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WORKING PAPER SERIES

Paper No. 09-07

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Harrod, Balassa and Samuelson (Re)Visit Eastern Europe

SVEUČILIŠTE U ZAGREBU

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Abstract

In this paper we investigate Harrod Balassa Samuelson (HBS) effect in 11 transition countries. A large number of empirical papers based on quite limited datasets has already been published on HBS in Eastern Europe. The major contribution of this paper is the fact that we estimate HBS with NACE6 quarterly national account data which enables us to divide data into tradable and nontradable sector as suggested by De Gregorio, Giovannini and Wolf (1994) without any unrealistic assumptions. Following Bergstrand (1991) together with relative productivity we also employ share of government consumption in GDP as an explanatory variable. Unlike in previous studies, results have indicated that it is possible to find univariate cointegrating vectors only in Bulgaria, Croatia and Lithuania, and panel cointegration test has indicated that it is possible to find strong evidence of cointegration in post 2000 sample. For the post 1995 period, rejection of the null hypothesis is dependent on the inclusion of government consumption as independent variable and methodology used (DOLS vs. OLS cointegration test).

Keywords
Harrod Balassa Samuelson effect, Price convergence, Transition countries, panel cointegration tests

JEL classification
F15, F21, F43

1. Introduction

Debate about Harrod-Balassa-Samuelson (hereafter HBS) hypothesis is definitely one of the major stylized facts about transition. Inquiry into HBS in Eastern Europe started with an attempt by Halpern and Wyplosz (1997) to explain peculiar movements of real exchange rates in Eastern Europe. At that time unprecedented appreciation trend of real exchange rates in transition countries needed theoretical explanation. Real exchange rates in transition countries appreciated between 7.5 percent in Slovenia and 800 percent in Latvia with Slovenia being an outlier. The dramatic appreciation was attributed to catch-up following initial undervaluation of transition countries and HBS theory (see Figure 1 and Figure 2).

The debate culminated with the question whether the strength of HBS effect in transition countries is strong enough to interfere with Maastricht criterions of European Monetary Union (EMU). A consensus with regard to the strength of the effect in relation to Maastricht rules remained unsolved. Three studies have suggested that there is interference between EMU rules and the HBS effect in certain countries (Halpern and Wyplosz 2001; De Broeck and Slok 2001; Lojschova 2003). Six papers have suggested that there is a substantial amount of evidence of cointegration between productivity and price levels, but that there is not any evidence of interference between convergence induced inflation and EMU rules (Cipriani 2001; Coricelli and Jazbec 2001; Egert 2002a; 2002b; Egert, Drine, Lommatzsch and Rault 2003; Mihaljek and Klau 2003). Two studies have not found any evidence of the HBS effect in Transition countries at all (Fischer 2002; Arratibel, Rodriguez-Palenzuela and Thimann 2002).

The goal of this paper is to readdress the question of HBS in Transition countries on a much larger and much more consistent data sample. We use Eurostat national account quarterly data (NACE6) which enables us to implement De Gregorio, Giovannini and Wolf (1994) definition of tradability of sectors and perform univariate and panel cointegration tests without any unrealistic assumptions that were used in previous studies. For example, several authors were confronted by data restrictions forcing them to rely on cross-sectional estimation for select individual years (Halpern and Wyplosz 1997; Krajnyak and Zettelmeyer 1998; Cihak and Holub 2001); others employed panel estimates with approximately 120 to 140 observations (Halpern Wyplosz 2001; Coricelli and Jazbec 2001; De Broeck and Slok 2001; Flek, Markova and Podpiera 2002; Jazbec 2002; Arratibel, Rodriguez-Palenzuela and Thimann 2002; Fischer 2002; Lojschova 2003); and time series models with higher frequency of (monthly) data (Egert 2002a; Egert 2002b; Egert et. al. 2003; Mihaljek and Klau 2003).

In an attempt to compile enough observations for testing, many authors made unrealistic assumptions about productivity growth. Using quarterly data Fischer (2002) used average labor productivity of total economy rather than relative sectoral productivity as prescribed in the theory. Using monthly data forced Egert (2002a) and (2002b) to assume the nontradable sector productivity is equal to zero. Egert et al. (2003) used interpolation of annual data for missing quarterly data and used the ratio of CPI and PPI instead of relative sector prices. Mihaljek and Klau (2002) used quarterly growth rates of residual between growth rates of annual GDP and quarterly industrial production as output of nontradable sector in several countries.

The remainder of the paper is divided in four sections. Section 2 provides a theoretical explanation of HBS effect. Section 3 discusses the data and provides an overview of the econometric methodology. Section 4 presents and discusses the results and Section 5 concludes.

2. Theory Review – Harrod-Balassa-Samuleson effect

The Harrod-Balassa-Samuelson model predicates that, under perfect capital mobility, shifts in labor productivity cause permanent changes in the real exchange rate. This model uses long-lived real productivity shocks to drive long run price differentials (Harrod 1933; Samuelson 1964; Balassa 1964).

The intuition behind the Harrod-Balassa-Samuelson hypothesis is that tradable goods tend to experience faster productivity growth than do nontradable goods. This implies that economies with more productive tradable good sectors tend to have higher price levels than less productive ones. On the other hand, nontradable goods, which tend to be service oriented, use less technological expertise. Therefore, productivity growth in the nontradable goods sector is slower than in the tradable goods sector. Thus, real productivity differences have an effect on the real exchange rate.

Consider a small country whose tradable goods' prices are pinned down by world price levels. A positive shock to the tradable goods sector will have no affect on domestic prices of tradable goods. However, because positive productivity shocks augment the marginal product of labor, tradable goods sector wages increase. In the absence of productivity growth in the nontradable sector, it must raise prices in order to match wage increases in the tradable sector. This causes the relative price of nontradable goods to rise, increasing the aggregate price level.

Throughout seventies and eighties empirical papers on the HBS effect had mostly been based on quite simple linear relationships focusing exclusively on supply side, mostly testing the relationship between the aggregate productivity level and price level in cross-section studies (Balassa 1964). The basic model was built on the relationship between the ratio of purchasing power index PPP and nominal exchange rate E as a function of income per capita YN :

$$\frac{PPP}{E} = f(\frac{Y}{N}) \tag{1}$$

Asea and Mendoza (1994) basically made the HBS model with aggregate price levels and productivity obsolete. They incorporated the HBS theory within a long-run balanced growth neoclassical implication of a general-equilibrium model with fully modeled utility functions and the demand side of the economy. The model suggested that relative prices are a function of the relative productivity and the marginal rate of substitution between the tradable and nontradable sectors. Inclusion of substitution implied that the relative sectoral productivity and not the aggregate level of productivity determines the relative price of nontradables.

As a result of their findings, the majority of later papers used relative productivity and prices between tradable and nontradable sectors. The two sector model is built within a conceptual framework based on a standard production function with three factors of production: capital K, labor L and technology A; two types of domestically produced goods: tradable T and non-tradable N; and two Cobb-Douglas production functions, one for each sector of an economy. Formally, consider the following production functions for the tradable and nontradable goods sector respectively (time subscripts are repressed for clarity):

$$Y_T = A_T K_T^{\alpha} L_T^{1-\alpha} \tag{2}$$

and

$$Y_N = A_N K_N^{\beta} L_N^{1-\beta}. \tag{3}$$

Under the assumption of perfect competition, perfect international capital mobility, with the real interest rate pinned down by the world interest rate, perfect mobility of factors between sectors within the economy and the law of one price in the tradable sector, it is trivial to prove that a change in relative price in the non-tradable sector is a function of a change in the relative productivity of sectors and/or relative factor intensities of sectors(Rogoff 1992):

$$\hat{p} = \frac{\beta}{\alpha}\hat{a}_T - \hat{a}_N \tag{4}$$

where $p = p_N/p_T$, lower case letters denote natural logs, and `hats' denote the time derivative of the variables, e.g. $\hat{a}_T = d\ln a_T/dt$.[1]

Another important contribution to the model was made by Bergstrand (1991). He has integrated demand side of the economy into research with a goal to explore the effects of the demand side (government spending) on relative prices. The logic behind the model was the assumption that government spending preferences are biased into direction of nontradables, which should result with connection between the share of government spending in GDP and relative prices.

Therefore, an econometric estimate of equation ((4)) is based on following equation:

[1] Interesting implication of the two sector model is the fact that even if we assume that productivity shocks in each sectors are of the same magnitude, $\hat{p} > 0.$, relative price of nontradable sector will increase if the traded sector is more capital intensive $\beta/\alpha > 1$ (Froot and Rogoff 1994).

$$p_t = f(a_t, x_t) \qquad (5)$$

where relative price of nontradable sector is function of relative productivity of tradable sector, $a = (a_T/a_N)$. The vector **x** represents share of government in GDP G/Y and other potentially relevant variables, e.g. GDP per capita, ToT, Oil, etc. Such tests are usually considered to represent an internal transmission mechanism of HBS effect[2] or -- if we assume that service intensive sector is equal to nontradable sector - the Baumol-Bowen effect (Froot and Rogof 1994). In order to estimate external transmission mechanism of the HBS effect we express all variables vis-á-vis a numeraire country (Froot and Rogof 1994). In this case, the equation takes the following form:

$$\frac{p_{i,t}}{p_{0,t}} = f(\frac{a_{i,t}}{a_{0,t}}, \frac{x_{i,t}}{x_{0,t}}) \qquad (6)$$

where the subscript `0' represents the numeraire country. Thus, the ratio of the relative prices in country i to the numeraire country is a function of the relative productivity (share of government etc.) between i and the numeraire country.

3. Data

Theoretical papers on the HBS are based on the precise division of commodities into tradables and nontradables. But, in reality only few real world commodities fall easily into the nontradable category. Starting with Officer's (1976) original paper, most researchers simply assumed that manufacturing and/or industry are tradable sectors while the services sector is a nontradable part of an economy. Large number of researchers have added agriculture to tradables, and almost the same number of them have simply excluded it due to administered prices. Infrastructure, such as energy, and water management in early papers were considered tradables, while starting from the early nineties they were generally excluded from analysis.

Hitherto, the tradability of sectors has been tested only once. De Gregorio, Giovannini and Wolf (1994, pp. 1230-1232) empirically tested the tradability of various sectors of an economy. Their empirical work was based on an OECD international sector database, comprising 14 countries and 20 sectors between 1970 and 1985. De Gregorio, Giovannini and Wolf (1994) used 10 percent ratio of exports to total production of sectors in order to estimate "tradedness". Although selected subjectively, 10 percent threshold provided high level of robustness. Cutting the threshold to 5 percent would have no effect, raising it to 20 percent would shift the quantitatively small non-metal mineral products from tradables to nontradables.

According to their test, agriculture, mining and most of manufacturing[3] had a share of exports in total production of between 23.6 and 59.9 percent, agriculture having the lowest and metal manufacturing the highest shares. On the other hand, the share of exports of services was lower than 5 percent. Transport had share of 27.8 percent, while other services had a share of exports in total production of 1.9 percent. Therefore, agriculture and mining were classified as tradables, as well as manufacturing and transportation. The remaining services, accounting for about 50-60 percent of GDP, were treated as nontradables.

De Gregorio, Giovannini and Wolf (1994) division of the economy did not become a standard for future research. In following papers, a sector division of the economy remained as heterogeneous as it was in preceding papers. The most important reason for divergence in the latter papers was not so much a theoretical or empirical disagreement as much as a shortage of sufficient number of sectoral observations.

We will follow De Gregorio, Giovannini and Wolf (1994) approach to sector division with several exceptions. In order to increase the number of quarterly observations, instead of the NACE 17 dataset we will use NACE 6 sectors. Although quarterly NACE 17 database has all the sectors necessary to perform De Gregorio et al's division of sectors according to "tradedness", data for much smaller number of Eastern European countries is available. On the other hand, NACE 6 database have much larger number of observations, but publishes data for transport aggregated together with other nontradable services which makes it impossible to treat transport as tradable sector. Therefore, due to the nature of NACE 6 data,

[2] Egert (2002) divides the effect on internal and external transmission mechanism

[3] The only exception within these three sectors was the manufacturing of non-metallic minerals with a share of 13.7 percent.

transport is treated as nontradable. Another departure from De Gregorio, Giovannini and Wolf (1994) is exclusion of agriculture due to administrative prices and large subsidies in the sector, which might affect the results of our test.

We use average labor productivity as a ratio of sector's gross value added (chain-link index at basic 2000 prices) and total employment. Average productivity of industrial sector is used as a proxy for productivity in the tradable sector, and aggregate productivity of four nontradable sectors is used as an proxy for productivity of nontradable sector.[4] Price index in industry is used as price level in tradable sector and weighted average of price indices of four nontradable sectors is used as price level for nontradables.[5] Share of government represents the ratio of final government expenditure of general government and GDP (chain-link index at basic 2000 prices) and numeraire country for the estimation of external transmission mechanism is Germany.

The quarterly database has been compiled for Germany (numeraire country) and 11 transition countries: Bulgaria, Croatia, Czech Rep., Estonia, Hungary, Latvia, Lithuania, Poland, Romania, Slovakia and Slovenia. Data for Germany is available from 1991:I, Estonia from 1993:I, Bulgaria, Latvia, Lithuania and Slovakia from 1995:I, Czech Rep. from 1996:I, Croatia from 1998:II, Hungary and Slovenia from 2000:I. Due to small number of observations Poland (2007:I-2007:IV) and Romania (2002:I-2006:IV) were excluded from further analysis.

4. Methodology

Our regression analysis focuses on various versions of the following model suggested by equation ((6))

$$p_t = \alpha_i + \beta a_t + x'\gamma + e_{i,t}. \tag{7}$$

where x' is a vector of other variables; $e_{i,t} : iid(0,\sigma^2)$; and β and γ are coefficients to be estimated.

The HBS theory requires that the above model be stationary. Given the small sample period, it is difficult to make conclusions using univariate cointegrating methods (such as the Engle-Granger (1987) or Jöhansson (1991) tests) which suffer from power problems in small time series. Therefore, we employ a number of panel cointegration models to test for the existence of a long run relationship equilibrium.

4.1. Univariate Cointegration Tests

We begin by testing the HBS model using both the intra- and inter-national versions of the theory, equations ((5)) and ((6)) respectively. We use use the Jöhansson VECM methodology to conduct cointegration tests given as

$$\Delta X_t = \Gamma(L)\Delta X_t + \Pi X_{t-1} + \mu + \delta t + \varepsilon_t \tag{8}$$

where $\Gamma(L)$ is a lag operator. The long-run multiplier matrix $\Pi = \Phi(1) - I$ can be decomposed into two ($p \times r$) matrices such that $\alpha\beta' = \Pi$; β represents the cointegrating vectors or the long-run equilibria of the system of equations; and α is the matrix of error-correction coefficients which measure the rate each variable adjusts to the long-run equilibrium. Maximum likelihood estimation of ((8)) can be carried out by applying reduced rank regression. $\hat{\beta}$ is given by the r-largest eigenvectors associated with the eigenvalues $\hat{\lambda}$: Hypothesis tests on $\hat{\beta}$ can be conducted using likelihood ratio (LR) tests with standard χ^2 inference.

[4] Aggregate gross added value of four nontradable sectors is divided by aggregate employment in four nontradable sectors according to equation: $Y_N / L_N = \sum_{S=1}^{4} Y_{N,S} / \sum_{S=1}^{4} L_{N,S}$, where S denotes four sectors.

[5] Sector shares of gross added value of the nontradable sectors are used as weights according to equation: $P_N = \sum_{S=1}^{4}((Y_{N,S} / Y_N) P_{N,S})$, where P denotes price index.

The LR test statistic is given by, $T\sum_{i=1}^{r}\ln[(1-\tilde{\lambda}_i)/(1-\hat{\lambda}_i)] : \chi^2_{r(p-s)}$ where $\tilde{\lambda}_i$ are the eigenvalues from the restricted MLE.

The test for cointegration is a test for the number of non-zero eigenvalues. The likelihood ratio statistic testing the rank of Π, or equivalently the number of non-zero eigenvalues, is the trace statistic is given by $-T\sum_{i=1+r}^{p}\ln(1-\hat{\lambda}_i)$, (Jöhansen, 1988, 1991).

4.2. Panel Cointegration Tests

Given the highly technical nature of panel cointegration, and the extensive literature, we briefly outline the panel cointegration methods used. The synopsis implies the reader is familiar with the cointegration basics.[6] Consider the following fixed-effect panel regression:

$$y_{i,t} = \alpha_i + x_{i,t}^{'}\beta + e_{i,t}, i = 1,\ldots,N, t = 1,\ldots,T, \tag{9}$$

where $y_{i,t}$ are 1 × 1, β is an $M \times 1$ vector of slope parameters, α_i are the intercepts, and $e_{i,t} : I(0)$ are the stationary disturbance terms. It is assumed that $x_{i,t}$ are $M \times 1$ I(1) processes which are themselves not cointegrated such that each variable is a random walk process

$$x_{i,t} = x_{i,t-1} + \varepsilon_{i,t}.$$

These assumptions imply that ((9)) represents a system of cointegrating regressions.

Recently, there has been a growth in panel estimation techniques based on the Engle and Granger (1987) and Jöhansen (1991) univariate methods of testing for cointegration in systems to take advantage of the power gains of increasing the number of observations in panel data. In this section we briefly review the OLS and DOLS estimators for panel cointegration (see Kao and Chiang, 2000).[7] Kao and Chiang (2000) derive the limiting distributions for the OLS, FMOLS and DOLS estimators for the regression specification given in ((9)). They also investigated the finite sample properties of each estimator through Monte Carlo simulation. They found that i. the OLS estimator has a non-negligible bias; ii. the FMOLS estimator does not improve on the OLS estimator in general; and iii. the DOLS estimator has the best properties of the three. Given the small gains to using FMOLS we concentrate on the OLS and DOLS estimators.

The OLS estimator of β is,

$$\hat{\beta}_{OLS} = \left[\sum_{i=1}^{N}\sum_{t=1}^{T}(x_{it}-\overline{x}_{it})(x_{it}-\overline{x}_{it})'\right]^{-1}\left[\sum_{i=1}^{N}\sum_{t=1}^{T}(x_{it}-\overline{x}_{it})(y_{it}-\overline{y}_{it})\right] \tag{10}$$

where a bar denotes the variables time mean. Given the large bias in the standard OLS regression, we also consider the bias adjusted OLS estimator, $\tilde{\beta}$:

$$\tilde{\beta}_{OLS} = \hat{\beta}_{OLS} - \bar{\mathrm{u}}_{OLS}$$

where $\bar{\mathrm{u}}_{OLS}$ is the mean, over time, OLS bias with

$$\hat{\mathrm{u}}_{OLS} = -3\hat{\Omega}_1^{-1}\mu'_{\varepsilon_1} + 6\hat{\Omega}_1^{-1}\Delta'_{\mu\varepsilon_1}$$

where Ω_1 is the estimated long run covariance matrix, μ_{ε_1} is the first row of the estimated mean error and $\Delta_{\mu\varepsilon_1}$ is the kernel estimates of the long run covariance matrices.

The DOLS estimator is given as,

$$y_{it} = \alpha_i + x_{it}^{'}\beta + \sum_{j=-q}^{q} c_{ij}\Delta x_{it-j} + \upsilon_{it}. \tag{11}$$

[6] For an overview of panel cointegration methods see Banerjee (1999) and Kao (1999). See Kao and Chiang (2000), Phillips and Moon (1999) and Pedroni (1995, 1999) for a more detailed analysis of the panel cointegration estimators. The discussion here is based on Banerjee (1999), Pedroni (1999) and Kao and Chiang (2000).

[7] Much of the discussion in the present paper is taken from Banerjee (1999) and Kao et al. (1999).

Given the relatively short term data we only allow for up to 1 lead and lags.

Kao (1999) describes two types of panel cointegration tests based on the univariate Engel-Granger (EG) cointegration tests. A Dickey-Fuller (DF) type test and the augmented Dickey-Fuller (ADF) test. As in the EG test, we check for panel cointegration by conducting unit root tests using the residuals from the panel cointegration estimators:

$$\hat{e}_{i,t} = \gamma_i \hat{e}_{i,t-1} + \nu_{i,t} \tag{12}$$

where the $\hat{e}_{i,t}$ are the estimated residuals from equation ((9)). The null hypothesis of *no* cointegration is, as in the EG test,

$$H_0 : \gamma_i = 1, H_A : \gamma_i = \gamma < 1, \forall\, i \tag{13}$$

notice that in the specification we restrict the estimated AR coefficients to be equal across all i regressions. The first three tests we consider are based on the DF and ADF tests used in the EG test for cointegration. The two DF based tests are:

$$DF_\rho = \frac{\sqrt{N}T(\hat{\rho}-1)+3\sqrt{3}}{\sqrt{10.2}} \tag{14}$$

and

$$DF_\rho^* = \frac{\sqrt{N}T(\hat{\rho}-1)+\dfrac{3\sqrt{N}\hat{\sigma}_\upsilon^2}{\hat{\sigma}_{0\upsilon}^2}}{\sqrt{3+\dfrac{7.2\hat{\sigma}_\upsilon^4}{\hat{\sigma}_{0\upsilon}^4}}} \tag{15}$$

where $\hat{\sigma}_\upsilon^2 = \Sigma_u - \Sigma_{u\varepsilon}\Sigma_\varepsilon^{-1}$ and $\hat{\sigma}_{0\upsilon}^2 = \Omega_u - \Omega_{u\varepsilon}\Omega_\varepsilon^{-1}$.

We also account for heterogeneity in the regressors by conducting the panel tests suggested by Pedroni (1999). These tests fall into two categories. Define γ_i to be first order AR coefficient of from the residuals of the ith cross unit, equation ((12)). The first set of tests restricts this coefficient to be equal across all i units, as in equation ((13)), which is similar to the restriction on the Levin-Lin AR parameter. The second set of tests relax the restriction on γ along the lines of the IPS panel unit root test:

$$H_0 : \gamma_i = 1, H_A : \gamma_i < 1, \forall\, i. \tag{16}$$

In either case, the null hypothesis is no cointegration. In the interest of saving space, we outline the procedure here, interested readers are encouraged to read the original paper. The tests we employ are two from the first category, equation ((13)), of the Pedroni panel cointegration statistics, the panel ν-statistic, PC_ν

$$PC_\nu = T^2 N^{32} \left(\sum_{i=1}^{N}\sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{i,t-1}^2\right)^{-1} \tag{17}$$

and ρ-statistic, PC_ρ

$$PC_\rho = T\sqrt{N} Z_{\hat{\rho}_{N,T}} \left(\sum_{i=1}^{N}\sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{i,t-1}^2\right)^{-1} \sum_{i=1}^{N}\sum_{t=1}^{T} \hat{L}_{11i}^{-2} (\hat{e}_{i,t-1}\Delta\hat{e}_{i,t} - \hat{\lambda}_i) \tag{18}$$

and the group ρ statistic, GC_ρ, based on the less restrictive group model, equation ((16)),

$$GC_\rho = T\sqrt{N} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \hat{e}_{i,t-1}^2\right)^{-1} \sum_{t=1}^{T} (\hat{e}_{i,t-1}\Delta\hat{e}_{i,t} - \hat{\lambda}_i) \tag{19}$$

where $\hat{\lambda}_i = (1/2)(\hat{\sigma}_i^2 - \hat{s}_i^2)$ where $\hat{\sigma}_i^2$ and $\hat{s}_i^2$ are the long run and contemporaneous respectively of the estimated residuals from the autoregression $\hat{u}_{i,t} = \hat{e}_{i,t} - \hat{\gamma}_i \hat{e}_{i,t-1}$. $\hat{L}_{11i}^{-2} = \hat{\Omega}_{21,i}\hat{\Omega}_{22,i}^{-1}\hat{\Omega}_{21,i}'$ is the multivariate estimate of the long run covariance matrix, Ω_i.

5. Results

5.1. Univariate Results

The first step is to check for the presence of unit processes in the data. For each country entire available data set[8] is used in univariate unit root test. We use the standard augmented Dickey-Fuller (ADF) model to check for $I(0)$ processes:

$$\Delta y_t = \alpha + bt + \rho y_{t-1} + \sum_{j=1}^{k} \phi_j \Delta y_{t-j} + \varepsilon_t. \tag{20}$$

Given that the ADF unit root test is well known, we do not go into much detail. The null hypothesis is that the series $\{y\}_0^T : I(1)$, that is, if the AR(1) coefficient equals one, $\rho = 1$. A nonstationary null hypothesis implies we cannot use standard t test statistics as the underlying distribution is assumed to have a non-constant mean and variance. Thus, standard t-statistics cannot be used and the inference is made based on the Dicky-Fuller distribution.

Results are presented in Table 3. As the concern is only whether or not each series is stationary, the ADF t-statistics are presented. We consider two sets of series: data for the intrAnational tests are given as, for any variable p, a, g, and for the intErnational version p_i/p_0, a_i/a_0, g_i/g_0 for each country *vis á vis* the numeraire country, Germany. The left side of the table presents results for international HBS and the right side contains results for the intranational version.

Results indicate that most of the variables are nonstationary, but only small number of countries have all three nonstationary variables. In international test, all three variables for Bulgaria (p_i/p_0, a_i/a_0, g_i/g_0) and two variables (p_i/p_0, a_i/a_0)for Lithuania and Croatia are nonstationary. In intranational test, Slovakia is the only country with all three nonstationary variables (p, a, g). Following unit root test results, cointegration tests are performed on international data for Bulgaria, Croatia and Lithuania and intranational data for Slovakia. Table 4 presents cointegration test results for Bulgaria, Croatia and Lithuania. Results suggest that relative price levels *vis á vis* Germany are cointegrated with relative prices and relative share of government in Bulgaria, and with relative prices only in Lithuania. In case of Croatia, results are ambiguous due to different number of cointegrating vectors implied by trace and max-eigenvalue statistics. Results for intranational version in Slovakia implied zero cointegrating vectors.

5.2. Panel Results

In order to create balanced panel three data sets are employed. In the first panel, data for Estonia, Bulgaria, Latvia, Lithuania and Slovakia during the period 1995:I-2008:3 are used. In the second panel, Czech Rep. is added to the panel and time structure is shortened to 1996:I-2008:3. In the third panel, Croatia, Hungary and Slovenia are added and time span is 2000:1-2008:3. The first and second panel allow us to test for HBS effect during the period of recovery of transition countries in the second part of nineties. Third panel has much larger number of cross-sections, but it only focuses on the 21st century.

Figure 3 and Figure 4 show average growth rates of relative productivity and prices for countries in the first and third panel during the two different time spans. It is more than obvious that most of transition countries in both panels experienced relative productivity growth that was slower compared to Germany. In the first panel, three countries and in the third panel seven countries had negative relative productivity growth *vis á vis* Germany. On the other hand all countries experienced positive growth of relative prices *vis á vis* Germany.

We begin with unit root tests to check the order of integration of each of the variables. We use the panel test introduced by Im, Pesaran, and Shin (1997). IPS test rely on a panel representation of the standard augmented Dicky-Fuller model:

[8]Data for Germany is available from 1991:I, Estonia from 1993:I, Bulgaria, Latvia, Lithuania and Slovakia from 1995:I, Czech Rep. from 1996:I, Croatia from 1998:II, Hungary and Slovenia from 2000:I. Due to small number of observations Poland (2007:I-2007:IV) and Romania (2002:I-2006:IV) were excluded from further analysis.

$$\Delta y_{it} = \alpha_i + \rho_i y_{it-1} + \sum_{j=1}^{k} \phi_j \Delta y_{it-j} + \varepsilon_{it}, i = 1,\ldots,N, t = 1,\ldots,T \tag{21}$$

As in the univariate case the null hypothesis is $\rho_i = 0$. Note, in this representation we allow for heterogeneous intercepts and AR(1) parameters. We are concerned with the Studentized t-statistics on the estimated ρ_i for the panel as a whole.

The IPS test's panel t-statistic is the mean of the individual univariate test statistics derived from the panel standard errors. Other panel tests include the Harris and Tzavalis (1999) and Maddala and Wu (1999) Fisher statistics. As the results are comparable, we do not report them here.

Results of the panel unit root tests and cointegration tests are presented in Table 5, Table 6 and Table 7. Part I of each table includes the unit root tests and Part II and III present the results of the cointegration tests. We consider both variations of the data, the intranational, based on equation ((4)), and international HBS (Germany as numeraire country) presented in equation ((5)). For each variation of the data, two HBS models are estimated: `Model 1' is the canonical HBS regression of price dispersion and productivity. In `Model 2' we include the government expenditure-GDP ratio for each country.

The panel unit root test statistics presented are the p-values of the null hypothesis that the variable is $I(1)$. The tests demonstrate that many of the variables in the first panel are indeed nonstationary at standard test critical values (Table 5). In the second panel, only relative productivity a_i and international relative productivity a_i/a_0 are nonstationary, while international relative government share g_i/g_0, government share g_i and relative prices p_i are stationary at 10% and international relative prices p_i/p_0 at 5%. In the third panel, only international government share g_i/g_0 is stationary at 10% significance level, while all other variables are nonstationary at conventional levels.

We report the *bias adjusted* OLS results in Part II and the DOLS estimates in the Part III of Table 5, Table 6 and Table 7. Beginning first with the estimated coefficients for productivity and government share, it is straightforward to see that the estimates are statistically significant in all panels.

Productivity coefficients are between 0.23 (DOLS estimate in intranational model 1 of the second panel (Table 6)) and 1.00 (OLS estimate of international model 1 of the third panel (Table 5)). Estimates for government consumption are between -0.40 (DOLS estimate of international model in the second panel (Table 6)) and 0.15 (DOLS estimate of intranational model in the third panel (Table 7)).

It is interesting to notice that compared to DOLS test, OLS have resulted with smaller estimates for productivity coefficient in the first two panels (longer time span, smaller number of countries) and bigger estimates in the third panel. Also, OLS estimated coefficient for government consumption is negative in first two panels and positive in the third, while DOLS estimate indicates that government coefficient is negative internationally and positive otherwise in all panels.

It is also clear that the third panel cointegration tests strongly reject the null hypothesis of no cointegration in all possible variations (Table 7). In the first panel, the null hypothesis of no cointegration is strongly rejected for international and intranational Model 2 (Table 5). In the second model, null hypothesis is rejected in the international model 2 and in DOLS estimate for intranational model 1 and 2(Table 6).

6. Summary

Unlike previous studies on HBS in transition countries we have performed univariate and panel cointegration test on much larger number of observation[9] and without unrealistic assumptions[10]. Also, Eurostat NACE 6 classification data on prices, added value and employment enabled us to divide data into tradable and nontradable sector according to De Gregorio, Giovannini and Wolf (1994) methodology.

Results have indicated that it is possible to find univariate cointegrating vectors only in Bulgaria, Croatia and Lithuania, and panel cointegration test has indicated that it is possible to find strong evidence

[9] Halpern Wyplosz (2001); Coricelli and Jazbec (2001); De Broeck and Slok (2001); Flek, Markova and Podpiera (2002); Jazbec (2002); Arratibel, Rodriguez-Palenzuela and Thimann (2002); Fischer (2002); Lojschova (2003)
[10] Fischer (2002); Egert (2002a, 2002b); Egert et al. (2003); Mihaljek and Klau (2002)

of cointegration in the panel of 9 countries during 2000:I-2008:III period. Two panels covering late nineties (5-6 countries), have resulted with ambiguous results, where rejection of the null hypothesis were depended on inclusion of government consumption into specification (First panel) or methodology used (Second panel).

References

1. Arratibel, Olga; Diego Rodríguez-Palenzuela, and Christian Thimann. 2002. "Inflation Dynamics and Dual Inflation in Accession Countries: A 'New Keynesian' Perspective" ECB working paper, No. 132.
2. Asea, Patrick and Enrique Mendoza. 1994. "The Balassa-Samuelson Model: A General Equilibrium Appraisal" Review of International Economics, 2:3, pp. 244-267.
3. Balassa, Bela (1964). "Purchasing Power Parity: A Reappraisal," *Journal of Political Economy*, 72, 584-96.
4. Balassa, Bela. 1964. "The Purchasing Power Parity Doctrine: A Reappraisal" Journal of Political Economy, 72:6, pp. 584-596.
5. Banerjee, Anindya (1999). "Panel Data Unit Roots and Cointegration: An Overview," *Oxford Bulletin of Economics and Statistics* 61, 607-29.
6. Bergstrand, Jeffrey. 1991. "Structural Determinants of Real Exchange Rates and national Price Levels: Some Empirical Evidence," American Economic Review, 81:1, pp. 325-334.
7. Cipriani, Marco. 2001. "The Balassa-Samuelson Effect in Transition Economies", mimeo, IMF.
8. Coricelli, Fabrizio and Bostjan Jazbec. 2001. "Real Exchange Rate Dynamics in Transition Economies," CEPR discussion paper No. 2869.
9. Čihák, Martin, Tomáš Holub. 2001. "Convergence of Relative Prices and Inflation in Central and Eastern Europe" IMF working paper 124.
10. De Broeck, Mark and Torsten Slø k. 2001. "Interpreting Real Exchange Rate Movements in Transition Countries," IMF working paper 56.
11. De Gregorio, Jose; Alberto Giovannini, and Holger Wolf. 1994. "International Evidence on Tradables and Nontradables Inflation," European Economic Review, 38, pp. 1225-1224.
12. Égert, Balázs. 2002a. "Estimating the impact of the Balassa-Samuelson effect on inflation and the real exchange rate during transition," Economic Systems, 26, pp. 1-16.
13. Égert, Balázs. 2002b. "Investing the Balassa-Samuelson hypothesis in the Transition, Economics of Transition," Economics of Transition, 10:2, pp. 273-309.
14. Égert, Balázs; Imed Drine, Kirsten Lommatzsch and Christophe Rault. 2003. "The Balassa-Samuelson effect in Central and Eastern Europe: myth or reality?" Journal of Comparative Economics, 31, pp. 552-572.
15. Engle, R.F. and C.W.J. Granger (1987), "Cointegration and Error-Correction: Representation, Estimation, and Testing", Econometrica 55 (March), pp. 251-27.
16. Fischer, Christoph. 2002. "Real Currency Appreciation in accession countries: Balassa-Samuelson and Investment Demand", BOFIT Discussion Papers No. 8.
17. Flek, Vladislav; Lenka Marková and Ji ř í Podpiera. 2002. "Sector Productivity and Real Exchange Rate Appreciation: Much Ado about Nothing?" Czech National Bank Working Paper Series 4/2002.
18. Froot, Kenneth A. and Kenneth Rogoff. 1994. "Perspectives on PPP and Long-Run Real Exchange Rates". NBER Working Paper No. 4952
19. Halpern, László and Charles Wyplosz (1997) "Equilibrium Exchange Rates in Transition Economies", IMF Staff Papers 44(4), p. 430-460.
20. Halpern, László, and Charles Wyplosz. 2001. "Economic Transformation and Real Exchange Rates in the 2000s: The Balassa-Samuelson Connection," Geneve: UN/ECE.
21. Harris, Richard D. F. and E. Tzavalis, (1999), "Inference for Unit Roots in Dynamic Panels Where the Time Dimension is Fixed," Journal of Econometrics 91, pp.201-226.
22. Harrod, Roy F. 1933. International Economics, Nisbet and Cambridge University Press.
23. Im, K.S., Pesaran, M.K. and Shin, Y., 2003. "Testing for Unit Roots in Heterogenous Panels", Journal of Econometrics 115, pp.53-74.

24. Jöhansen, Soren, 1991. "Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models," Econometrica, Econometric Society, vol. 59(6), pp. 1551-80.
25. Jazbec, Bostjan. 2002. "Balassa-Samuelson Effect in Transition Economies - The Case of Slovenia" William Davidson Working Paper No. 507.
26. Kao, Chihwa and Min-Hsien Chiang (2000). "On the Estimation and Inference of a Cointegrated Regression in Panel Data," *Advances in Econometrics* 15,179-222.
27. Kao, Chihwa (1999). "Spurious Regression and Residual-Based Tests for Cointegration in Panel Data," *Journal of Econometrics* 90, 1-44.
28. Kao, Chihwa and Chiang, Min-Hsien and Chen, Bangtian (1999). "International R&D Spillovers: An Application of Estimation and Inference in Panel Cointegration,"*Oxford Bulletin of Economics and Statistics* 61, 691-709.
29. Krajnyák, Kornélia, and Jeromin Zettelmeyer. 1998. "Competitiveness in Transition Economies: What Scope for Real Appreciations," IMF Staff Papers, 45:2.
30. Levin, A, Lin, C.F., 1993. "Unit Root Tests in Panel Data: Asymptotic and Finite-sample Properties", Unpublished Manuscript, University of California, San Diego
31. Lojschova, Adriana. 2003. "Estimating the Impact of the Balassa-Samuelson Effect in Transition Economies," Vienna: Institute for Advanced Studies.
32. Maddala, G.S. and Wu, S. (1999) A comparative study of unit root tests with panel data and a new simple test. Oxford Bulletin of Economics and Statistics. 61: 631-652.
33. McCoskey, Suzanne and Kao, Chihwa (1998). "A Residual-Based Test of the Null of Cointegration in Panel Data," *Econometric Reviews* 17, 57-84.
34. Mihaljek, Dubravko, and Marc Klau. 2003. "The Balassa-Samuelson effect in Central Europe: A Disaggregated Analysis," BIS Working Paper 143.
35. Officer, Lawrence H. 1976. "The Productivity Bias in Purchasing Power Parity: An Econometric Investigation" IMF Staff Paper 23, pp. 545-579.
36. Pedroni, Peter L. (1993). *Panel Cointegration, Endogenous Growth and Business Cycles in Open Economies*, Ph.D. Thesis, Columbia University.
37. Pedroni, Peter L. (1999). "Critical Values for Cointegration Tests in Heterogeneous Panels with Multiple Regressors," *Oxford Bulletin of Economics and Statistics* 61, 653-70.
38. Phillips, Peter C. B. and Moon, Hyungsik R. (1999). "Nonstationary Panel Data Analysis: An Overview of Some Recent Developments," University of California, Santa Barbara,Working Papers in Economics: 98R/17.
39. Rogoff, Kenneth (1996). "The Purchasing Power Parity Puzzle," *Journal of Economic Literature*, XXXIV, 647-668.
40. Rogoff, Kenneth. 1992. "Traded Goods Consumption Smoothing and the Random Walk Behavior Walk Behavior of the Real Exchange Rate" NBER Working paper No. 4119.
41. Samuelson, Paul A. 1964. "Theoretical Notes on Trade Problems," Review of Economics and Statistics, 46:2, pp. 145-54.

Tables

Table 1: De Gregorio et al. classification in terms of NACE 17

T/NT	Sector
T	Agriculture, hunting and forestry
T	Fishing
T	Mining and quarrying
T	Manufacturing
NT	Electricity, gas and water supply
NT	Construction
NT	Wholesale and retail trade
NT	Hotels and restaurants
T	Transport, storage and communication
NT	Financial intermediation
NT	Real estate, renting and business activities
NT	Public administration and defence; compulsory social security
NT	Education
NT	Health and social work
NT	Other community, social, personal service activities
Excluded	Activities of households

Table 2: De Gregorio et al. classification adjusted to NACE 6

T/NT	Sector
Excluded	Agriculture, hunting, forestry and fishing
T	Total industry (excluding construction)
NT	Construction
NT	Wholesale and retail trade; hotels and restaurants; transport,
	storage and communication
NT	Financial intermediation; real estate, renting and business activities
NT	Public administration and defence; education; health and social work;
	other; private households with employed persons

Table 3: Univariate Unit Root Tests

	International				Intranational			
		Trend	Mean	none		Trend	Mean	none
EST	p_i/p_0	-1.87	-1.46	-2.25**	p_i	-2.14	-1.09	-1.70**
	a_i/a_0	-3.35*	-3.37**	--	a_i	-5.24***	--	--
	g_i/g_0	-3.75**	--	--	g_i	-3.86**	--	--
BUL	p_i/p_0	-1.85	-1.15	-1.24	p_i	-4.21***	--	--
	a_i/a_0	-1.70	-0.07	1.30	a_i	-2.59	-1.66	-1.90*
	g_i/g_0	-2.71	-2.62*	-0.97	g_i	-0.47	-1.20	0.08
LAT	p_i/p_0	-1.86	-2.33	-2.59**	p_i	-1.76	-2.24	-2.57**
	a_i/a_0	-0.87	0.43	1.36	a_i	-0.87	1.88	-0.28
	g_i/g_0	-3.83**	--	--	g_i	-4.24***	--	--
LIT	p_i/p_0	-3.38*	-2.84*	1.25	p_i	-3.73**	--	--
	a_i/a_0	-2.39	-1.61	-0.55	a_i	-3.02	-0.07	-0.95
	g_i/g_0	-2.61	-0.37	-2.00	g_i	-3.28	0.03	3.11
SLO	p_i/p_0	-5.45***	--	--	p_i	-3.52*	0.07	0.14
	a_i/a_0	-2.75*	-2.86	-0.40	a_i	-2.75*	-0.93	-1.60
	g_i/g_0	-2.95	-1.38	-0.18	g_i	-3.16*	-0.58	-1.26
SLK	p_i/p_0	-2.96	-2.19	-1.55	p_i	-3.24*	-0.85	0.76
	a_i/a_0	-5.01***	--	--	a_i	-1.92	-0.08	1.21
	g_i/g_0	-4.49***	--	--	g_i	-1.45	0.31	1.62
CZE	p_i/p_0	-1.75	0.55	-0.09	p_i	-3.61**	--	--
	a_i/a_0	-4.21***	--	--	a_i	-3.67**	--	--
	g_i/g_0	-0.77	-0.73	-1.13	g_i	-1.11	-0.31	0.93
CRO	p_i/p_0	-2.00	-2.64	-1.05	p_i	-1.55	-2.31	-0.44
	a_i/a_0	-3.23	-1.53	0.19	a_i	-4.36***	--	--
	g_i/g_0	-3.43	-1.62	-2.90***	g_i	-3.96**	--	--
HUN	p_i/p_0	-0.96	-1.58	0.76	p_i	-3.53*	-0.27	1.55
	a_i/a_0	-4.11**	--	--	a_i	-6.21***	--	--
	g_i/g_0	-1.28	-0.26	-0.99	g_i	-1.13	0.42	1.61
GER	--	--	--	--	p_i	-3.23*	-1.80	-1.76
	--	--	--	--	a_i	-3.10	0.93	1.15
	--	--	--	--	g_i	-3.08	-0.53	0.95

Notes: ADF Studentized t-statistics are presented; ***, **, and * represent rejection of the null hypothesis at the 1%, 5%, and 10% level respectively.

Table 4: Univariate Cointegration Tests

	Trace statistics	No. of coint. vectors	Max-eigen. statistics	No. of coint. vectors	Cointegrating vector p_i/p_0	a_i/a_0	g_i/g_0	constant	trend
BUL	66.71	1	53.15	1	1.00	0.32	-1.09	-0.03	
						(-0.06)	(-0.05)	(-0.02)	
LIT	29.91	1	26.26	1	1.00	0.35		0.11	-0.003
						(0.02)			(0.001)
CRO	13.51	1	10.85	0	1.00	0.16			
						(0.04)			

Notes: t statistics for coefficient estimates in brackets

Table 5: Panel Unit Root and Cointegration Tests 1995:I-2008:III

	I. Panel unit Root Tests					
	p_i/p_0	a_i/a_0	g_i/g_0	p_i	a_i	g_i
IPS	0.22262	0.33349	0.30627	0.39354	0.14812	0.26909

	International		**Intranational**	
	II. OLS Cointegration Tests			
	Model 1	*Model 2*	*Model 1*	*Model 2*
Productivity	0.4981	0.4630	0.4164	0.3869
	(0.000)	(0.000)	(0.000)	(0.000)
Government	--	-0.3178	--	-0.3202
	--	(0.000)	--	(0.000)
DF_ρ	0.000	0.000	0.000	0.000
DF_ρ^*	0.000	0.000	0.0029	0.000
PC_ν	0.000	0.000	0.000	0.000
PC_ρ	0.2154	0.000	0.4593	0.0002
GC_ρ	0.2175	0.000	0.4597	0.0002
	III. DOLS Cointegration Tests			
Productivity	0.5278	0.5890	0.2401	0.3430
	(0.000)	(0.0075)	(0.000)	(0.1188)
Government	--	-0.3862	--	0.1254
	--	(0.0029)	--	(0.2310)
DF_ρ	0.000	0.000	0.000	0.000
DF_ρ^*	0.000	0.000	0.0270	0.000
PC_ν	0.000	0.000	0.000	0.000
PC_ρ	0.1419	0.000	0.4784	0.0012
GC_ρ	0.1442	0.000	0.4786	0.0013

Notes: p-values for coefficient estimates in parenthesis,
panel statistics are p-values under the null hypothesis of *no* cointegration.

Table 6: Panel Unit Root and Cointegration Tests 1996:I-2008:III

	I. Panel unit Root Tests					
	p_i/p_0	a_i/a_0	g_i/g_0	p_i	a_i	g_i
IPS	0.02054	0.47435	0.06381	0.07033	0.27321	0.09494

	International		**Intranational**	
	II. OLS Cointegration Tests			
	Model 1	*Model 2*	*Model 1*	*Model 2*
Productivity	0.4707	0.4409	0.4042	0.3787
	(0.000)	(0.000)	(0.000)	(0.000)
Government	--	-0.2489	--	-0.2516
	--	(0.000)	--	(0.000)
DF_ρ	0.000	0.000	0.000	0.000
DF_ρ^*	0.000	0.000	0.000	0.000
PC_ν	0.000	0.000	0.000	0.000
PC_ρ	0.3968	0.0078	0.0673	0.1105
GC_ρ	0.3978	0.0084	0.0692	0.1128
	III. DOLS Cointegration Tests			
Productivity	0.5119	0.5906	0.2336	0.3403
	(0.000)	(0.000)	(0.000)	(0.0002)
Government	--	-0.3955	--	0.1333
	--	(0.000)	--	(0.0037)
DF_ρ	0.000	0.000	0.000	0.000
DF_ρ^*	0.000	0.000	0.000	0.000
PC_ν	0.000	0.000	0.000	0.000
PC_ρ	0.0983	0.000	0.0034	0.0176
GC_ρ	0.1005	0.000	0.0037	0.0185

Notes: p-values for coefficient estimates in parenthesis,
panel statistics are p-values under the null hypothesis of *no* cointegration.

Table 7: Panel Unit Root and Cointegration Tests 2000:I-2008:III

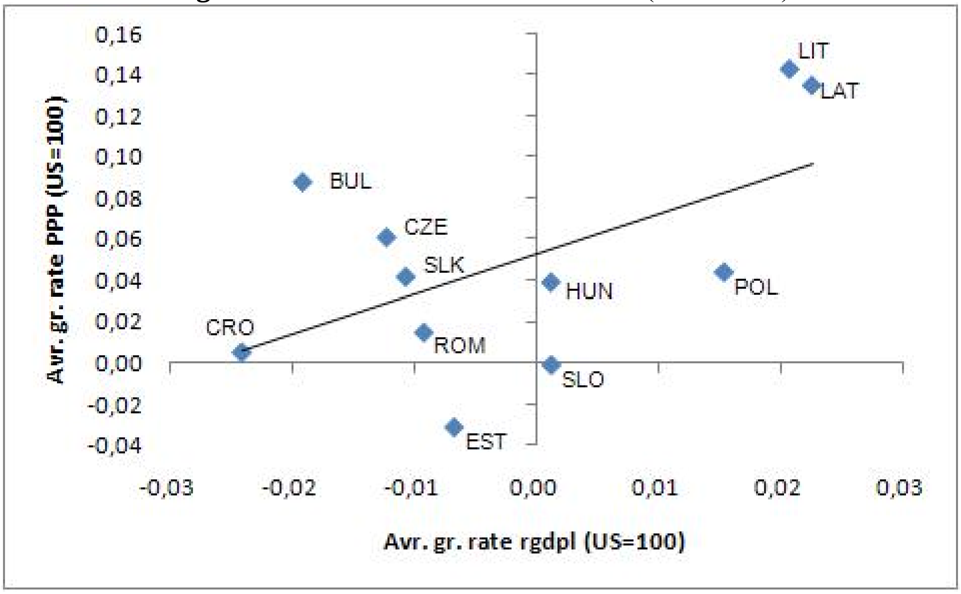
	I. Panel unit Root Tests					
	p_i/p_0	a_i/a_0	g_i/g_0	p_i	a_i	g_i
IPS	0.48225	0.13584	0.05998	0.35497	0.37292	0.14508

	International		**Intranational**	
	II. OLS Cointegration Tests			
	Model 1	*Model 2*	*Model 1*	*Model 2*
Productivity	1.0027	0.8754	1.0026	0.8279
	(0.000)	(0.000)	(0.000)	(0.000)
Government	--	0.1269	--	0.1742
	--	(0.000)	--	(0.000)
DF_ρ	0.000	0.000	0.000	0.000
DF_ρ^*	0.000	0.000	0.000	0.000
PC_ν	0.000	0.000	0.000	0.000
PC_ρ	0.000	0.000	0.000	0.000
GC_ρ	0.000	0.000	0.000	0.000
	III. DOLS Cointegration Tests			
Productivity	0.5495	0.5442	0.2643	0.3366
	(0.000)	(0.000)	(0.000)	(0.000)
Government	--	-0.1706	--	0.1503
	--	(0.000)	--	(0.000)
DF_ρ	0.000	0.000	0.000	0.000
DF_ρ^*	0.000	0.000	0.000	0.000
PC_ν	0.000	0.000	0.000	0.000
PC_ρ	0.000	0.000	0.000	0.000
GC_ρ	0.000	0.000	0.000	0.000

Notes: p-values for coefficient estimates in parenthesis,
panel statistics are p-values under the null hypothesis of *no* cointegration.

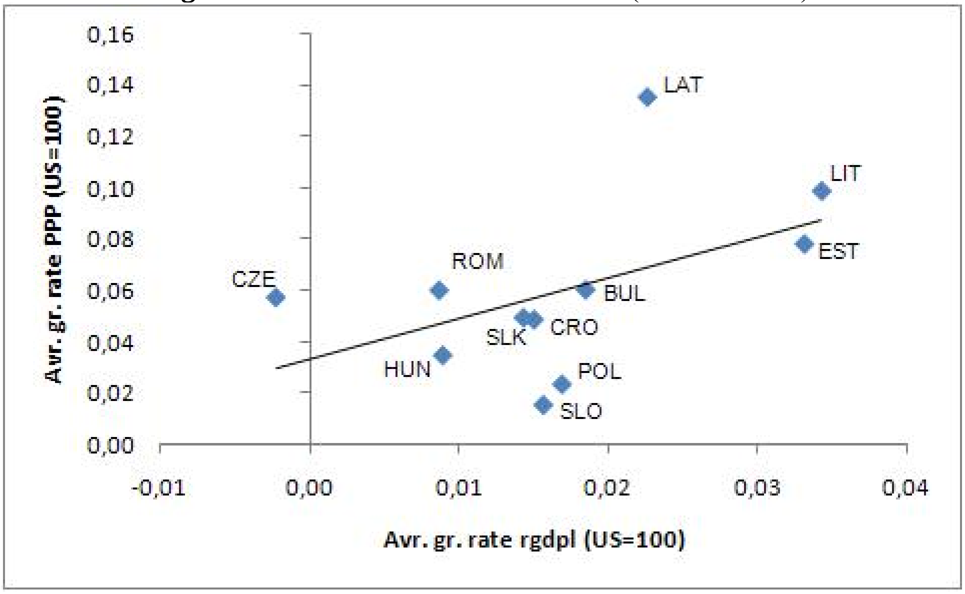
Figures

Figure 1: Growth rate of PPP and GDP (1990-2004)



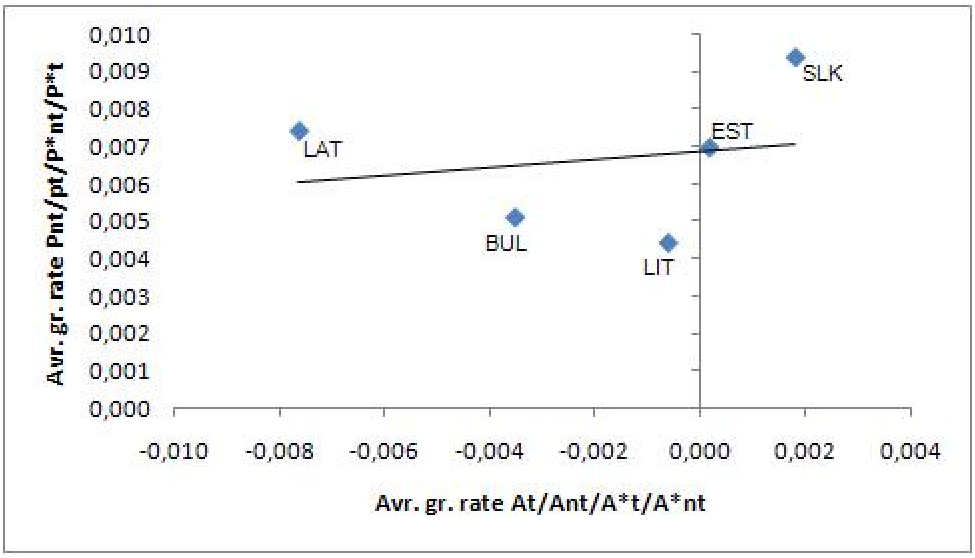
Source: Penn World Table 6.2

Figure 2: Growth rate of PPP and GDP (NADIR-2004)



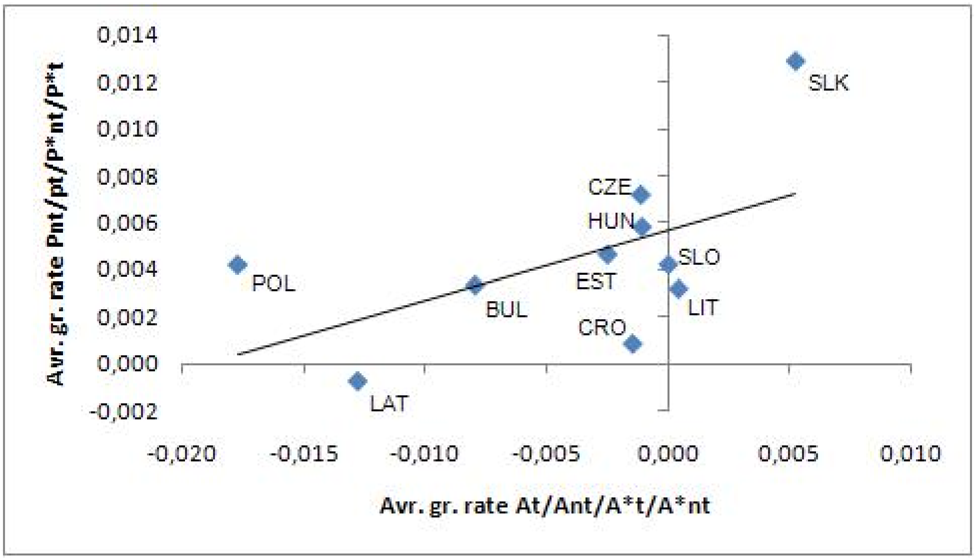
Source: Penn World Table 6.2

Figure 3: Growth rate of p_i/p_0 and a_i/a_0 (quarterly data 1995:I-2008:III)



Source: Eurostat

Figure 4: Growth rate of p_i/p_0 and a_i/a_0 (quarterly data 2000:I-2008:III)



Source: Eurostat