used to scale the productivity growth gap in Poland will differ from the one used for the Czech Republic, given that their trade weights with the respective Eurozone countries are different. Thus, for example, given that the Czech Republic has a larger trade weight with Germany than Poland (Cz=73%, Pl=61%), strong non-tradable inflation in Germany associated with a large productivity growth gap will have a larger impact on the *REER* in the Czech Republic and, hence, should also imply that its *effective* productivity growth gap differential with the Eurozone is smaller, ceteris paribus.

Figure 2: The B-S effect in Central Europe (% pts per annum)



	Czech Rep.	Hungary	Poland	Slovakia
B-S contribution (%pts)	0.39	0.40	1.14	0.11
REER appreciation (%)	2.44	1.97	2.43	2.55

NB: Maastricht criterion for inflation convergence states, a country joining EMU should have an inflation rate no more than 1.5% above the average of the three lowest rates of inflation in the EU.

In the context of Figure 2, recall that movements in the *REER* can be interpreted as inflation differentials expressed in a common currency.[24] Figure 2 illustrates that the role of productivity differentials in explaining inflation differentials was most prominent in Poland, where B-S was able to account for almost 50% of the average annual inflation differential observed vis-à-vis the Eurozone (1.14% points out of 2.43%); conversely, in Slovakia B-S seems to account for no more than 5%. Looking at the size of B-S over the past years individually does not suggest any strong up- or downward trends in Central Europe. Rather there seems to be a more cyclical pattern (see Appendix 14). Moreover, the graphs show that averaging the B-S effect over time gives us a somewhat distorted picture of the impact of productivity differentials on inflation differentials. In Slovakia, for example, large negative sectoral productivity growth gaps have led to proposed negative impacts on the real exchange rate (see Appendix 14 for more on this). Thus by averaging the B-S effect over time, positive and negative influences act to offset each other. Nevertheless, averages do give us an indication of the positive net effect that productivity developments have had on the real exchange rates in Central Europe.

[24] Actual inflation differentials with Central Europe have declined substantially in recent years. Nominal currency appreciation, however, has played a pivotal role in reducing inflation. During ERM II this would no longer be possible.

Table 4 presents a number of comparable studies estimating B-S in transition economies. The most comparable is the one by Egert et. al (2002), which uses quarterly data to investigate five Central European transition economies (Czech Rep., Hungary, Poland, Slovakia, Slovenia) over the period 1991-2001. This study estimates B-S to be larger compared to the findings presented here.

Table 4: Study Comparison

	Data	**Explanatory variable**	**Estimation method**	**B-S effect (% points)**
Coricelli-Jazbec (2001)	19 transition econ's 1990-1998	P^N/P^T	Fixed Effects Panel coint.	0.9 - 1.2
De Broeck-Slok (2001)	25 transition econ's 1993-1998	*REER*	Pooled mean group est.	0.2 - 0.6
Egert et al. (2002)	5 transition econ's 1991Q1-2001Q2	*RER (DM)*	VAR and Panel coint.	0.5 - 3.5
Fischer (2002)	10 transition econ's 1993-1999	*REER*	Fixed Effects Panel coint.	0.7 - 2.2
Halpern-Wyplosz (2001)	9 transition econ's 1991-1998	P^N/P^T	GLS	3

Most of the difference is likely to be attributable to the difference in data sets. Egert et al. (2002) proxy the tradable sector solely through industrial production and set non-tradable productivity growth equal to zero, implicitly assuming non-tradable productivity growth to be equal in Central Europe compared to Germany. As shown earlier, however, non-tradable productivity growth has been substantially stronger in Central Europe compared to Germany (and the Eurozone as a whole), causing Egert's results to be biased and inconsistent.

The size of this bias is partially clouded by Egert's choice of objective variable (RER DM). The sectoral productivity growth gap in Germany averaged 2.4% p.a. compared to 2.1% in the Eurozone. Hence, our estimates for B-S would have been even lower had we used the RER (DM) as our dependent variable.

The measurement bias caused by using industrial production to proxy the tradable sector is less clear cut. In Hungary industrial productivity exaggerates tradable productivity growth, whereas the opposite is true in Poland. In the Czech Republic and Slovakia the discrepancy is comparatively small (see Appendix 15). Finally, including the early 1990s in their data set is likely to have caused additional omitted variable bias, on account of the large structural changes which occurred at the time (Halpern-Wyplosz, 2001; Coricelli-Jazbec, 2004).

Despite the fact that the B-S effect seems to be weaker compared to other studies, it remains statistically highly significant, even when including various control variables into the short-run equation. Nevertheless, it seems clear from Figure 2 that supply-side factors alone are not able to account for the inflation differentials observed between Central Europe and the Eurozone.

Comparing our long-run and short-run equations, we obtain the plausible result that demand side factors dominate in the short-run (β_{BS}=0.12; β_{pcGDP}=0.44), whereas in the long-run supply factors take the lead (β_{BS}=0.41; β_{Govt}=0.24). Note also that the positive coefficients on the demand side factors support the proposition of a bias towards non-tradables.

VII. Conclusion

The Central European transition economies have experienced strong *REER* appreciation during their transition to market-based economies. The purpose of this study was to evaluate to what extent this was the result of an equilibrium adjustment, in order to judge whether the convergence criteria remain appropriate for these newly acceded countries. Inflation differentials associated with B-S reflect an equilibrium phenomenon and do not require a policy response. Rather than calling for flexibility of nominal exchange rates, the presence of a strong B-S effect may thus underline the drawbacks of the Maastricht criterion on inflation for Central European countries.

The B-S effect is clearly evident in all Central European accession countries, while not always at the level suggested by earlier studies. Moreover, B-S can be expected to persist in the medium- to long-run as the Central European economies continue on their convergence path with their western counterparts.

It seems clear, nevertheless, that supply factors alone are not able to account for the inflation differentials observed vis-à-vis the Eurozone. Equilibrium *REER*s seem to also depend on other factors, some of which are in fact actively managed by the respective countries, such as government spending. Indeed, Fischer et al. (1998) suggest that government spending for CE countries could plausibly decline by a considerable amount in the future. This in turn could act to offset the inflation pressure induced by supply-side factors, such as B-S. This also raises the issue of potential welfare costs associated with euro-adoption, should governments be forced to reduce government expenditure in the face of increased inflation pressure.

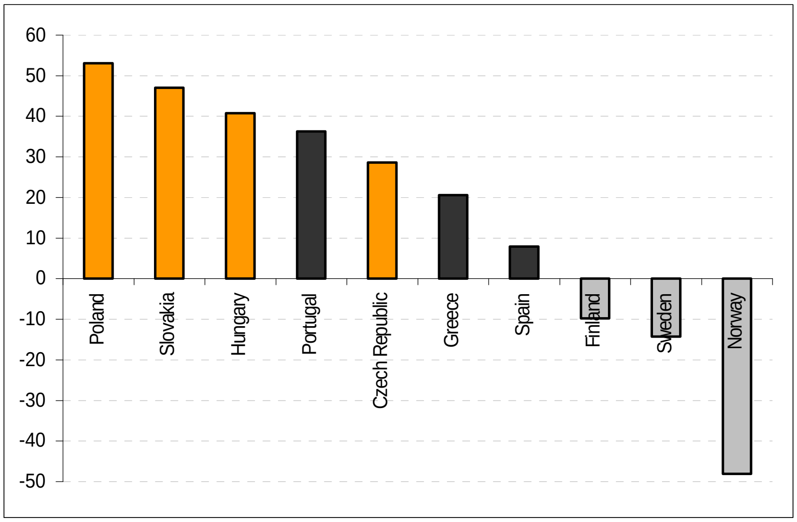
On the other hand, the tendency for real appreciation, particularly during ERM II, is likely to be reinforced by capital inflows, which will affect the *REER* both via the nominal exchange rate and via the B-S effect (empirical evidence suggests that FDI tends to raise productivity growth in the tradable sector considerably more than in the non-traded sector). Hence, in the words of Halpern-Wyplosz (2001, p.15), "the risk of currency crises in acceding countries is far from negligible."

The question arises whether a specific policy response can be prescribed to the EMU accession process of transitional economies. The "one-size-fits-all" regime of the Maastricht criteria may not be a sufficient approach to the upcoming expansion of EMU membership. Indeed, recent and current violations of Maastricht by some founding members of the EMU have already required "flexible" response policies, raising expectations for a similar, more flexible, approach towards accession candidates with particular transitional characteristics.

One option would be to ease the inflation constraint on accession countries, possibly by adjusting the current benchmark on inflation. Perhaps, one could consider tying the inflation constraint to the three countries with the highest inflation rates in the Eurozone, rather than the to the lowest inflation rates in the EU. In doing so, inflation rates would have to converge to the likes of Greece, Spain and Portugal, rather than Finland, Sweden and Norway. The former are countries with lower per capita incomes and higher underlying growth rates - similar to what we observe in Central Europe (see Figure 3). Conversely, the latter have per capita GDPs well in excess of the Eurozone average. Thus one could think of Greece, Spain and Portugal to be in a more comparable stage of "catch-up" with the remaining Euro-area countries, making them a natural choice against which to benchmark the newly acceded Central European transition economies. After all, why should countries such as Sweden and Norway, which are not members of the EMU, provide a benchmark for Eurozone accession candidates?

Nonetheless, the advantages in adjusting current convergence criteria to better "fit" current economic developments need of course be carefully weighed against potential costs involved in doing so. One could argue that amending existing criteria may not damage their credibility more than tolerating further breaches in the future. Yet increased inflation differentials within the Eurozone will make a unified monetary policy increasingly difficult to manoeuvre. Hence it may well be in the interest of current EMU members not to amend current convergence criteria, but to force postponement of accession until the respective candidates have largely concluded their path of convergence. Certainly this would serve the fundamental objective of the EMU to keep inflation pressures in check.

Figure 3: Per Capita GDP differential versus Eurozone (2005)



Source: Eurostat
NB: data is expressed in relation to Eurozone = 0

VIII. Limitations and Possible Improvements

Using the trade weighted real effective exchange rate versus the Eurozone as our objective variable added a considerable amount of accuracy to our estimates. One limitation of this approach, however, is that the *REER*s were based on constant weights, despite trade patterns having continued to shift significantly throughout the transition process. Hence, improved estimates of *BS* would be obtainable by allowing trade weights to vary over time.

Further improvements could also be made with respect to our GDP per capita time series obtained form Eurostat. Despite being labelled as quarterly data, visual inspection suggests estimation frequency to be annual. Thus, on average, an improved proxy for quarterly figures should be feasible by "smoothing out" the series, prior to using the variable for estimation purposes.

Finally, another noteworthy point, which needs to be considered when interpreting estimation results, is the use of quarterly data. In order to test for the responsiveness of coefficient estimates with respect to data frequency, De Gregorio et al. (1994) re-estimate their results, allowing data frequency to decline by averaging observed growth rates. In doing so, the coefficient on productivity differentials increases, whereas the coefficients on demand side factors become insignificantly different from zero. Data frequency is thus likely to affect coefficient estimates not only by providing more degrees of freedom, thereby improving the reliability of estimation results and enabling a smaller cross-section dimension. Rather, higher frequency data is also likely to conceal longer-run trends in the data. In other words, it may be more reasonable to think of perfect capital mobility to hold over the long-run, and thus the supply curve becoming flatter as the observation frequency declines.

Appendix 1

The Balassa-Samuelson Effect in Industrialised Countries

A paper by the NY Fed (2001) suggests that nearly two-thirds of the US dollar's performance in the 1990s can be explained by cross-country sectoral productivity gaps. It indicates that productivity growth in the tradable sector far surpassed growth in the non-tradable sector and that the ratio of productivity growth between tradables and non-tradables has been consistently widening in the US over the past thirty years.

Likewise in Japan, which experienced a real yen-dollar appreciation of 82 per cent since 1973, the leading explanation is the increase in the relative prices of non-tradables brought about by substantial productivity increases in its tradable sector (Marston, 1987).

Using a panel of OECD countries, Cannzoneri et al. (1999) find supportive evidence of the connection between sectoral productivity growth gaps and relative non-tradables inflation. Their evidence on PPP in the traded sector, however, is less favourable and seems to vary depending on the exchange rate "cross" used. In a later paper (2002) the authors estimate the size of inflation differentials arising from B-S for EMU countries. Their findings suggest significant sectoral productivity gaps across the EMU and predict regional inflation differentials five-times the size compared to the US.[25]

[25] The US regions in this study corresponded to aggregations used by Bayoumi and Eichengreen (1993).

Appendix 2

Relative Non-tradable Prices as a Proxy for the Real Exchange Rate

The justification of using non-tradable prices as a proxy for the real exchange rate rests on the assumption that the sectoral productivity growth gap in the "foreign country" is insignificant and therefore that relative non-tradable prices in that country are constant.

To see this it is best to reconsider equation (3.2), reproduced below

$$P = \left(P_T\right)^{\alpha}\left(P_{NT}\right)^{1-\alpha}$$

$$P^* = \left(P_T^*\right)^{\alpha *}\left(P_{NT}^*\right)^{1-\alpha *} \tag{A2.1}$$

Rearranging we are able to obtain:

$$P = P_T\left(\frac{P_{NT}}{P_T}\right)^{1-\alpha}$$

$$P^* = P_T^*\left(\frac{P_{NT}^*}{P_T^*}\right)^{1-\alpha *} \tag{A2.2}$$

Finally substituting this into the definition of the real exchange rate gives us:

$$RER = \frac{eP^*}{P} = \frac{eP_T^*}{P_T} \times \frac{\left(P_{NT}^* / P_T^*\right)^{1-\alpha^*}}{\left(P_{NT} / P_T\right)^{1-\alpha}} \tag{A2.3}$$

Hence assuming that purchasing power parity holds for traded but not for non-traded goods:

$$P_T = eP_T^*$$

$$\hat{P}^{NT} \neq e\hat{P}^{NT*} \tag{A2.4}$$

and that relative non-tradable prices are constant at some level c in the foreign country:

$$P_{NT}^* / P_T^* = c \tag{A2.5}$$

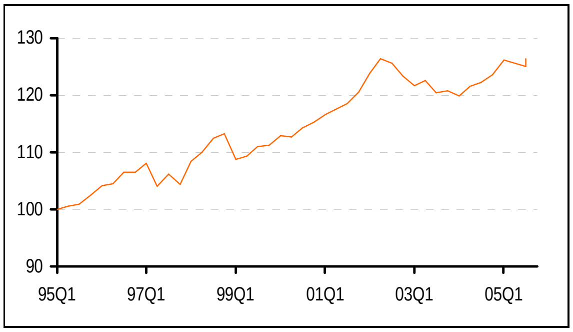
then changes in the real exchange rate are solely determined by changes in the relative price of non-tradables in the home country.

Section IV, however, illustrates that the sectoral productivity growth gap in the Eurozone has been substantial, rendering assumption (A2.5) incorrect. Thus, ignoring inflation developments in the Eurozone will lead to biased coefficient estimates of the Balassa-Samuelson effect in Central Europe.

Appendix 3

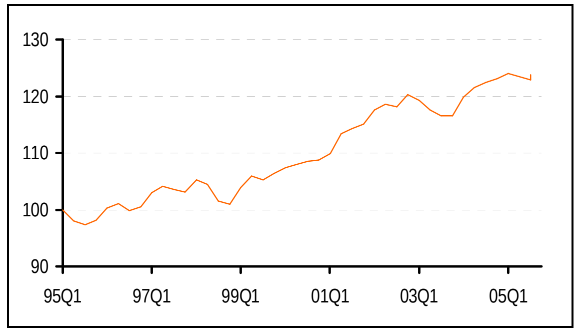
CPI-based, trade weighted real effective exchange rates vis-à-vis the euro-area (REER)

Chart 1: Czech Rep.: Real Effective Exchange Rate



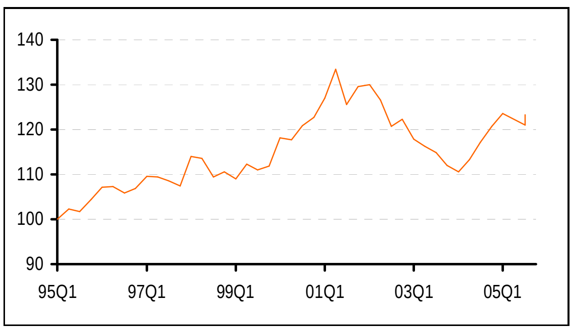
Source: Czech Statistics Office, OECD MEI, IMF DOTS, Datastream

Chart 2: Hungary: Real Effective Exchange Rate



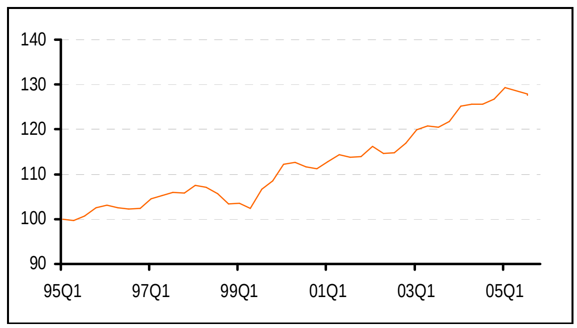
Source: Hungarian Statistics Office, OECD MEI, IMF DOTS, Datastream

Chart 3: Poland: Real Effective Exchange Rate



Source: Polish Statistics Office, OECD MEI, IMF DOTS, Datastream

Chart 4: Slovak Rep.: Real Effective Exchange Rate



Source: Slovak Statistics Office, OECD MEI, IMF DOTS, Datastream

Data constructed using:

EU harmonised consumer price index series for Ger, Fra, It, Nld and the respective CE economy (Cz, Hu, Pl, Sk);

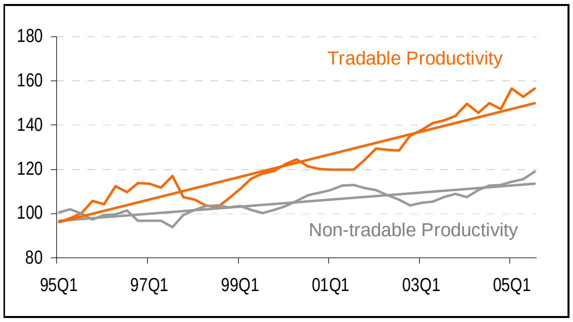
Euro exchange rates (synthetic index prior to 1999);

Trade weights from the IMF Direction of Trade statistics.

Appendix 4

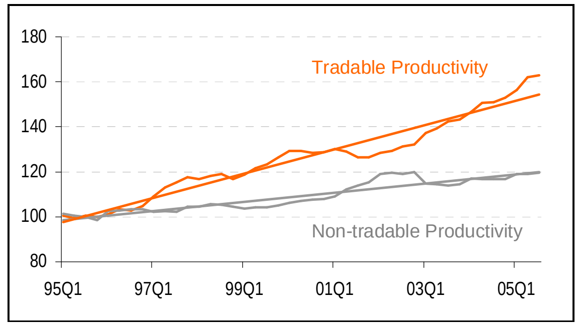
Sectoral productivity indices in Central Europe

Chart 5: Czech Rep.: Tradable vs. Non-tradable Productivity



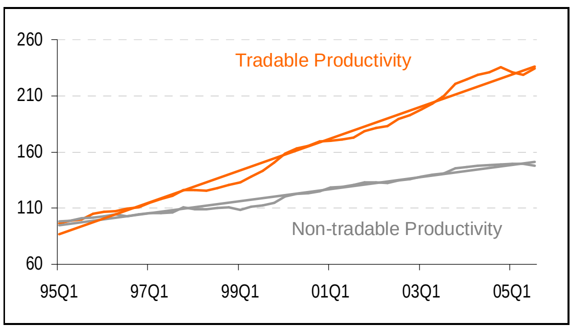
Source: Czech Statistics Office

Chart 6: Hungary: Tradable vs. Non-tradable Productivity



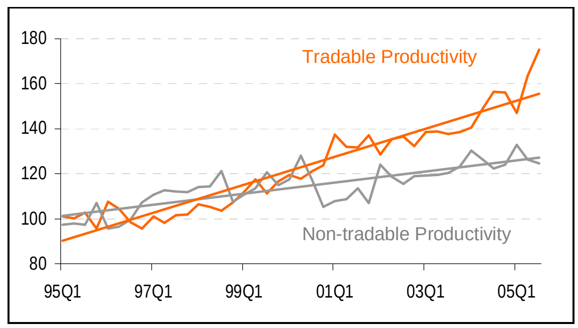
Source: Hungarian Statistics Office

Chart 7: Poland: Tradable vs. Non-tradable Productivity



Source: Polish Statistics Office and Eurostat

Chart 8: Slovak Rep.: Tradable vs. Non-tradable Productivity



Source: Slovak Statistics Office

Data constructed using:
Sectoral output (GVA) and total employment data. All data series were seasonally adjusted either by the respective statistics office or using "Census 1" (see Appendix 16). Sector weights reflect the respective sector's weight in overall production.

The data was subsequently grouped into traded and non-traded sectors as follows:

Traded sector:
Manufacturing; mining; hotels; transportation and communications.
Not considered: agriculture, forestry and fishing due to distortions caused by CAP.

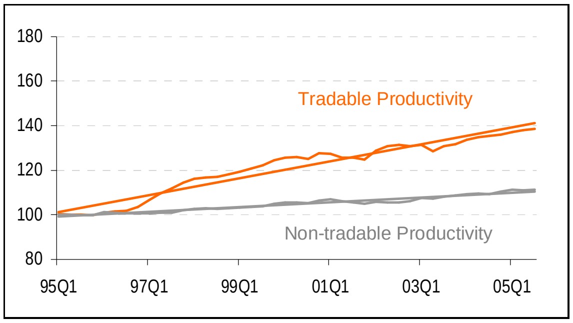
Non-traded sector:
Construction; electricity, gas and water; wholesale, retail trade and repair; education; financial intermediation; health and social work; real estate, renting and business activities; other community, social and personal activities.
Not considered: public administration, defence and compulsory social security, due to difficulty in controlling for large shifts in public sector employees.

NB: for recent years (2001-2005) the above data is freely available over the internet from the respective national statistics offices. Most of the remaining data, however, was available only upon request, via fax or email.

Appendix 5

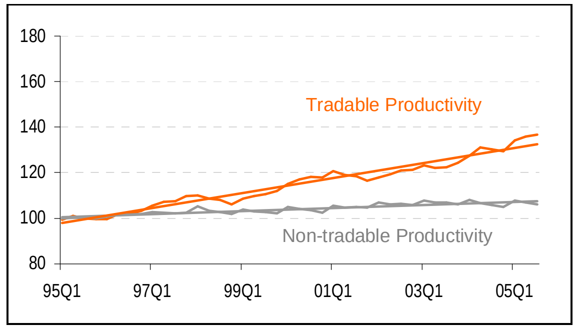
Sectoral productivity indices in the Eurozone

Chart 9: France: Tradable vs. Non-tradable Productivity



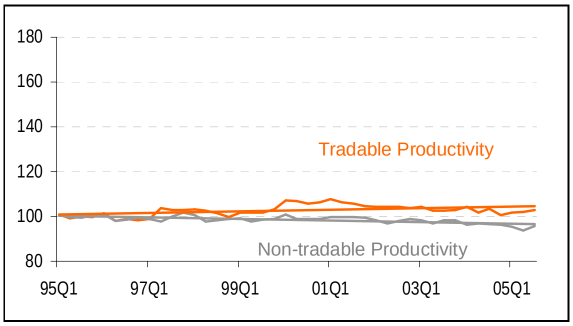
Source: Eurostat and French Statistics Office

Chart 10: Germany: Tradable vs. Non-tradable Productivity



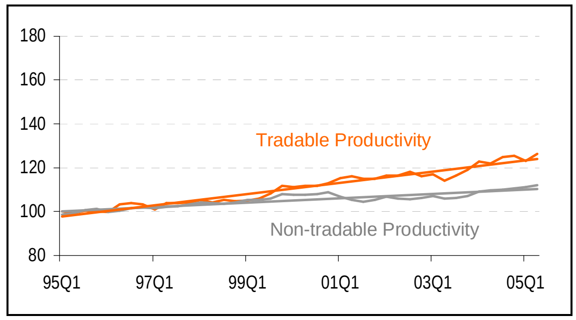
Source: Eurostat and German Statistics Office

Chart 11: Italy: Tradable vs. Non-tradable Productivity



Source: Eurostat

Chart 12: Netherlands:Tradable vs.Non-tradable Productivity



Source: Eurostat and Dutch Statistics Office

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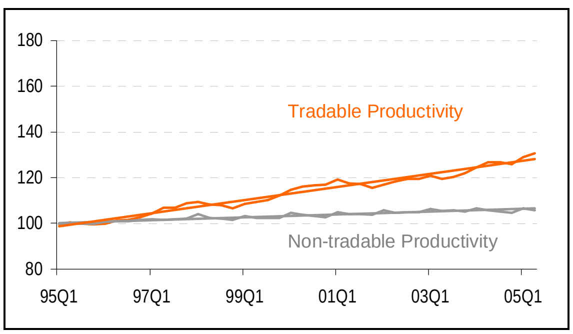
Chart 13: Eurozone: Tradable vs. Non-tradable Productivity



Source: Weighted average of France, Germany, Italy and the Netherlands.

Data constructed using:
Sectoral output (GVA) and total employment data, which was grouped into traded and non-traded sectors as outlined in Appendix 4.

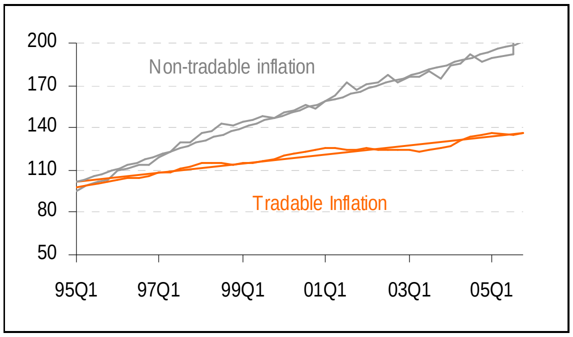
In chart 13, country weights reflect the respective country's average share in overall trade with Central Europe (Cz, Hu, Pl, Sk). For estimation purposes the data series in chart 13 were reproduced using country specific trade weights, resulting in four different Eurozone proxies (one for each of the respective Central European countries – see Appendix 13 for details).

NB: Sectoral employment data for Germany and the Netherlands were available only upon request from the respective national statistics offices.

Appendix 6

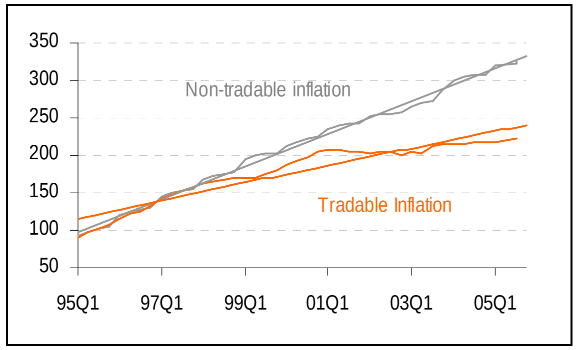
Sectoral price indices for Central Europe

Chart 14: Czech Rep.: Tradable vs. Non-tradable Inflation



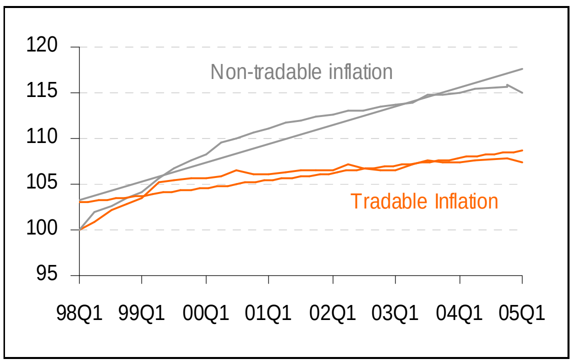
Source: Czech Statistics Office

Chart 15: Hungary: Tradable vs. Non-tradable Inflation



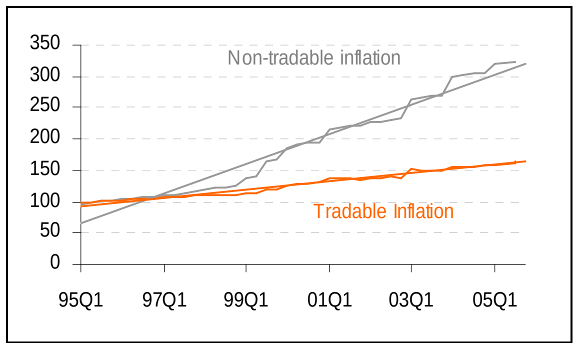
Source: Hungarian Statistics Office

Chart 16: Poland: Tradable vs. Non-tradable Inflation



Source: Polish Statistics Office

Chart 17: Slovak Rep.: Tradable vs. Non-tradable Inflation



Source: Slovak Statistics Office

Data constructed using:

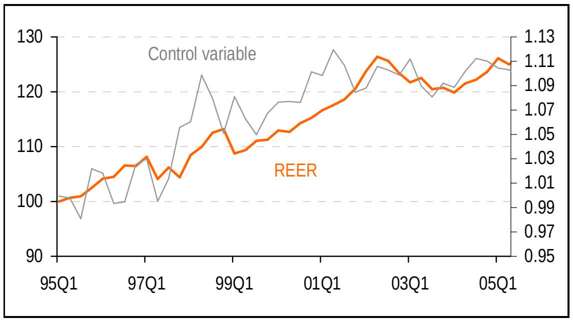
Month-on-month growth rates of the CPI-breakdown. Tradable and non-tradable sectors were grouped in line with the sector breakdown presented in Appendix 4.

Appendix 7

Control Variables (relative to Eurozone)

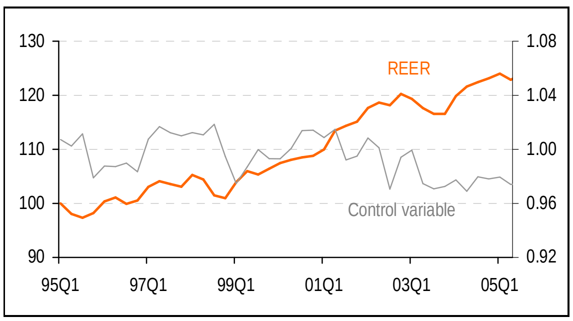
Terms of Trade

Chart 18: Czech Rep.: Terms of Trade



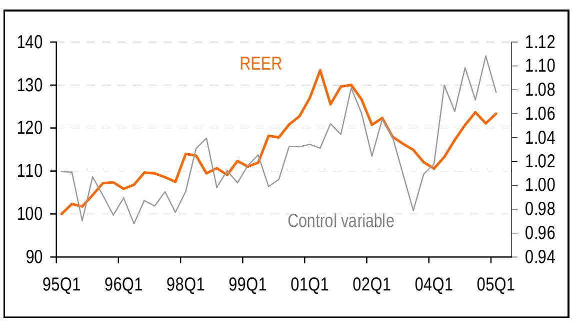
Source: Czech Statistics Office, IMF IFS, Eurostat

Chart 19: Hungary: Terms of Trade



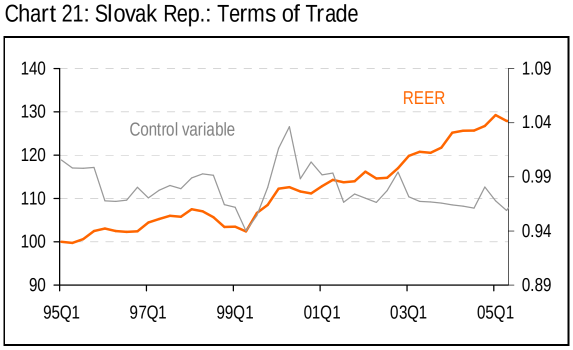
Source: IMF IFS, Eurostat

Chart 20: Poland: Terms of Trade



Source: IMF IFS, Eurostat

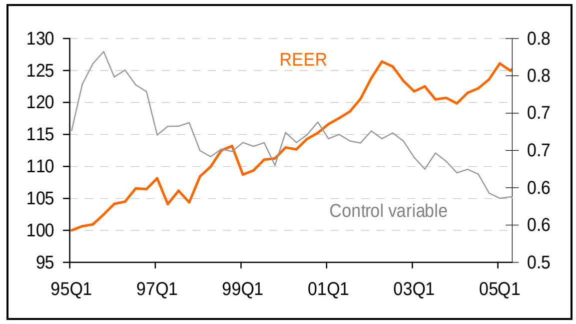
Chart 21: Slovak Rep.: Terms of Trade



Source: Slovak Statistics Office, IMF IFS, Eurostat

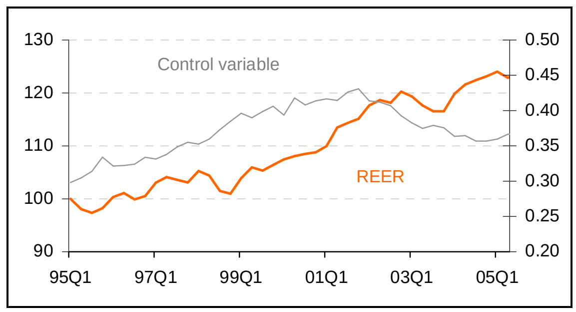
Money-to-GDP ratio

Chart 22: Czech Rep.: Money-to-GDP ratio



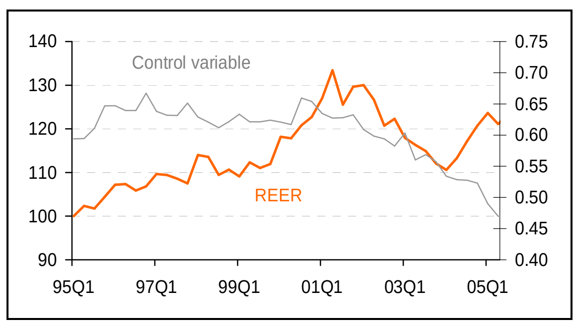
Source: IMF International Financial Statistics

Chart 23: Hungary: Money-to-GDP ratio



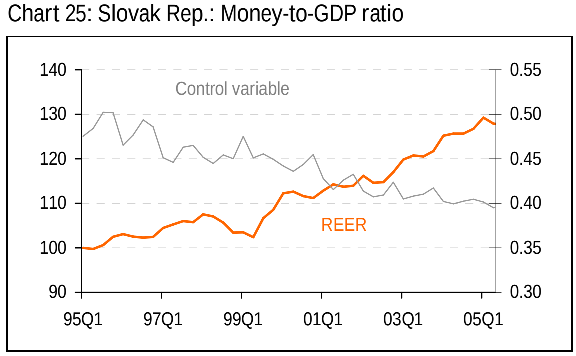
Source: IMF International Financial Statistics

Chart 24: Poland: Money-to-GDP ratio



Source: IMF International Financial Statistics

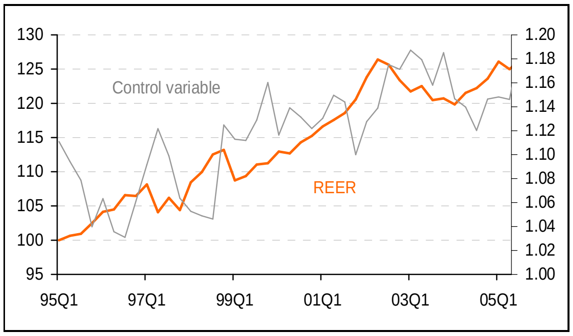
Chart 25: Slovak Rep.: Money-to-GDP ratio



Source: IMF International Financial Statistics

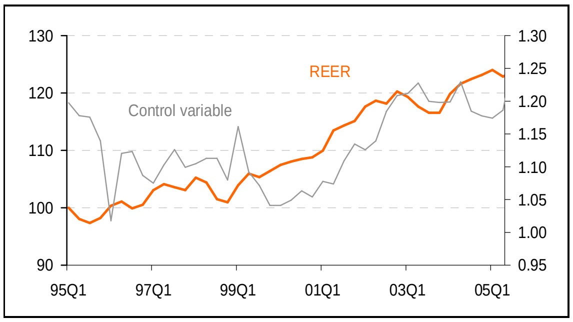
Government Expenditure-to-GDP ratio

Chart 26: Czech Rep.: Govt. Exp.- to- GDP ratio



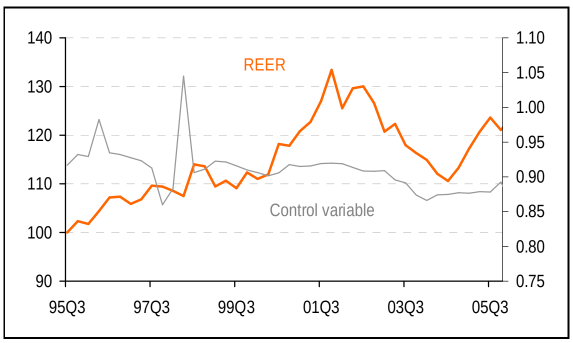
Source: Eurostat and French Statistics Office

Chart 27: Hungary: Govt. Exp.- to- GDP ratio



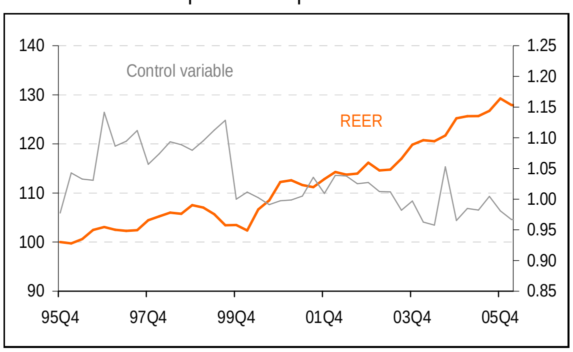
Source: Eurostat

Chart 28 : Poland: Govt. Exp.- to- GDP ratio



NB: Weighted average of France, Germany, Italy and the Netherlands.

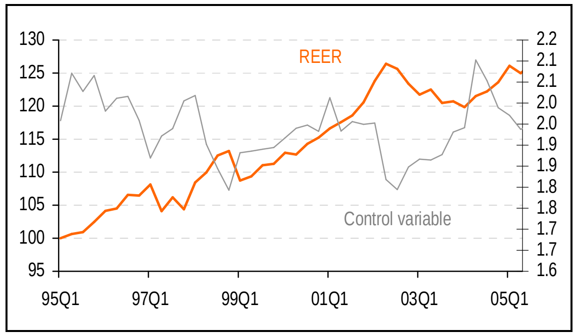
Chart 29: Slovak Rep.: Govt. Exp.- to- GDP ratio



Source: Polish Statistics Office

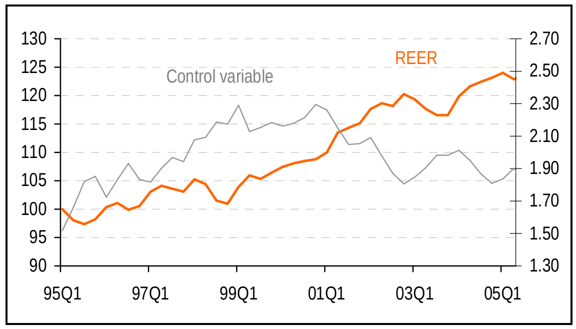
Openness

Chart 30: Czech Rep.: Openness



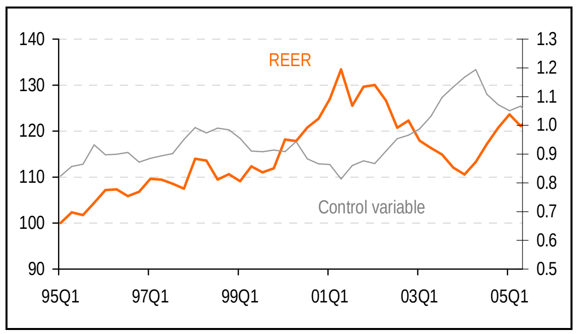
Source: Eurostat and German Statistics Office

Chart 31: Hungary: Openness



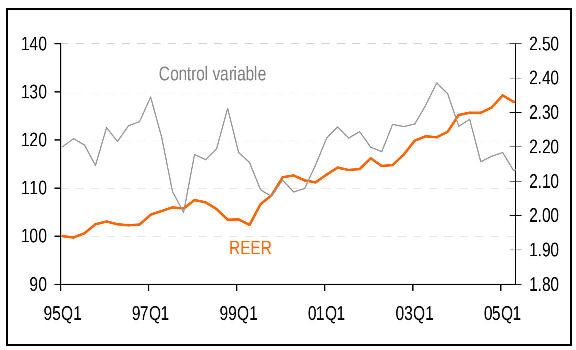
Source: Eurostat and Dutch Statistics Office

Chart 32: Poland: Openness



Source: Polish Statistics Office

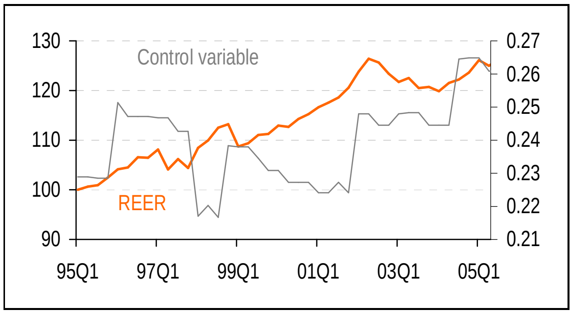
Chart 33: Slovak Rep.: Openness



Source: Polish Statistics Office

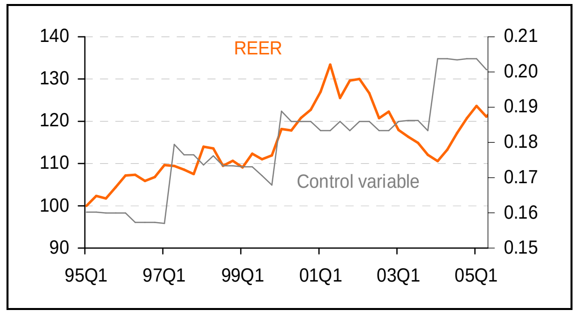
GDP per capita

Chart 34: Czech Rep.: GDP per capita



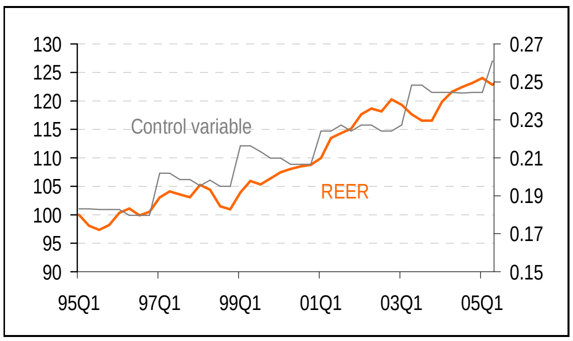
Source: Eurostat and French Statistics Office

Chart 35 : Poland: GDP per capita



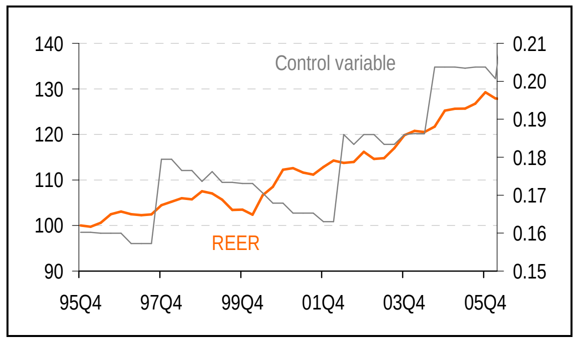
NB: Weighted average of France, Germany, Italy and the Netherlands.

Chart 36: Hungary: GDP per capita



Source: Eurostat

Chart 33: Slovak Rep.: GDP per capita



Source: Polish Statistics Office

NB: All variables are expressed relative to the Eurozone

Terms of Trade = export prices over import prices. Export (import) prices were obtained by dividing nominal exports (imports) by real exports (imports).

Rationale: Given differentiated products, export and import price inflation will not be of the same magnitude in practice. The terms of trade control for this effect.

Source: National Statistics Offices; International Financial Statistics; Eurostat

Government Expenditure-to-GDP ratio = Final consumption expenditure of general government divided by GDP (all in real terms).

Rationale: The share of government expenditure in GDP reflects the impact of fiscal policy on the *REER* through changes in relative demand for non-tradables. Empirically, government expenditure is biased towards non-traded goods and services, hence increased government expenditure should raise the relative demand for non-tradables, resulting in an increase in the relative price of non-traded goods and services.

Source: Eurostat

GDP per capita

Rationale: Assuming non-homothetic tastes, demand may affect relative price of non-tradables through altering the composition of output. Given services (the majority of non-traded goods) are luxuries in consumption, whereas commodities (mostly traded) are necessities, GDP growth causes consumers to spend a relatively higher proportion of their additional income on non-tradables causing an increase in the relative price level of non-tradables (Bergstrand;1991).

Source: Eurostat; National Statistics Offices

Openness = exports plus imports divided by GDP.

Rationale: The degree of openness of the economy determines by how much the real exchange rate responds to monetary and real shocks, the response is larger in more closed economies (De Broeck-Slok; 2001).

Source: Eurostat

Money-to-GDP ratio = broad money (M3) divided by GDP.

Rationale: Captures monetary shocks that, in conjunction with nominal rigidities, translate into real exchange rate movements (De Broeck-Slok; 2001).

Source: International Financial Statistics

NB: Some of the explanations for the economic rational of the control variables were taken directly from De Broeck-Slok (2001). A more thorough discussion can be found in Mussa (1984) and Frenkel and Mussa (1985).

Appendix 8

Unit Root Testing

The IPS test is comparable to the test by Levin and Lin (1992, 1993) but relaxes the homogeneity constraint imposed by LL, by allowing ρ_i to vary across the i cross-section units. LL also do not allow for different lags for the different cross-sections in the model when estimating (5.1). The main drawback of both the IPS and the LL tests, however, is that both tests assume that each i is cross-sectionally independent, implying that there is no cointegration between pairs or groups of individuals (countries) in the cross sections. This assumption is particularly questionable when dealing with financial variables for different countries, where we may assume different countries to be economically linked in terms of long-run equilibrium relationships. As a consequence the panel unit root tests may over-reject the null of non-stationarity (see Banerjee, Cockerill and Russell; 2001),

REER, constant, trend			Adj. terms		
	ADF stat	*p*	mean	variance	*z*-stat.
cz	-2.509	0	-2.173	0.655	
hu	-3.237	1	-2.177	0.687	
pl	-1.918	0	-2.173	0.655	
sk	-2.276	0	-2.173	0.655	
average	-2.485		-2.174	0.663	**-0.764**

govt, constant			Adj. terms		
	ADF stat	*p*	mean	variance	*z*-stat.
cz	-1.624	0	-1.523	0.770	
hu	-1.764	0	-1.523	0.770	
pl	-1.855	2	-1.476	0.830	
sk	-2.722	0	-1.523	0.770	
average	-1.991		-1.511	0.785	**-1.083**

B-S, constant, trend			Adj. terms		
	ADF stat	*p*	mean	variance	*z*-stat.
cz	-3.039	3	-2.117	0.735	
hu	-2.748	0	-2.173	0.655	
pl	-2.945	0	-2.173	0.655	
sk	-1.624	0	-2.173	0.655	
average	-2.589		-2.159	0.675	**-1.047**

pcGDP, constant, trend			Adj. terms		
	ADF stat	*p*	mean	variance	*z*-stat.
cz	-2.337	0	-2.173	0.655	
hu	-2.189	7	-1.986	0.872	
pl	-3.686	0	-2.173	0.655	
sk	-2.122	0	-2.173	0.655	
average	-2.584		-2.126	0.709	**-1.086**

tot, constant			Adj. terms		
	ADF stat	*p*	mean	variance	*z*-stat.
cz	-1.947	0	-1.523	0.770	
hu	-1.673	2	-1.476	0.830	
pl	-0.737	9	-1.329	0.996	
sk	-3.345	0	-1.523	0.770	
average	-1.926		-1.463	0.842	**-1.009**

m3, constant, trend			Adj. terms		
	ADF stat	*p*	mean	variance	*z*-stat.
cz	-2.998	0	-2.173	0.655	
hu	-1.204	0	-2.173	0.655	
pl	-1.288	0	-2.173	0.655	
sk	-5.304	3	-2.117	0.735	
average	-2.698		-2.159	0.675	**-1.313**

openness, constant			Adj. terms		
	ADF stat	*p*	mean	variance	*z*-stat.
cz	-2.769	0	-1.523	0.770	
hu	-2.966	0	-1.523	0.770	
pl	-1.604	0	-1.523	0.770	
sk	-2.767	0	-1.523	0.770	
average	-2.527		-1.523	0.770	**-2.288**

$$z\text{ - stat} = \frac{ADF - \mu}{\sqrt{var}} \sim N(0,1)$$

Appendix 9

Panel Estimation Techniques

When applying pooled OLS to panel data, the model does not distinguish between the within and the between dimension of the data. Besides a purely random component, which is assumed to be uncorrelated over time (ε_{it}), the error term is thus also likely to contain an individual specific component, which does not vary over time (α_i):

$$y_{it} = \mu + x_{it}'\beta + a_i + \varepsilon_{it} \qquad \text{(A9.1)}$$

Irrespective of whether or not this individual specific component is correlated with the explanatory variables, the composite error term, $\varepsilon_{it}+ \alpha_i$ will exhibit a particular form of autocorrelation (unless $\sigma_\alpha^2=0$). Hence, the standard errors for the OLS estimators will be incorrect. Intuitively, the coefficient estimates will be inefficient because in pooling the observations, OLS estimation ignores the structure of the error covariance matrix.

The fixed effects estimation, through the “within transformation”, eliminates any time invariant variables from the model by subtracting the averages from each of the variables:

$$y_{it} - \bar{y}_i = (x_{it} - \bar{x}_i)'\beta + (\varepsilon_{it} - \bar{\varepsilon}_i) \qquad \text{(A9.2)}$$

In doing so, estimated coefficients are identified only through changes in the respective variables over the sample period. Thus, effectively, the between-dimension is ignored and coefficient estimates are solely characterised through the within-dimension of the data.

Finally, the random effects estimator takes into account both the between- and the within-dimension of the data, treating α_i as random component of the error ($u_{it} = \varepsilon_{it}+ \alpha_i$). By applying OLS to the transformed model (A9.3):

$$y_{it} - \theta\,\bar{y}_i = \mu(1-\theta) + (x_{it} - \theta\,\bar{x}_i)'\beta + u_{it} \qquad \text{(A9.3)}$$

we are able to obtain the generalised least squares estimator, which in effect is the weighted average of the within and between estimators:

$$\hat{\beta}_{GLS} = \Delta\hat{\beta}_B + \left(I_k - \Delta\right)\hat{\beta}_{FE} \qquad \text{(A9.4)}$$

where Δ represents the efficient weighting matrix, proportional to the inverse of the covariance matrix of $\hat{\beta}_B$. In order for the coefficient estimates to be consistent, however, there must be no correlation between α_{it} and x_{it}.

Hausman Test: Fixed versus Random Effects:

As mentioned above, the appropriateness of the random effects estimator crucially depends upon whether or not the individual specific component of the error term (α_i) is correlated with one or more of the explanatory variables. The Hausman test is used to test for orthogonality of the random effects and the regressors. The null hypothesis is that the error term is uncorrelated with the explanatory variables:

$$\mathrm{H}_0 : E(\varepsilon_{it} x_{is}) = 0 \quad \text{for all } s,t \tag{A9.5}$$

Under the null both the fixed effects and the random effects estimations are consistent but fixed effects estimation is inefficient, whereas under the alternative, fixed effects estimation is consistent but random effects is not. Hence, under the null hypothesis the two estimates should not differ systematically, and a test can be based on the difference. Another important part of the test statistic is the covariance matrix of the difference vector $(\hat{\beta}_{FE} - \hat{\beta}_{RE})$:

$$Var(\hat{\beta}_{FE} - \hat{\beta}_{RE}) = Var(\hat{\beta}_{FE}) + \hat{V}ar(\hat{\beta}_{RE}) - 2Cov(\hat{\beta}_{FE}, \hat{\beta}_{RE}) \tag{A9.6}$$

Hausman's essential result hereby is that the covariance of an efficient estimator with its difference from an inefficient estimator is zero, which implies that:

$$\begin{aligned} &Cov(\hat{\beta}_{FE} - \hat{\beta}_{RE}, \hat{\beta}_{RE}) = Cov(\hat{\beta}_{FE}, \hat{\beta}_{FE}) - \hat{V}ar(\hat{\beta}_{RE}) = 0 \\ &\Rightarrow Cov(\hat{\beta}_{FE}, \hat{\beta}_{RE}) = \hat{V}ar(\hat{\beta}_{RE}) \end{aligned} \tag{A9.7}$$

Substituting (2.13) into (2.12), we obtain:

$$Var(\hat{\beta}_{FE} - \hat{\beta}_{RE}) = Var(\hat{\beta}_{FE}) - \hat{V}ar(\hat{\beta}_{RE}), \tag{A9.8}$$

which is the required covariance matrix for the test statistic, which is based on the Wald criterion:

$$\xi_H = (\hat{\beta}_{FE} - \hat{\beta}_{RE})' \left[\hat{V}(\hat{\beta}_{FE}) - \hat{V}(\hat{\beta}_{RE})\right]^{-1} (\hat{\beta}_{FE} - \hat{\beta}_{RE}) = 60.01 \sim \chi^2(3) \tag{A9.9}$$

Our result of 60.01 implies that the null hypothesis can be rejected at any meaningful level of significance.

In light of the test result we may conclude that the only consistent estimation is the fixed effects estimation, given that it eliminates the individual specific component of the error term, which is correlated with the explanatory variables.

Table 4: Hausman Test

Ho: difference in coefficients in not systematic				sqrt(diag(V_b-V_B))
Coefficients	**FE**	**RE**	**FE-RE**	**S.E.**
log(BS)	0.406	0.247	0.159	0.064
log(tot)	0.907	0.538	0.369	0.055
log(govt)	0.244	-0.024	0.268	0.129

b - consistent under Ho and Ha

B - inconsistent under Ho, efficient under Ha

chi2(3) = (b-B)'[(V_b-V_B)^(-1)](b-B) = 60.01

Prob>chi2 = 0.000

Appendix 10

Cointegration Testing

B-S, ToT

	ADF stat.	mean adj.	var. adj.	z -stat.
cz	-3.595	-1.662	1.559	
hu	-1.848	-1.662	1.559	
pl	-2.742	-1.662	1.559	
sk	-1.990	-1.662	1.559	
average	-2.544	-1.662	1.559	**-0.706**

B-S, Govt, pcGDP

	ADF stat.	mean adj.	var. adj.	z -stat.
cz	-3.518	-2.156	1.286	
hu	-3.240	-2.156	1.286	
pl	-2.427	-2.156	1.286	
sk	-4.095	-2.156	1.286	
average	-3.320	-2.156	1.286	**-1.026**

B-S, Govt

	ADF stat.	mean adj.	var. adj.	z -stat.
cz	-3.113	-1.662	1.559	
hu	-2.229	-1.662	1.559	
pl	-3.240	-1.662	1.559	
sk	-3.248	-1.662	1.559	
average	-2.957	-1.662	1.559	**-1.038**

B-S, ToT, Govt

	ADF stat.	mean adj.	var. adj.	z -stat.
cz	-3.821	-2.156	1.286	
hu	-4.035	-2.156	1.286	
pl	-3.021	-2.156	1.286	
sk	-3.665	-2.156	1.286	
average	-3.636	-2.156	1.286	**-1.305**

B-S, pcGDP

	ADF stat.	mean adj.	var. adj.	z -stat.
cz	-2.399	-1.662	1.559	
hu	-3.238	-1.662	1.559	
pl	-1.990	-1.662	1.559	
sk	3.014	-1.662	1.559	
average	-1.153	-1.662	1.559	**0.407**

B-S, ToT, Govt, pcGDP

	ADF stat.	mean adj.	var. adj.	z -stat.
cz	-3.686	-2.571	1.028	
hu	-3.601	-2.571	1.028	
pl	-3.763	-2.571	1.028	
sk	-4.340	-2.571	1.028	
average	-3.848	-2.571	1.028	**-1.259**

B-S, m3

	ADF stat.	mean adj.	var. adj.	z -stat.
cz	2.492	-1.662	1.559	
hu	0.075	-1.662	1.559	
pl	-2.071	-1.662	1.559	
sk	-4.133	-1.662	1.559	
average	-0.909	-1.662	1.559	**0.603**

B-S, ToT, Govt, m3

	ADF stat.	mean adj.	var. adj.	z -stat.
cz	-3.550	-2.571	1.028	
hu	-3.705	-2.571	1.028	
pl	-4.002	-2.571	1.028	
sk	-4.128	-2.571	1.028	
average	-3.846	-2.571	1.028	**-1.258**

B-S, ToT, pcGDP

	ADF stat.	mean adj.	var. adj.	z -stat.
cz	-5.411	-2.156	1.286	
hu	-3.245	-2.156	1.286	
pl	-1.807	-2.156	1.286	
sk	-3.716	-2.156	1.286	
average	-3.545	-2.156	1.286	**-1.224**

B-S, ToT, Govt, pcGDP, m3

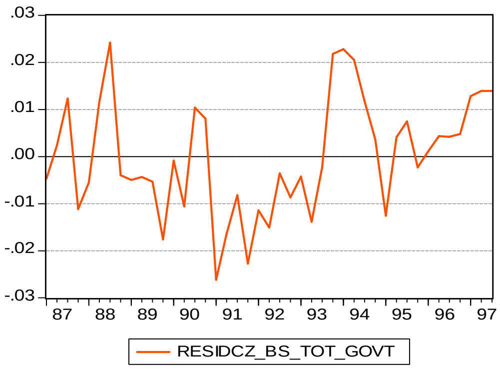
	ADF stat.	mean adj.	var. adj.	z -stat.
cz	-5.215	-3.241	0.800	
hu	-3.219	-3.241	0.800	
pl	-3.801	-3.241	0.800	
sk	-4.410	-3.241	0.800	
average	-4.161	-3.241	0.800	**-1.029**

Appendix 11

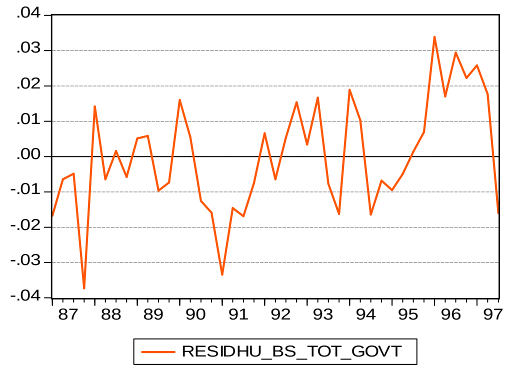
Residual Plots of Long-run Equation:

$$\varepsilon_t = \log(REER)_t - 2.03 - 0.41BS_t - 0.91\log\left(\frac{tot}{tot^*}\right)_t - 0.24\log\left(\frac{govt}{govt^*}\right)_t$$

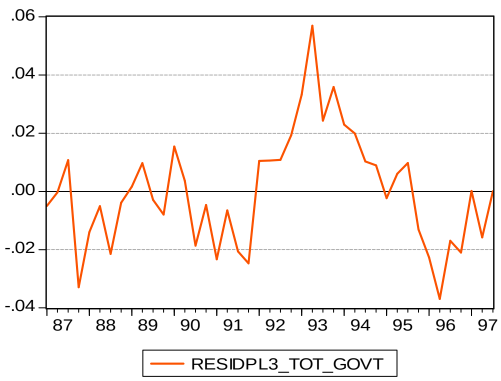
Czech Republic



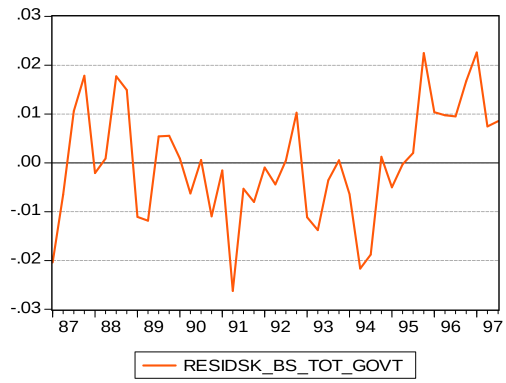
Hungary



Poland



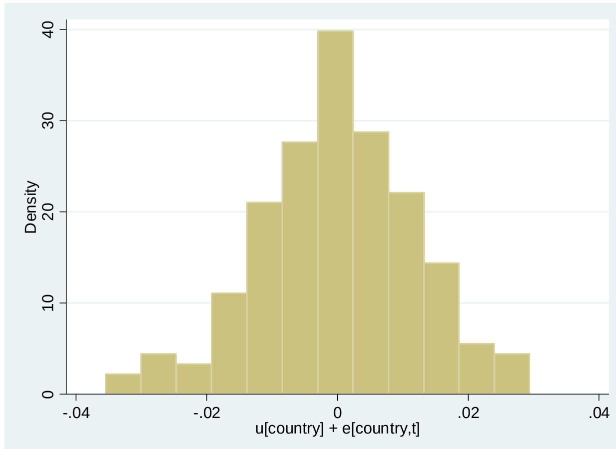
Slovakia



Appendix 12

ECM - Diagnostic Tests

Normality:



Skewness/Kurtosis test for Normality			Joint Test	
Variable	Pr(Skewness)	Pr(Kurtosis)	adj. chi2(2)	Prob>chi2
ecmresid	0.277	0.392	1.940	0.379

Serial Correlation:

resid	**Serial (1)**		**Serial (2)**		**Serial (4)**		**Serial (8)**		**Serial (12)**	
Regressors	*Coeff.*	*p-value*	*Coeff.*	*p-value*	*Coeff.*	*p-value*	*Coeff.*	*p-value*	*Coeff.*	*p-value*
c	0.000	0.964	0.000	0.963	0.000	0.810	0.000	0.610	0.001	0.376
resid (-1)	-0.109	0.167	-0.105	0.188	-0.096	0.234	-0.088	0.283	-0.102	0.225
resid (-2)			0.065	0.405	0.059	0.455	0.020	0.801	0.024	0.775
resid (-3)					-0.056	0.482	-0.046	0.562	-0.017	0.835
resid (-4)					0.048	0.537	0.037	0.643	0.070	0.395
resid (-5)							-0.005	0.954	-0.016	0.843
resid (-6)							0.055	0.477	0.065	0.407
resid (-7)							0.054	0.476	0.035	0.652
resid (-8)							0.043	0.574	0.022	0.772
resid (-9)									0.048	0.531
resid (-10)									0.081	0.294
resid (-11)									-0.015	0.840
resid (-12)									-0.109	0.153
F-test	1.930	0.167	1.320	0.270	0.890	0.474	0.470	0.872	0.580	0.858

Heteroscedasticity:

In order to avert any potential problems related to heteroscedastic errors, heterscedastic robust standard errors were used in estimation.

Appendix 13

Alternative Regression Models

log(*REER1*) average weights	**BS1**		**BS2**		**BS3**		**BS4**	
	average weights				**specific weights**			
	excl. α		**incl. α**		**excl. α**		**incl. α**	
Regressors	*Coeff.*	*s.e.*	*Coeff.*	*s.e.*	*Coeff.*	*s.e.*	*Coeff.*	*s.e.*
log(*BS*)	0.575	0.133	0.587	0.257	0.596	0.118	0.728	0.219
log(*tot*/*tot**)	1.360	0.349	1.567	0.358	1.294	0.344	1.509	0.353
log(*govt*/*govt**	0.342	0.225	0.289	0.235	0.396	0.222	0.359	0.233
constant	2.066	0.007	2.075	0.007	2.051	0.009	2.061	0.009
R^2 (within)	0.199		0.135		0.226		0.164	

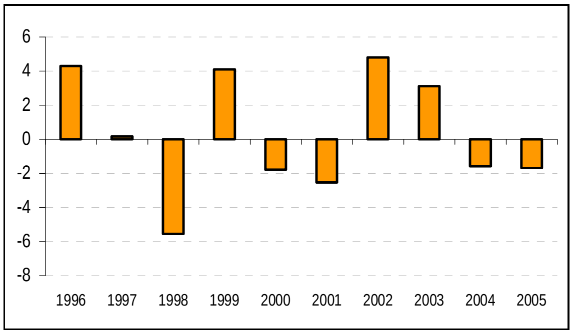
log(*REER2*) specific weights	**BS1**		**BS2**		**BS3**		**BS4**	
	average weights				**specific weights**			
	excl. α		**incl. α**		**excl. α**		**incl. α**	
Regressors	*Coeff.*	*s.e.*	*Coeff.*	*s.e.*	*Coeff.*	*s.e.*	*Coeff.*	*s.e.*
log(*BS*)	0.319	0.069	0.336	0.134	0.328	0.061	0.406	0.114
log(*tot*/*tot**)	0.824	0.181	0.938	0.186	0.789	0.178	0.907	0.184
log(*govt*/*govt**	0.234	0.116	0.207	0.122	0.263	0.115	0.244	0.121
constant	2.034	0.004	2.039	0.004	2.026	0.004	2.031	0.005
R^2 (within)	0.239		0.172		0.267		0.201	

The use of country specific weights in constructing the *REER* implies that the *REER* in each Central European country is constructed using the respective country-specific trade weights with respect to the Eurozone countries. In the context of the *BS* variable, the use of country-specific weights entails constructing different Eurozone proxies for the traded and non-traded productivity indices for each of the Central European countries. In other words, the Eurozone productivity growth gap proxy used to scale the productivity growth gap in Poland will differ from the one used for the Czech Republic, given that their trade weights with the respective Eurozone countries are different. Thus, for example, given that the Czech Republic has a larger trade weight with Germany than Poland (Cz=73%, Pl=61%), strong non-tradable inflation in Germany associated with a large productivity growth gap will have a larger impact on the *REER* in the Czech Republic and, hence, should also imply that its *effective* productivity growth gap differential with the Eurozone is smaller, ceteris paribus.

Appendix 14

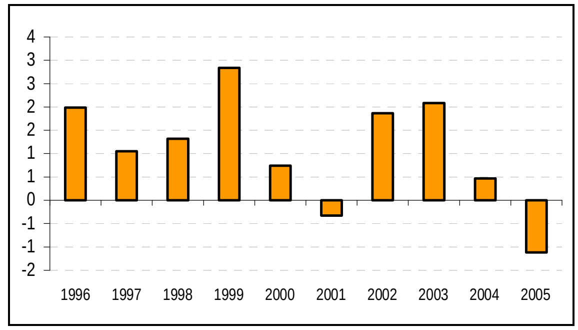
The Balassa-Samuelson effect over time

Chart 38: Czech Rep.: B-S effect on REER (% pts.)



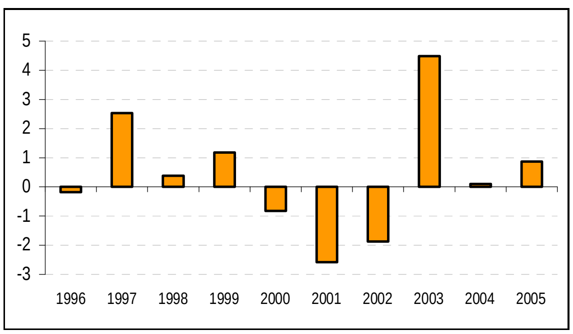
Source: National Statistics Offices, IMF IFS, Eurostat, Datastream, IMF DOTS

Chart 39 : Poland: B-S effect on REER (% pts.)



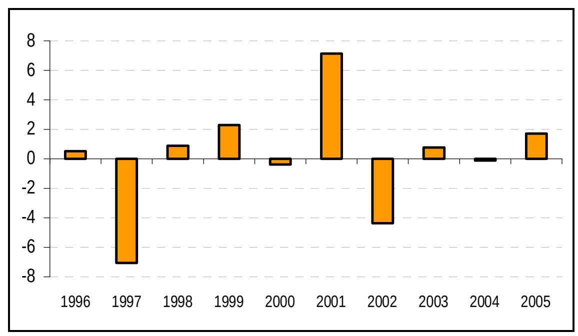
Source: National Statistics Offices, IMF IFS, Eurostat, Datastream, IMF DOTS

Chart 40: Hungary: B-S effect on REER (% pts.)



Source: National Statistics Offices, IMF IFS, Eurostat, Datastream, IMF DOTS

Chart 41: Slovak Rep.: B-S effect on REER (% pts.)



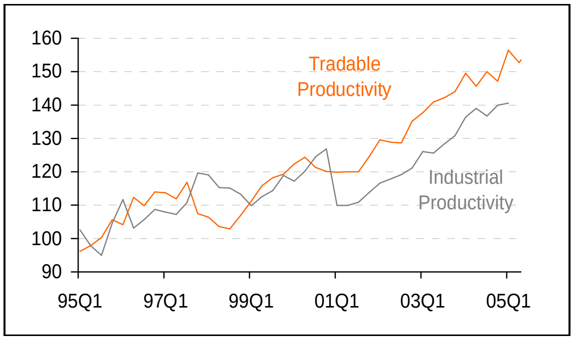
Source: National Statistics Offices, IMF IFS, Eurostat, Datastream, IMF DOTS

Notice that the proposed contributions of the B-S effect towards inflation differentials are not necessary visible in the real exchange rate path of the respective countries. In particular, for Slovakia, B-S appears to have a modest effect on the real exchange rate during 1995-2005, with exception of 1997 and 2001, where, according to coefficient estimates, B-S induced a change in the *REER* of -7.1% and 7%, respectively. During 1997 Slovakia did indeed experience a highly negative sectoral productivity growth gap, which however seems to have been largely offset by the strong nominal currency appreciation at the time (13.5%). Likewise, in the case of the proposed strong positive impact of B-S in 2001, the nominal depreciation of 3.8% and real appreciation of 3% are able to account for this.

Appendix 15

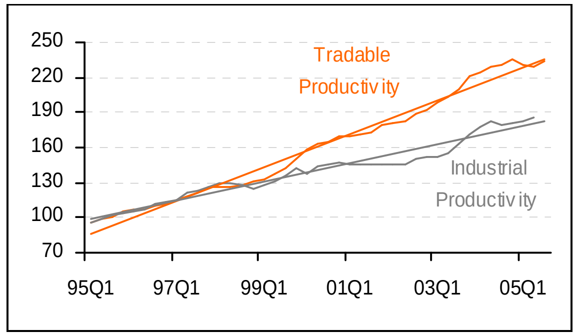
Industry vs. Tradable Productivity

Chart 42: Czech Rep.: Industrial vs. Tradable Productivity



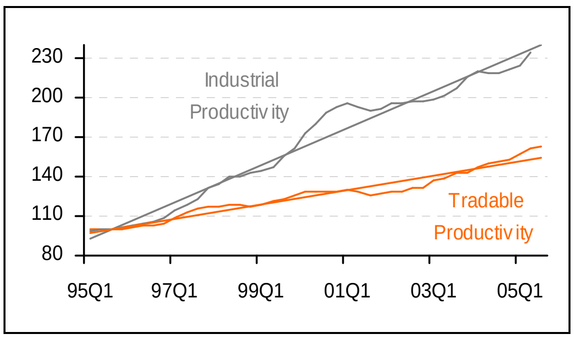
Source: Czech Statistics Office

Chart 43 : Poland: Industrial vs. Tradable Productivity



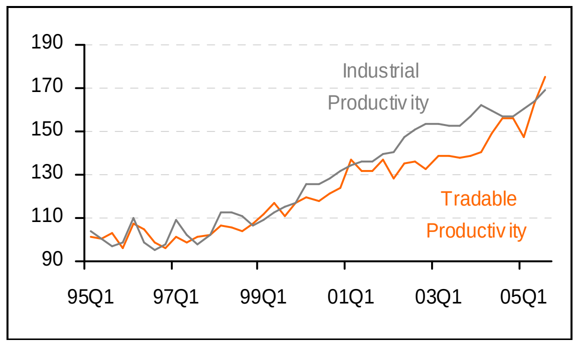
Source: Polish Statistics Office

Chart 44: Hungary: Industrial vs. Tradable Productivity



Source: Hungary Statistics Office

Chart 45: Slovak Rep.: Industrial vs. Tradable Productivity



Source: Slovak Statistics Office

Appendix 16

Seasonal adjustment procedure

Census 1 is a procedure to decompose a time series in its seasonal, trend-cycle and irregular component. The procedure assumes for the population of the variable a simple multiplicative or additive form:
$Y_t = S_t \cdot TC_t \cdot Z_t$ or $Y_t = S_t + TC_t + Z_t$
where Y_t is the observable variable, S_t the seasonal component, TC_t the trend trend-cycle component and Z_t the irregular component.

Decomposition Procedures

1. **Moving averages:** computes for each point in time a centered moving average with length of one season.
   $y_{ma,t} = \frac{1}{p} \cdot \left(\sum_{i=t-(p-1)/2}^{t+(p-1)/2} y_t \right)$ for uneven length of season periods p
   $y_{ma,t} = \frac{1}{p} \cdot \left(\sum_{i=t-(p/2)-1}^{t+(p/2)-1} y_t + \frac{y_{t+p/2} + y_{t-p/2}}{2} \right)$ for even length of season periods

2. **Ratios/Differences:** subtracts or divides actual observation and centered moving average:
   $r_t = y_t / y_{ma,t}$ for multiplicative decomposition
   $d_t = y_t - y_{ma,t}$ for additive decomposition

3. **Seasonal factors:** average ratios and differences for each point of the season
   $sm_T = \left(\prod_{i=0}^{n-1} r_{T+i\cdot p} \right)^{1/n}$ or $sa_T = \frac{1}{n} \cdot \sum_{i=0}^{n-1} d_{T+i\cdot p}$ $\forall T \in \{1,..., p\}$
   where n is the number of observed seasons, T an individual season point (say January) and zero the index for the first point in time for which a ratio could be computed. For the multiplicative model one uses typically the medial geometric average (excluding smallest and largest ratio).

4. **Seasonally adjusted series:** Adds or multiplies the original series with the seasonal factors
   $y_{sa,t} = y_t \cdot sm_T$ for multiplicative decomposition
   $y_{sa,t} = y_t + sa_T$ for additive decomposition

5. **Trend-cycle component:** 5-point centered moving average on the sa series:
   $y_{tc,t} = (1/9) \cdot y_{sa,t-2} + (2/9) \cdot y_{sa,t-1} + (3/9) \cdot y_{sa,t} + (2/9) \cdot y_{sa,t+1} + (1/9) \cdot y_{sa,t+2}$

6. **Irregular component:** Subtracting or dividing trend-cycle component from actual value
   $\varepsilon_t = y_1 / y_{tc,t}$
   $\varepsilon_t = y_t - y_{tc,t}$

NB: this section was directly sourced from **www.statsoft.com/textbook/sttimser**.

Bibliography

Asea, P., Mendoza (1994). "The Balassa Samuelson model: a general equilibrium appraisal", *Review of International Economics*, 2, 244-67

Balassa, B. (1964). "The Purchasing Power Parity Doctrine: A Reappraisal", *Journal of Political Economy*, 72

Banerjee, Dolado, Galbraith, Hendry (1993). "Co-integration, Error-Correction, and the Econometric Analysis of Non-Stationary Data", Oxford University Press: Oxford.

Bergstrand, J.H. (1991). "Structural Determinants of Real Exchange Rates and National Price Levels: Some Empirical Findings", *The American Economic Review*, Vol. LXXXI, Vol. 81, No. 1

Bhagwati, J. (1984). "Why are Services Cheaper in the Poor Countries?", *The Economic Journal*, Vol. 94.

Canzoneri, M.B., Cumby, R.E., Diba, B. (1999). "Relative labor productivity and the real exchange rate in the long run: Evidence for a panel of OECD countries", *Journal of International Economics*, Vol. 47, 245-266.

Canzoneri, M.B., Cumby, R.E., Diba, B. (2002). "Productivity Trends in Europe: Implications for Real Exchange Rates, Real Interest Rates, and Inflation", *Review of International Economics*, 10(3), 497-516.

Coricelli, F., Jazbec, B. (2001). "Real Exchange Rate Dynamics in Transition Economies", *CEPR Discussion Paper*, No.2869, updated and published in 2004 in *Structural Change and Economic Dynamics*,15.

De Broeck, M., Slok, T. (2001). "Interpreting real exchange rate movements in transition countries", *IMF Working Paper*, No. 01/56.

De Gregorio, J., Giovannini, A., Wolf, H.C. (1994). "International evidence on tradables and nontradables inflation", *European Economic Review*, Vol. 38, No. 6

De Gregorio, J., Giovannini, A., Krueger, Th.H. (1994). "The Behaviour of Nontradable-Goods Prices in Europe: Evidence and Interpretation", *Review of International Economics*, Vol. 2, No. 3, 284-305

Deutsche Bundesbank (2001). "Monetary aspects of the enlargement of the EU", *Monthly Report*, 53(10)

Deutsche Bundesbank (2002). "Fundamental determinants of real exchange rate movements in the central and east European accession countries", *Monthly Report*, 54(10)

Devereux, M. B. (1997). "Real Exchange Rates and Macroeconomics: Evidence and Theory", *The Canadian Journal of Economics*, Vol. 30, No. 4a, 773-808.

Égert, B. (2002a). “Estimating the Balassa-Samuelson effect on inflation and the real exchange rate during the transition”, *Economic Systems*, 26, 1–16.

Égert, B. (2002b). “Investigating the Balassa-Samuelson hypothesis in transition: Do we understand what we see?”, *BOFIT Discussion Papers*, No.6.

Égert, B., Drine, I., Lommatzsch, K., Rault, C. (2003). “The Balassa-Samuelson Effect in Central and Eastern Europe: Myth or Reality?”, *Journal of Comparative Economics*, Vol. 31, 552-572.

Frenkel, J., Mussa, M. (1985). “Asset Markets, Exchange Rates, and the Balance of Payments”, Ch. 14 in Grossman, G., Rogoff, K., *Handbook of International Economics, Vol.2*, Amsterdam: North Holland, pp. 679-747

Green, W. H. (2003). *“Econometric Analysis Fifth Edition”*, Prentice Hall: New Jersey.

Harris, R. and Sollis, R. (2003). “*Applied Time Series Modelling and Forecasting*”, John Wiley & Sons, Ltd: West Sussex.

Im, K.S., Pesaran, M.H. and Shin, Y. (1995, 1997). Testing for unit roots in heterogeneous panels, *Journal of Econometrics*, 44, 215-228.

Jakab, Z.M., Kovács (1999). “Determinants of Real Exchange Rate Fluctuations in Hungary”, *NBER Working Paper*.

Levin, A. and Lin, C.F. (2002). Unit Root Test in Panel Data: Asymptotic and Finite Sample Properties, *Journal of Econometrics* 108, 1-24.

Marston, R. (1987). “Real exchange rates and productivity growth in the United Sates and Japan”, *NBER Working Paper*, No. 1922

Mussa, M. (1984). “ The Theory of Exchange Rate Determination” in Bilson, J., Marston, R., *Exchange Rate Theory and Practice*, *NBER Conference Report*, Chicago: Chicago University Press, pp. 13-58.

Obstfeld, M., Rogoff, K. (1996). *Foundations of International Macroeconomics*, The MIT Press: Cambridge.

Pedroni, P. (1995). Panel Cointegration: Asymptotic and Finite Sample Properties of Pooled Time Series Tests with an Application to the PPP Hypothesis, *Econometric Theory*, 20, 2004, 597–625.

Pedroni, P. (1999). Critical Values for cointegration tests in heterogeneous panels with multiple regressors, *Oxford Bulletin of Economics and Statistics*, 61, 653-678.

Samuelson, P. (1964). “Theoretical Notes on Trade Problems”, *Review of Economics and Statistics*, 46.

Szapáry, G. (2000). “Maastricht and the Choice of Exchange Rate Regime in Transition Countries During the Run-up to EMU”, *NBH Working Paper 2000/7*

Tille, C., Stoffels, N., Gorbachev, O. (2001). "To What Extent Does Productivity Drive the Dollar?", *Current Issues in Economics and Finance,* Federal Reserve Bank of New York, 7(8)

Verbeek, M. (2004). "*A Guide to Modern Econometrics*", John Wiley & Sons, Ltd: West Sussex.