

Sonora, Robert J.; Tica, Josip

Article

Harrod, Balassa, and Samuelson (re)visit Eastern Europe

Cogent Economics & Finance

Provided in Cooperation with:

Taylor & Francis Group

Suggested Citation: Sonora, Robert J.; Tica, Josip (2014) : Harrod, Balassa, and Samuelson (re)visit Eastern Europe, Cogent Economics & Finance, ISSN 2332-2039, Taylor & Francis, Abingdon, Vol. 2, Iss. 1, pp. 1-16,
<https://doi.org/10.1080/23322039.2014.920557>

This Version is available at:

<https://hdl.handle.net/10419/147709>

Standard-Nutzungsbedingungen:

Die Dokumente auf EconStor dürfen zu eigenen wissenschaftlichen Zwecken und zum Privatgebrauch gespeichert und kopiert werden.

Sie dürfen die Dokumente nicht für öffentliche oder kommerzielle Zwecke vervielfältigen, öffentlich ausstellen, öffentlich zugänglich machen, vertreiben oder anderweitig nutzen.

Sofern die Verfasser die Dokumente unter Open-Content-Lizenzen (insbesondere CC-Lizenzen) zur Verfügung gestellt haben sollten, gelten abweichend von diesen Nutzungsbedingungen die in der dort genannten Lizenz gewährten Nutzungsrechte.

Terms of use:

Documents in EconStor may be saved and copied for your personal and scholarly purposes.

You are not to copy documents for public or commercial purposes, to exhibit the documents publicly, to make them publicly available on the internet, or to distribute or otherwise use the documents in public.

If the documents have been made available under an Open Content Licence (especially Creative Commons Licences), you may exercise further usage rights as specified in the indicated licence.



<http://creativecommons.org/licenses/by/3.0/>



Received: 19 November 2013
Accepted: 21 April 2014
Published: 10 June 2014

*Corresponding author: Robert J. Sonora, Department of Economics, School of Business Administration, Fort Lewis College, Durango, CO, USA
E-mail: sonora_t@fortlewis.edu

Reviewing Editor:
Steve Cook, Swansea University

Further author and article information is available at the end of the article

RESEARCH ARTICLE

Harrod, Balassa, and Samuelson (re)visit Eastern Europe

Robert J. Sonora^{1*} and Josip Tica²

Abstract: In this paper, we investigate the Harrod–Balassa–Samuelson (HBS) hypothesis in 11 Central and Eastern European transition countries. Unlike previous research, we test the HBS hypothesis with NACE 6 quarterly data which enables us to divide data into tradable and nontradable sectors without requiring unrealistic assumptions on the nature of the data. Contrary to previous results, we are only able to find evidence for univariate HBS effects in Bulgaria, Croatia, Hungary, and Poland. However, using panel cointegration tests, we find strong statistical evidence for the HBS hypothesis within countries and across countries. Our results also demonstrate that cross-country HBS holds under the assumption that the law of one price for tradables does not hold. Finally, we find, contrary to theory, that government consumption negatively impacts relative prices. The policy implications are that failing to acknowledge the peculiarities of the transition process results in suboptimal monetary policy.

Keywords: Harrod–Balassa–Samuelson effect, price convergence, transition countries, panel cointegration tests

JEL classifications: E31, F31, F41

1. Introduction

Debate about the Harrod–Balassa–Samuelson (hereafter HBS) hypothesis is one of the major stylized facts about transition.¹ Inquiry into HBS effects in Central and Eastern Europe was initiated by Halpern and Wyplosz's (1997) attempt to explain peculiar movements of real exchange rates in Eastern Europe. At that time, the unprecedented appreciation trend of real exchange rates in transition countries needed theoretical explanation. Real exchange rates in transition countries appreciated between 7.5% in Slovenia and 800% in Latvia. The dramatic appreciation was attributed to catch-up following initial undervaluation of transition countries and HBS theory.

The debate culminated with the question whether the strength of HBS effect in transition countries is strong enough to interfere with Maastricht criteria of the European Monetary Union (EMU). Three studies suggest there is interference between EMU rules and the HBS effect within certain countries, see Halpern and Wyplosz (2001), De Broeck and Sløk (2001), and Lojschova (2003). Other research suggests there is a substantial amount of evidence for cointegration between productivity and price levels, but no evidence of interference between convergence induced inflation and EMU rules, see Cipriani (2001), Coricelli and Jazbec (2001), Égert (2002a, 2002b), Égert, Drine, Lommatzsch, and Rault (2003), and Mihaljek and Klau (2003, 2008). On the other hand, Fischer (2002), Arratibel, Rodríguez-Palenzuela, and Thimann (2002), and Petrović (2012) do not find any evidence for the HBS effect in transition countries. The overall consensus is that the HBS effect is present, but is not large enough to interfere with Maastricht rules.

The goal of this paper is to readdress the question of HBS in transition countries on a larger and more consistent data sample. Using Eurostat's national account quarterly data (NACE 6) enables us

to implement the De Gregorio, Giovannini, and Wolf (1994) definition of sector tradability and perform univariate and panel cointegration tests without relying on the assumptions that were used in previous studies. Several authors confronted by data restrictions were forced to use cross-sectional estimates over any given year, as in Officer (1976).² Others have employed panel estimates to circumvent the power problem associated with short time series.³ Though somewhat restricted by data availability, some research has been conducted using time series models with monthly data, Égert (2002a, 2002b), Égert et al. (2003), and Mihaljek and Klau (2003).

In an attempt to compile enough observations for testing, several 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. Monthly data limited Égert (2002a, 2002b) to assume the nontradable sector productivity is equal to zero. Égert 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 (2003) used quarterly growth rates of residual between growth rates of annual GDP and quarterly industrial production as output of nontradable sector in several countries. Research that made significant methodological contributions such as Chong, Jordà, and Taylor (2012) used relative GDP per capita as a proxy for relative productivity instead of relative tradable to nontradable productivity.⁴

In this paper, we employ univariate and panel cointegration tests using Eurostat's NACE 6 data to reduce the noise and problems associated with restrictive *ex-ante* assumptions; a small number of observations; and heterogenous sources of data. Overall, we find that univariate tests of HBS are less successful than the extant literature. However, we do find considerable evidence for single country and cross-country HBS effects using panel methods. We demonstrate that in the international HBS hypothesis compared to tradable prices, the HBS effect is two to three times smaller with respect to real exchange movements, which is in line with Engel (1999) and Cincibuch and Podpiera (2006). Finally, we find, contrary to theory, that government consumption negatively impacts relative prices.

The policy implication is that inflation generating processes in transition countries are more complex compared to the mainstream closed-economy macroeconomic models usually employed to address issues of optimal monetary policy and price stability. With respect to transition countries, failure to account for mechanisms that drive changes in relative sectoral productivity and the terms of trades may result in suboptimal policies.

The remainder of the paper is divided in five sections: Section 2 provides a theoretical explanation of HBS effect. Section 3 discusses the data and provides an overview of the data properties. Section 4 summarizes the empirical strategy. In Section 5, we discuss the univariate and panel econometric methods used and present the results, and Section 6 provides some brief summary remarks.

2. The HBS Theory

The HBS model predicts 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 (Balassa, 1964; Harrod, 1933; Samuelson, 1964).

To motivate the discussion, consider a small country with perfect factor mobility whose tradable goods' prices are pinned down by world price levels. A positive shock to the tradable goods sector will have no effect on domestic prices of tradable goods. However, because positive productivity shocks augment the marginal product of labor, wages in the tradable sector rise. In the absence of productivity growth in the nontradable sector, with internal labor mobility, wages in that sector must rise to match wage increases in the tradable sector pushing up prices in that sector. This causes the relative price of nontradable goods to rise, increasing the aggregate price level.

The intuition behind the HBS hypothesis is that tradable goods, for example, manufacturing tend to experience faster productivity growth than do nontradable goods. On the other hand, nontraded

goods, which tend to be service oriented, use less technological expertise. This implies that economies with a more productive tradable good sector should have higher price levels than less productive ones. Therefore, because the nontradable sector experiences slower productivity growth than the traded sectors, deviations in the real exchange rate are long lived.

Assume Cobb–Douglas production with perfect competition, perfect international capital mobility—the domestic real interest rate equals the world real interest rate—perfect mobility of factors between sectors within the economy, and the law of one price in the tradable sector. It can be shown 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), time subscripts are repressed for clarity,

$$\dot{p}_N - \dot{p}_T = \frac{\beta}{\alpha} \dot{a}_T - \dot{a}_N \quad (1)$$

where p_N , p_T , a_N , and a_T are the index of nontradable prices, the index of tradable goods, productivity of nontradables and tradables, respectively.⁵ α , $\beta \in (0, 1)$ are the capital shares in the tradable and nontradable sectors, respectively. For any variable, X , $\dot{X} = d \ln X / dt$ is the time derivative of that variable. Equation 1 represents an *internal* transmission mechanism of the HBS effect or—if we assume that service intensive sector is equal to nontradable sector—the Baumol–Bowen effect (Froot & Rogoff, 1995).

In order to estimate the *external* transmission mechanism of the HBS effect, we express all variables vis-à-vis a numeraire country (Froot & Rogoff, 1995)

$$(\dot{p}_N - \dot{p}_T) - (\dot{p}_N^* - \dot{p}_T^*) = \left(\frac{\beta}{\alpha} \dot{a}_T - \dot{a}_N \right) - \left(\frac{\beta^*}{\alpha^*} \dot{a}_T^* - \dot{a}_N^* \right) \quad (2)$$

where “*”s denote foreign, or numeraire, country variables. Equation 2 is the international analog to the intranational HBS model in Equation 1. If purchasing power parity holds in the tradable sector, Equation 2 represents the real exchange rate. If not, the real exchange rate is derived from the general price level p which is a weighted average of tradable and nontradable sector prices using Equation 1.

$$\dot{p} = \dot{p}_T + \gamma \left(\frac{\beta}{\alpha} \dot{a}_T - \dot{a}_N \right) \quad (3)$$

where $\gamma > 0$ is the weight of nontradable goods in general price index. The right-hand side of Equation 3 is substituted for domestic p and foreign p^* general price level in the inverted real exchange rate equation $\dot{q} = \dot{p} - (\dot{e} + \dot{p}^*)$. Therefore, in the second version of the international HBS theory, the real exchange rate is a function of relative prices of tradable sector and relative productivity vis-à-vis the numeraire country is given by

$$\dot{q} = \dot{p}_T - (\dot{e} + \dot{p}_T^*) + \gamma \left(\frac{\beta}{\alpha} \dot{a}_T - \dot{a}_N \right) - \gamma^* \left(\frac{\beta^*}{\alpha^*} \dot{a}_T^* - \dot{a}_N^* \right) \quad (4)$$

An additional contribution to the model was made by Bergstrand (1991) who integrated government expenditure to examine the effects of the demand side on relative prices. The logic behind the model was the assumption that government spending preferences are biased in the direction of nontradables, which should connect the share of government spending to relative prices.

3. The Data Properties

Quarterly data have been compiled for Germany and 11 Eastern European transition countries: Bulgaria, Croatia, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Romania, Slovakia, and

Slovenia. Data for Germany are available from 1991.1, Estonia from 1993.1, Bulgaria, Latvia, Lithuania, and Slovakia from 1995.1, Czech Republic from 1996.1, Croatia from 1998.2, Hungary and Slovenia from 2000.1, and Poland from 2001.1. Due to small number of observations, Romania (2002.1–2006.4) were excluded from further analysis. From the data availability, we divide the data into panels. The first spans 1995.1–2009.1 and includes Bulgaria (BUL), Estonia (EST), Latvia (LAT), Lithuania (LIT), and Slovakia (SLK). The second covers the period 2001.1–2009.1 and adds Croatia (CRO), the Czech Republic (CZE), Hungary (HUN), and Slovenia (SLO) to the above countries.

Theoretical papers on the HBS are based on a precise division of goods into tradables and nontradables. However, for empirical applications, the definition of traded and nontraded goods is less parsimonious. Beginning with Officer (1976), most researchers simply assume that the manufacturing and/or industrial sectors are tradable while the services serve as nontradables. A large number of researchers add agriculture to tradables and almost the same number of them exclude it as a centrally administered price, particularly in economies with strong agricultural policies.

To date, the tradability of sectors has been examined only once. De Gregorio et al. (1994, pp. 1230–1232) empirically tested the degree of tradability in various sectors of an economy. They used a 10% ratio of exports to total production threshold in each sector in order to estimate the “tradedness” for 20 sectors across 20 OECD countries over the period 1970–1985. Although the 10% threshold is somewhat “ad hoc,” it proved to be quite robust. Cutting the threshold to 5% would have no effect, and raising it to 20% would change only the quantitatively small non-metal mineral products from the tradable to nontradable sector.

According to their test, agriculture, mining, and most of manufacturing had a share of exports in total production between 23.6 and 59.9%, agriculture having the lowest and metal manufacturing the highest shares.⁶ On the other hand, the share of exports of services was lower than 5%. Transportation had a share of 27.8%, while other services had an export share of 1.9%. Therefore, agriculture and mining were classified as tradables, as well as manufacturing and transportation. The remaining services, accounting for about 50–60% of GDP, were treated as nontradables, see Table A1 in Appendix A.

While De Gregorio et al.’s (1994) division of the economy did not become a standard for future research, we will follow their approach to sector division with one exception. To increase the number of quarterly observations we use the NACE 6 data-set. The quarterly NACE 17 database has all the branches necessary to perform De Gregorio et al.’s division of sectors according to “tradedness,” but data are only available for a small number of countries. On the other hand, the NACE 6 database has a 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 as presented in Table A2 in Appendix A.

We use average labor productivity as a ratio of sector’s gross value added (chain indexed 2000 prices) and total employment. Productivity of agriculture, hunting and forestry, and industrial branches are used to proxy for productivity in the tradable sector, and productivity of four nontradable branches is used as an proxy for productivity of nontradable sector.⁷

Weighted average of price indices in agriculture, hunting and forestry, and industry is used as price level in tradable sector and weighted average of price indices of four nontradable branches is used as price level for nontradables.⁸

For government expenditures, we use the real government expenditure/GDP ratio. To construct the real exchange rate and the tradable sector real exchange rate, we use all NACE branches of the price index and quarterly nominal exchange rate at the end of period expressed as national currency per euro. The numeraire country used in the estimation of the external mechanism is Germany.⁹

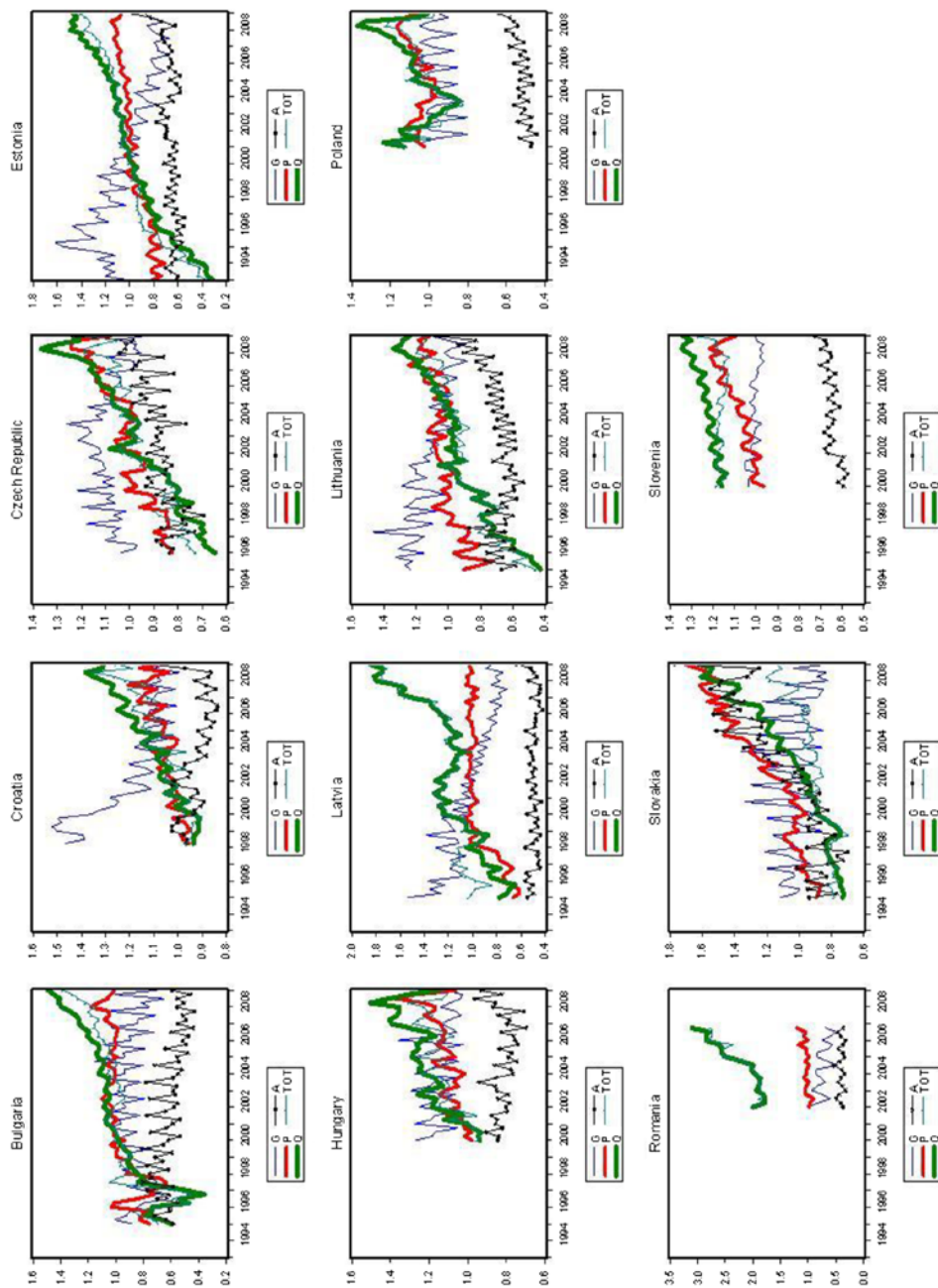


Figure 1. Data time series (vis-à-vis Germany).

Source: Eurostat.

Figure 1 depicts the nontradable–tradable goods price, P , the relative productivity, A , the real exchange rate, Q , terms of trade, TOT , and the government share of real GDP, G . In all countries, we see that each experienced an appreciation of their real exchange rate and terms of trade over the sample period which suggests that each country's currency was undervalued and/or nontradable goods prices were held artificially low in the centrally planned economies. This is borne out by a rise in relative prices over the same period. What is also striking is the size of the government sector in each economy.

The data descriptive statistics, found in Table 1, are also informative. The table only considers data for the period 2001.1–2009.2 and includes the mean and standard deviation of each variable. Intranational results are in the top half and the international statistics are in the bottom, with all data relative to the numeraire country, Germany. Concentrating first on the intranational data, we can see that there is little homogeneity within regions; for example, the Baltic countries of Estonia, Latvia, and Lithuania data are not clustered. Another striking feature is the mean of Slovakia's relative prices and productivity is quite a bit higher than anywhere else—though this could be a function of high nontradable prices and low nontradable good productivity. Turning our attention to the international results, we see similar results for the data which are discussed in intranational data, which are intuitively attractive as all the transition economy data are weighted by German variables. With respect to the other variables, another Baltic country, Latvia, has the highest terms of trade and real exchange rate, hinting at an overvalued currency. Overall, these statistics suggest a large degree of price and real exchange rate heterogeneity over the sample countries.

Table 1. Descriptive Statistics: 2001.1–2009.2

	Relative price		Relative productivity		Government share		Real exchange rate		Terms of trade	
	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.
<i>Intranational statistics</i>										
BUL	.04	1.058	.081	.687	.03	.187	–	–	–	–
CRO	.062	1.091	.063	1.113	.018	.23	–	–	–	–
CZE	.091	1.086	.141	1.126	.022	.226	–	–	–	–
EST	.055	1.067	.092	.841	.026	.177	–	–	–	–
LAT	.036	1.033	.048	.602	.025	.185	–	–	–	–
LIT	.062	1.105	.158	.884	.021	.217	–	–	–	–
HUN	.077	1.144	.075	.998	.02	.237	–	–	–	–
POL	.049	1.081	.077	.639	.013	.2	–	–	–	–
SLK	.207	1.395	.331	1.546	.036	.208	–	–	–	–
SLO	.07	1.115	.07	.797	.009	.208	–	–	–	–
<i>International statistics</i>										
BUL	.043	1.039	.089	.566	.124	.903	.134	1.176	.106	1.085
CRO	.055	1.072	.061	.914	.065	1.111	.11	1.162	.091	1.109
CZE	.09	1.066	.088	.921	.079	1.089	.121	1.08	.068	1.034
EST	.052	1.048	.078	.69	.108	.855	.152	1.21	.108	1.165
HUN	.077	1.124	.062	.819	.09	1.145	.111	1.242	.077	1.147
LAT	.028	1.015	.042	.494	.103	.889	.235	1.348	.229	1.331
LIT	.067	1.086	.109	.722	.079	1.048	.123	1.074	.094	1.014
POL	.056	1.062	.045	.523	.076	.965	.127	1.078	.101	1.032
SLK	.194	1.369	.214	1.257	.151	1.001	.221	1.203	.099	0.974
SLO	.07	1.096	.031	.652	.024	1.005	.052	1.241	.027	1.167
SLO	.07	1.096	.031	.652	.024	1.005	.052	1.241	.027	1.167

4. Methodology

We begin by applying univariate unit root and cointegration tests to test for the existence of HBS effects in transition countries individually. The relatively short sample periods and the nature and availability of sectoral data for transition countries lend itself to expand analysis to utilize panel unit root and cointegration tests, below, and enable us to cope with the power problem associated with small samples. Our regression analysis focuses on estimating Equations 1, 2, and 4 with and without share of government in GDP as additional explanatory variable.

The intranational version of the model Equation 1 is estimated

$$p_t = \beta_0 + \beta_1 a_t + \beta_2 g_t + e_t \quad (5)$$

where p_t is the tradable to nontradable relative price, a_t is the relative productivity and g_t is the share of government in GDP; $e_t \sim iid(0, \sigma_e^2)$. *Ex ante* theory suggests the estimated coefficients to have the following signs, $\beta_1 > 0$ and $\beta_2 > 0$.

Two versions of the international HBS hypothesis are estimated. The first assumes PPP holds for tradables as in Equation 2. This model is the international analog to Equation 5 and uses the variables p , a , and g expressed vis-à-vis the numeraire country, we denote these relative variables with “hats,” e.g. $\hat{a} = a_i/a_0$ for any country i . Estimates of the coefficients are predicted to have similarly signs as in the intranational HBS model, but of different magnitudes.

The second international model relaxes the PPP assumption for tradable goods, Equation 4, and uses the real exchange rate q as the regressand

$$q_t = \beta_0 + \beta_1 \hat{a}_t + \beta_2 \hat{g}_t + \beta_3 ToT_t + u_t \quad (6)$$

where q is the real exchange rate, ToT is the real exchange rate of tradable sector (terms of trade); and \hat{a} and \hat{g} represent the deviations discussed above. $u_t \sim iid(0, \sigma_u^2)$. As in the above models, theory hypothesizes $\beta_1 > 0$, $\beta_2 > 0$, and $\beta_3 > 0$.

5. Cointegration Tests and Results

5.1. Univariate Model

The HBS theory requires the above model to be cointegrated. That is there is long-run equilibrium relationship between the variables. We begin by testing the HBS model using both the intra- and inter-national versions of the theory, Equations 1, 2, and 4, respectively, using the Johansen (1991) VECM cointegration methodology on the univariate series to test for the existence of the long-run equilibrium relationship given by the HBS hypothesis.¹⁰

In order to avoid spurious regressions, the analysis begins with unit root tests and proceeds to testing for cointegration for those countries with nonstationary series. Results of the standard Augmented Dickey-Fuller (ADF) unit root test indicate that most of the variables are nonstationary, but only in a small number of countries all variables are nonstationary. In the intranational unit root test, no country have p and a of the same order of integration. In the international test with PPP assumption, three variables for Poland ($\hat{p}, \hat{a}, \hat{g}$) and two variables (\hat{p}, \hat{a}) for Lithuania and Croatia are nonstationary. In the international test without PPP assumption, only Bulgaria and Hungary q and \hat{a} are $I(1)$.¹¹

Following the unit root test results, cointegration tests are performed on international data for Croatia, Lithuania, Bulgaria, and Poland, which are presented in Table 2. The estimates demonstrate that relative price levels vis-à-vis Germany are cointegrated with relative productivity in Croatia. In the case of Poland, results are ambiguous due to different number of cointegrating vectors implied by the Johansen trace and max-eigenvalue statistics. In the international model

Table 2. Univariate Cointegration Tests International Version (vis-à-vis Germany)

	Trace stat	No. of coint. vectors	Max eigen stat.	No. of coint. vectors	Cointegrating vector							
					$\hat{\rho}$	q	$\hat{\alpha}$	$\hat{\eta}$	ToT	Constant	Trend	
CRO	17.55	1	15.00	1	1.00		.80 (7.82)					
POL	15.56	1	8.06	0	1.00		-.38 (-1.80)	-.28 (-.43)			-.30 (-2.36)	
BUL	35.90	4	27.08	4		1.00	.28 (1.77)	.37 (3.27)	-.74 (-8.28)		.32	-.004
HUN	33.97	1	16.68	0		1.00	.19 (1.91)	.39 (2.91)	-1.17 (-15.72)		-.07 (-2.51)	

Note: t Statistics for coefficient estimates in parenthesis.

without the PPP assumption, results for Bulgaria indicate strong evidence of cointegration while results for Hungary are ambiguous. Lithuania is omitted from Table 2 due to lack of cointegrating vectors.¹²

Results of the univariate unit root test and cointegration tests imply that estimated relationship between relative prices and productivity is either statistically insignificant and/or spurious in majority of the countries in the sample. We conclude the results of the univariate tests do not provide sufficient evidence for the HBS hypothesis in our sample. However, the lack of success in the univariate tests could be the result of the power problem caused by the short sample period. In order to avoid the power problem, we next employ panel unit root and cointegration tests.

5.2. Panel Model

Given the highly technical nature of panel cointegration, we briefly outline the panel cointegration methods used.¹³ The following synopsis assumes the reader is familiar with the cointegration basics. We begin by considering the panel analog to the univariate Engle–Granger cointegration method which begins by estimating the long-run cointegrating relationship given in the fixed-effect panel regression

$$y_{i,t} = \alpha_i + x'_{i,t}\beta + e_{i,t}, i = 1, \dots, N, t = 1, \dots, T \tag{7}$$

where $\{y_{i,t}\}$ are 1×1 ; β is an $M \times 1$ vector of slope parameters; α_i are fixed effects; and the disturbance term $e_{i,t} \sim I(0)$. 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 Equation 7 represents a system of cointegrating regressions.

As in the Engle–Granger (Engle & Granger, 1987) cointegration test, the first step is to estimate the cointegrating vector, estimated from Equation 7 above, $(1, -\hat{\alpha}, -\hat{\beta})'$. Three methods to estimate the long-run cointegrating vector have been proposed: “standard” OLS, Fully Modified OLS (FMOLS), and Dynamic OLS (DOLS). Kao and Chiang (2000) derive the limiting distributions for the OLS, FMOLS, and DOLS estimators for the regression specification given in Equation 7. They also investigated the finite sample properties of each estimator through Monte Carlo simulation. They found that (1) the OLS estimator has a non-negligible bias; (2) the FMOLS estimator does not improve on the OLS estimator in general; and (3) the DOLS estimator has the best properties of the three. Given the small gains to using FMOLS we restrict our analysis to the OLS and DOLS estimators. Appendix B provides a more detailed summary of the methods and Appendix C reviews the relevant cointegration test statistics.

5.3. Panel Results

Figures 2 and 3 show the average growth rates of relative productivity and price levels relative to vis-à-vis Germany for the two sample panels. It is clear that most transition countries in each panel experienced a higher relative price and productivity growth than Germany. As is the case for univariate cointegration methods, the order of integration for each of the variables must be analyzed. For each Panel, we begin the formal analysis with the Im, Pesaran, and Shin (2003) panel unit root tests which is a panel representation of the standard ADF model.

This representation allows for heterogeneous intercepts and AR(1) parameters. We are concerned with the Studentized *t*-statistics on the estimated AR(1) parameter for the panel as a whole. Table 3 presents results for the five variables in the international test and three variables in intranational test for the first panel spanning from 1995.1. The bottom half presents results for the panel spanning from 2001.1. All international variables in both panels are expressed vis-à-vis Germany.

The panel unit tests demonstrate that all intranational variables (*p*, *a*, and *g*) in the 1995 panel are indeed nonstationary at standard test critical values. In the international version of the first panel, relative prices \hat{p} and government consumption \hat{g} are nonstationary at standard test critical values and productivity \hat{a} is stationary at 10% level. In the second panel, relative prices *p* and government

Figure 2. Growth rate of \hat{p} and \hat{a} (Quarterly data 1995.1–2009.2).

Source: Eurostat.

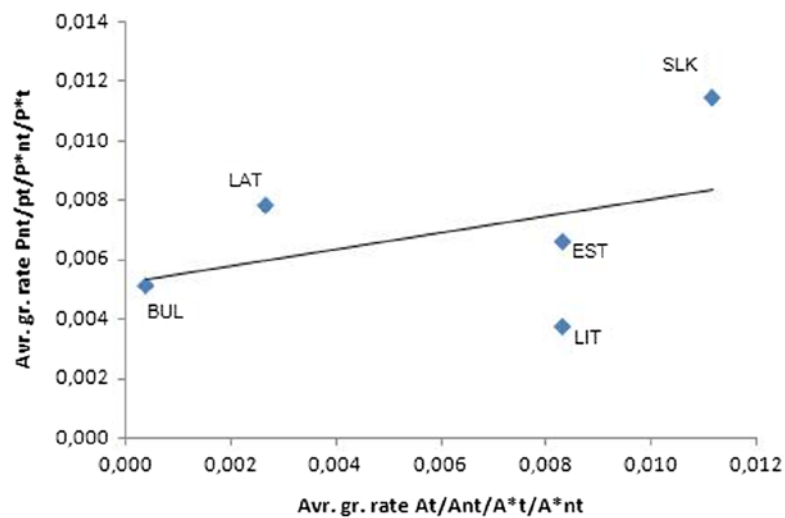


Figure 3. Growth rate of \hat{p} and \hat{a} (Quarterly data 2001.1–2009.2).

Source: Eurostat.

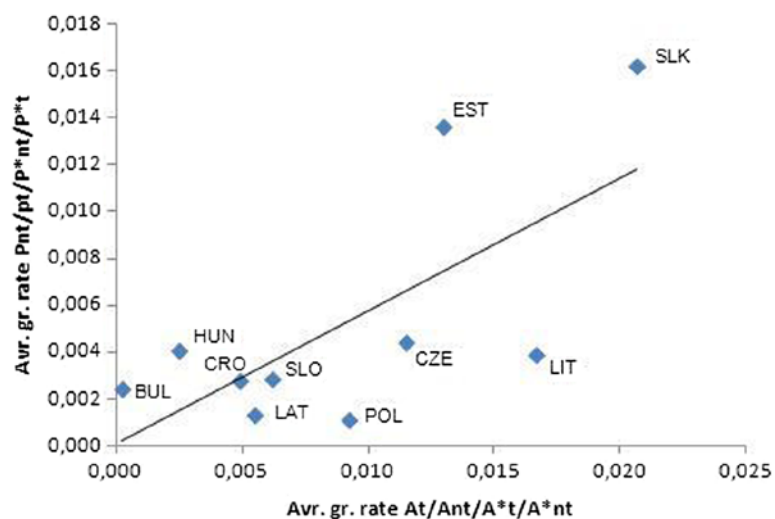


Table 3. *p*-Values Panel Unit Root Tests

	<i>p</i>	<i>a</i>	<i>g</i>	\hat{p}	\hat{a}	\hat{g}	<i>q</i>	ToT
1995:I–2009:I								
t_{IPS}	.242	.104	.178	.425	.052	.486	.000	.010
2001:I–2009:I								
t_{IPS}	.278	.032	.250	.156	.000	.093	.000	.000

consumption *g* are nonstationary in intranational version. In the international part, price levels \hat{p} are nonstationary and government consumption is stationary at 10%.

Table 4 presents the results of the panel cointegration test. Results for the intranational HBS are in the top third of the table and are denoted Version I. The tests for the international version of the model assuming that PPP holds in the tradable sector, Equation 2, Version II can be found in middle third of the table. The bottom third of the table results from the international version of the model without PPP assumption for the tradable sector, Equation 4. Results from each of the two models, Model 1, without a government shares, and Model 2, including government shares, are presented for each of the panels. Both the bias-adjusted OLS and DOLS cointegrating vectors are presented, with *p*-values in parenthesis. *p*-values for the overall cointegration test statistics denoted DF_{ρ} and DF_{ρ}^* for Kao's (1999) ADF-based tests and the Pedroni (1999) tests PC_{ν} , PC_{ρ} , and GC_{ρ} , are also tabulated. Details of each test can be found in Appendix C.¹⁴

Cointegration results indicate that null hypothesis of no cointegration is strongly rejected in 1995 panel intranational versions with the Kao (1999) test, while the Pedroni (1999) test rejects the null hypothesis only in the DOLS estimates. Estimated coefficients in Model 2 of the 1995 intranational panel are .34 for productivity *a* and –1.20 for government consumption *g* and .36 for productivity in the Model 1. Results of the 2001 intranational panel are ambiguous due to the fact that productivity is stationary at 10%.

In the international test with PPP assumption cointegration results for the 1995 panel are ambiguous due to the fact that productivity \hat{a} is stationary at 10%. Both cointegration tests indicate rejection of the null hypothesis at the conventional levels. Estimated coefficients for productivity are between .50 and .63 and for government consumption between –.22 and –.27. Compared to the intranational test, coefficients for both explanatory variables are larger in absolute terms. Estimates for the 2001 panel are strongly spurious with \hat{a} stationary at 1% significance level.

For the international test without PPP assumption for tradables, two major variables, the real exchange rate and the real exchange rate of tradable sector are stationary at 1% in both 1995 and 2001 panels. Noting that productivity \hat{a} in 2001 panel is also stationary at 1%, it is possible to analyze results of the Model 1 in 2001 international panel as fixed-effect panel of stationary variables that will by definition result with stationary residuals. In Model 1 of the 2001 panel, estimated coefficients for productivity are .23 (DOLS) and .30 (OLS), while estimated coefficients for tradable sector real exchange rate are 1.08 (DOLS) and 1.06 (OLS). Compared to international version with PPP, estimated coefficients for productivity are smaller and much closer to intranational results.

Compared to estimated coefficients for the real exchange rate for the tradable sector, estimated coefficients for productivity are, on average, 2–3 times smaller. Several conclusions can be drawn from these results:

- First, given that sample periods for transition economies are of insufficient length to overcome the power problem, we demonstrate the use of panel cointegration tests to identify the HBS effect in transition countries.

Table 4. Panel Cointegration Tests

	Panel 1995:I–2009:I				Panel 2001:I–2009:I			
	Model 1		Model 2		Model 1		Model 2	
	OLS	DOLS	OLS	DOLS	OLS	DOLS	OLS	DOLS
Version I. Intranational: p_i as dependent variable								
a_i	.584 (.000)	.359 (.000)	.369 (.000)	.342 (.002)	.493 (.000)	.270 (.000)	.464 (.000)	.289 (.316)
g_i	–	–	–1.169 (.000)	–1.200 (.000)	–	–	.020 (.447)	–.971 (.282)
DF_p	.000	.000	.000	.000	.000	.000	.000	.000
DF_p^*	.000	.000	.000	.000	.000	.000	.000	.000
PC_v	.000	.000	.000	.000	.000	.000	.000	.000
PC_p	.435	.028	.280	.000	.000	.000	.000	.000
GC_p	.435	.029	.282	.000	.000	.000	.000	.000
Version II. International: \hat{p} as dependent variable								
\hat{a}	.627 (.000)	.502 (.000)	.556 (.000)	.509 (.000)	.495 (.000)	.392 (.000)	.483 (.000)	.456 (.000)
\hat{g}	–	–	–.229 (.002)	–.272 (.001)	–	–	–.120 (.001)	–.343 (.000)
DF_p	.000	.000	.000	.000	.000	.000	.000	.000
DF_p^*	.000	.000	.000	.000	.000	.000	.000	.000
PC_v	.000	.000	.000	.000	.000	.000	.000	.000
PC_p	.000	.000	.000	.000	.000	.000	.000	.000
GC_p	.000	.000	.000	.000	.000	.000	.000	.000
Version III. International: q as dependent variable								
\hat{a}	.296 (.000)	.350 (.000)	.265 (.000)	.340 (.000)	.296 (.000)	.230 (.000)	.254 (.000)	.269 (.000)
\hat{g}	–	–	–.071 (.018)	–.142 (.000)	–	–	–.104 (.000)	–.269 (.000)
ToT	1.222 (.000)	1.235 (.000)	1.212 (.000)	1.178 (.000)	1.056 (.000)	1.081 (.000)	1.078 (.000)	1.020 (.000)
DF_p	.000	.000	.000	.000	.000	.000	.000	.000
DF_p^*	.000	.000	.000	.000	.000	.000	.000	.000
PC_v	.000	.000	.000	.000	.000	.000	.000	.000
PC_p	.000	.000	.000	.000	.000	.000	.000	.000
GC_p	.000	.000	.000	.000	.000	.000	.000	.000

Note: p -values for estimated coefficients are in parenthesis. For the panel cointegration tests, p -values for the null of no cointegration are presented.

- Second, the more robust methodology enables us to find evidence for HBS effects in data compiled in the way suggested by De Gregorio et al. (1994) without making unrealistic assumptions about the data. Given that previous research made strong a priori assumptions in comparable analysis, these finding provide more robust evidence of the HBS effect.
- Next, we find the PPP assumption for the tradable sector is too restrictive for the transition process. Findings for the international test without the PPP constraint in the tradable sector imply that part of the appreciation in real exchange rates in transition countries is attributed to appreciation in the tradable sector. Thus, reforms affect the slopes of the export demand curves of transition countries enabling them to increase exports parallel to real exchange rate appreciation.¹⁵

- Fourth, although the HBS effect is present and significant, its effect on real exchange rates is smaller compared to contribution of the changes in the relative prices of the tradable sector. The relative size of the coefficients on relative productivity and terms of trade indicate that movements in the prices of tradables explain a much larger share of appreciation in transition countries—implying a secondary but still statistically significant role for the HBS effect.
- Finally, the impact of government consumption on relative prices and the real exchange rate is contrary to theory and is idiosyncratic to the transition process and remains an avenue for further theoretical and empirical investigation.

6. Summary

Unlike previous studies on HBS in transition countries, we perform univariate and panel cointegration tests on much larger number of observations and with weaker set of *ex ante* assumptions. Also, the Eurostat NACE 6 classification data on prices, value added, and employment enable us to divide data into tradable and nontradable sector according to the De Gregorio et al. (1994) methodology.

Univariate results demonstrate it is possible to find cointegrating relationships only for Bulgaria, Croatia, Hungary, and Poland vis-à-vis Germany. Panel cointegration tests demonstrate evidence for cointegration in the intranational panel with government and productivity as explanatory variables. In the international model with PPP in the tradable sector, evidence for cointegration is unclear given the ambiguity of the unit root test results. In the international 2001 model without the PPP assumption for tradable sector, all variables are stationary at 1% (except government consumption) which means that OLS estimates will result in consistent and stationary residuals.

The estimated productivity coefficients are within the expected range found in other cross-section or panel estimates for transitional countries. In terms of the Maastricht rules, our results are in line with prevailing consensus that HBS is present, but it is not large enough to interfere with EMU rules.

Estimated government coefficients exhibit transition economy-specific behavior. According to theory, government consumption is biased towards nontradable goods and as a result of that, countries with higher government/GDP ratio are expected to have higher relative prices of nontradable sector (Bergstrand, 1991). Our results indicate negative results for government consumption in all cointegrated relationships. A possible explanation for this is that transition reforms have reduced government/GDP ratios while simultaneously increasing relative prices. In the international test without PPP assumption for tradable sector, it is possible to compare contribution of the relative productivity and deviations from PPP to the real exchange rate movements. Results indicate that contribution of changes of relative prices of tradables affected real exchange rates two to three times stronger compared to relative productivity.

Both the existence of evidence of HBS mechanism and terms of trade changes in transition countries imply that mainstream model thinking usually employed in analysis of real exchange rates, price stability, and convergence as well as international competitiveness might be misleading in the case of transition countries. Failure to acknowledge for the peculiarities of transition process in terms of relative price behavior can create major problems in managing economic policy for these countries.

Acknowledgments

We would like to thank participants at the 84th Annual Conference of the Western Economic Association Meeting in Vancouver, Canada and at the 20 years of Transition in Central and Eastern Europe Conference in London, UK for helpful comments. Especially helpful were suggestion made by John Devereaux, Young-Kyu Moh, and Dubravko Mihaljek. Any mistakes are, of course, ours alone

Author details

Robert J. Sonora¹
E-mail: sonora_t@fortlewis.edu

Josip Tica²

E-mail: jtica@efzg.hr

¹ Department of Economics, School of Business Administration, Fort Lewis College, Durango, CO, USA.

² Faculty of Economics and Business, Department of Macroeconomics and Economic Development, University of Zagreb, Trg J.F. Kennedy 6, 10000 Zagreb, Croatia.

Article Information

Cite this article as: Harrod, Balassa, and Samuelson (re)visit Eastern Europe, R.J. Sonora & J. Tica, *Cogent Economics & Finance* (2014), 2: 920557.

Notes

1. Throughout this paper the word “transition” refers to an economy’s transformation from a Soviet command style economy to a market-based one, as is used in the transition literature.
2. Halpern and Wyplosz (1997), Krajnyák and Zettelmeyer (1998), and Čihák and Holub (2001).
3. See, Halpern and Wyplosz (2001), Coricelli and Jazbec (2001), De Broeck and Sløk (2001), Flek, Marková, and Podpiera (2002), Jazbec (2002), Arratibel et al. (2002), Fischer (2002), and Lojschova (2003).
4. See Asea and Mendoza (1994) for the full derivation of the assumptions required for relative average productivity to proxy relative tradable to nontradable sector productivity.
5. An implication of the two sector model is the fact that even if we assume that productivity shocks in each sector are of the same magnitude, relative price of nontradable sector will increase if the traded sector is more capital intensive $\beta/\alpha > 1$ (Froot & Rogoff, 1995).
6. The only exception within these three sectors was the manufacturing of non-metallic minerals with a share of 13.7%.
7. Aggregate gross added value of tradable or nontradable sector is divided by aggregate employment of tradable or nontradable sector according to equation: $\frac{Y}{L} = \frac{\sum_{k=1}^S Y_k}{\sum_{k=1}^S L_k}$, where S denotes number of NACE branches within the tradable or nontradable sector.
8. Shares of gross added value of NACE branches are used as weights according to equation: $P = \sum_{k=1}^S \theta_k P_k$, where $\theta_k < 1$ is the output share of sector S , $\sum_{k=1}^S \theta_k = 1$, and P denotes price index and S is the number of NACE branches within the tradable or nontradable sector.
9. Results with eurozone as numeraire country are available from the authors upon request.
10. The Johansen (1991) method has become a standard in estimating and testing for cointegrating relationships; in the interest of saving space, we do not summarize it here.
11. Results are available from the authors on request.
12. Given that eurozone is the largest single bloc of trading countries for transition economies, we also analyzed HBS using the EZ as a single trading partner. While using the eurozone is intuitively appealing, it is difficult to reconcile it with the theoretical implications of the HBS hypothesis. Results for all tests using eurozone as the numeraire country are available from the authors on request.
13. For an overview of panel cointegration methods, see Banerjee (1999) and Kao (1999). See Kao and Chiang (2000), Phillips and Moon (1999), and Pedroni (1993, 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).
14. As with the univariate tests, results using eurozone as the numeraire country are available upon request. Overall, using eurozone, the results are similar to those found when Germany is numeraire, though the estimated coefficients tend to be slightly larger when using Germany as numeraire. We interpret this as due to the close relationship between the central and south-eastern European countries with Germany compared to the entirety of eurozone.
15. This results imply that export demand curves have rotated in a such way that—at least in the part of the export sector—Marshall–Lerner condition does not hold any more.

16. Much of this discussion is taken from Banerjee (1999) and Kao, Chiang, and Chen (1999).

References

- Arratibel, O., Rodríguez-Palenzuela, D., & Thimann, C. (2002). *Inflation dynamics and dual inflation in accession countries: A “new Keynesian” perspective*. ECB Working Paper, No. 132.
- Asea, P., & Mendoza, E. (1994). The Balassa–Samuelson model: A general-equilibrium appraisal. *Review of International Economics*, 2, 244–267.
<http://dx.doi.org/10.1111/roie.1994.2.issue-3>
- Balassa, B. (1964). The purchasing-power parity doctrine: A reappraisal. *Journal of Political Economy*, 72, 584–596.
<http://dx.doi.org/10.1086/jpe.1964.72.issue-6>
- Banerjee, A. (1999). Panel data unit roots and cointegration: An overview. *Oxford Bulletin of Economics and Statistics*, 61, 607–629.
<http://dx.doi.org/10.1111/obes.1999.61.issue-S1>
- Bergstrand, J. (1991). Structural determinants of real exchange rates and national price levels: Some empirical evidence. *American Economic Review*, 81, 325–334.
- Chong, Y., Jordà, Ò., & Taylor, A. M. (2012). The Harrod–Balassa–Samuelson hypothesis: Real exchange rates and their long-run equilibrium. *International Economic Review*, 53, 609–634.
<http://dx.doi.org/10.1111/iere.2012.53.issue-2>
- Čihák, M., & Holub, T. (2001). *Convergence of relative prices and inflation in central and Eastern Europe*. IMF Working Paper 124.
- Cincibuch, M., & Podpiera, J. (2006). Beyond Balassa–Samuelson: Real appreciation in tradables in transition countries. *The Economics of Transition*, 14, 547–573.
<http://dx.doi.org/10.1111/ecot.2006.14.issue-3>
- Cipriani, M. (2001). *The Balassa–Samuelson effect in transition economies*. Mimeo, IMF.
- Coricelli, F., & Jazbec, B. (2001). *Real exchange rate dynamics in transition economies*. CEPR Discussion Paper No. 2869.
- De Broeck, M., & Sløk, T. (2001). *Interpreting real exchange rate movements in transition countries*. IMF Working Paper 56.
- De Gregorio, J., Giovannini, A., & Wolf, H. C. (1994). International evidence on tradables and nontradables inflation. *European Economic Review*, 38, 1225–1244.
[http://dx.doi.org/10.1016/0014-2921\(94\)90070-1](http://dx.doi.org/10.1016/0014-2921(94)90070-1)
- Engel, C. (1999). Accounting for US real exchange rate changes. *Journal of Political Economy*, 107, 507–538.
<http://dx.doi.org/10.1086/250070>
- Engle, R. F., & Granger, C. W. J. (1987). Co-integration and error correction: Representation, estimation, and testing. *Econometrica*, 55, 251–277.
<http://dx.doi.org/10.2307/1913236>
- Égert, B. (2002a). Estimating the impact of the Balassa–Samuelson effect on inflation and the real exchange rate during transition. *Economic Systems*, 26, 1–16.
- Égert, B. (2002b). Investing the Balassa–Samuelson hypothesis in the transition, economics of transition. *Economics of Transition*, 10, 273–309.
- 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*, 31, 552–572.
- Fischer, C. (2002). *Real currency appreciation in accession countries: Balassa–Samuelson and investment demand*. BOFIT Discussion Papers No. 8.
- Flek, V., Marková, L., & Podpiera, J. (2002). *Sector productivity and real exchange rate appreciation: Much ado about nothing?*. Czech National Bank Working Paper Series 4/2002.
- Froot, K. A., & Rogoff, K. (1995). Perspectives on PPP and long-run real exchange rates. In G. Grossman & K. Rogoff (Eds.), *Handbook of international economics* (Vol. 3, pp. 1647–1688). Amsterdam: North-Holland.
- Halpern, L., & Wyplosz, C. (1997). Equilibrium exchange rates in transition economies. *Staff Papers—International*

- Monetary Fund*, 44, 430–460.
<http://dx.doi.org/10.2307/3867461>
- Halpern, L., & Wyplosz, C. (2001). *Economic transformation and real exchange rates in the 2000s: The Balassa–Samuelson connection*. Geneva: UN/ECE.
- Harrod, R. F. (1933). *International economics*. Nisbet and Cambridge University Press.
- Im, K. S., Pesaran, M. K., & Shin, Y. (2003). Testing for unit roots in heterogenous panels. *Journal of Econometrics*, 115, 53–74.
[http://dx.doi.org/10.1016/S0304-4076\(03\)00092-7](http://dx.doi.org/10.1016/S0304-4076(03)00092-7)
- Jazbec, B. (2002). *Balassa–Samuelson effect in transition economies—The case of Slovenia*. William Davidson Working Paper No. 507.
- Johansen, Soren (1991). Estimation and hypothesis testing of cointegration vectors in gaussian vector autoregressive models. *Econometrica*, 59, 1551–1580.
<http://dx.doi.org/10.2307/2938278>
- Kao, C. (1999). Spurious regression and residual-based tests for cointegration in panel data. *Journal of Econometrics*, 90, 1–44. [http://dx.doi.org/10.1016/S0304-4076\(98\)00023-2](http://dx.doi.org/10.1016/S0304-4076(98)00023-2)
- Kao, C., & Chiang, M.-H. (2000). On the estimation and inference of a cointegrated regression in panel data. *Advances in Econometrics*, 15, 179–222.
[http://dx.doi.org/10.1016/S0731-9053\(00\)15007-8](http://dx.doi.org/10.1016/S0731-9053(00)15007-8)
- Kao, C., Chiang, M.-H., & Chen, B. (1999). International R&D spillovers: An application of estimation and inference in panel cointegration. *Oxford Bulletin of Economics and Statistics*, 61, 691–709.
<http://dx.doi.org/10.1111/obes.1999.61.issue-S1>
- Krajnyák, K., & Zetelmeyer, J. (1998). Competitiveness in transition economies: What scope for real appreciations. *IMF Staff Papers*, 45, 309–362.
- Levin, A., & Lin, C. F. (1993). *Unit root tests in panel data: Asymptotic and finite-sample properties*. Unpublished Manuscript. San Diego: University of California.
- Lojschova, A. (2003). *Estimating the Impact of the Balassa–Samuelson effect in transition economies*. Vienna: Institute for Advanced Studies.
- Mihaljek, D., & Klau, M. (2003). *The Balassa–Samuelson effect in Central Europe: A disaggregated analysis*. BIS Working Paper 143.
- Mihaljek, D., & Klau, M. (2008). *Catching-up and inflation in transition economies: The Balassa–Samuelson effect revisited*. BIS Working Paper 270.
- Officer, L. H. (1976). The productivity bias in purchasing power parity: An econometric investigation. *Staff Papers—International Monetary Fund*, 23, 545–579.
<http://dx.doi.org/10.2307/3866641>
- Pedroni, P. L. (1993). *Panel cointegration, endogenous growth and business cycles in open economies* (PhD thesis). Columbia University.
- Pedroni, P. L. (1999). Critical values for cointegration tests in heterogeneous panels with multiple regressors. *Oxford Bulletin of Economics and Statistics*, 61, 653–670.
<http://dx.doi.org/10.1111/obes.1999.61.issue-S1>
- Petrović, P. (2012). Harrod–Balassa–Samuelson effect and the role of distribution sector: An empirical case study of Serbia and EMU. *Zbornik radova Ekonomskog fakulteta Rijeka*, 30, 57–87.
- Phillips, P. C. B., & Moon, H. R. (1999). *Nonstationary panel data analysis: An overview of some recent developments*. Working Papers in Economics: 98R/17. Santa Barbara: University of California.
- Rogoff, K. (1992). *Traded goods consumption smoothing and the random walk behavior of the real exchange rate*. NBER Working Paper No. 4119.
- Samuelson, P. A. (1964). Theoretical notes on trade problems. *The Review of Economics and Statistics*, 46, 145–154.
<http://dx.doi.org/10.2307/1928178>

Appendix A. Data Explanation Tables

Table A1. De Gregorio et al. NACE 17 Classification

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

Table A2. De Gregorio et al. NACE 6 Adjusted Classification

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

Appendix B. Statistical Methodology

B.1. OLS and DOLS estimators

We briefly review the OLS and Dynamic OLS (DOLS) methods for estimating the long-run cointegrating vector for panel cointegration (see Kao & Chiang, 2000).¹⁶ The OLS estimator of β is,

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

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, $\tilde{\beta}$:

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

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

$$\bar{\Lambda}_{OLS} = -3\hat{\Omega}_1^{-1} \mu'_{\epsilon_1} + 6\hat{\Omega}_1^{-1} \Delta'_{\mu\epsilon_1}$$

where Ω_1 is the estimated long-run covariance matrix, μ_{ϵ_1} is the first row of the estimated mean error and $\Delta_{\mu\epsilon_1}$ is the kernel estimates of the long-run covariance matrices.

The DOLS estimator is estimated from,

$$y_{it} = \alpha_i + x'_{it} \beta + \sum_{j=-\tau}^{\tau} c_{ij} \Delta x_{it-j} + v_{it} \quad (B.2)$$

which augments the standard OLS regression with leads and lags of vector of explanatory variables, x . Given the relatively short-term data, we only allow for up to $\tau=1$ lead and lags.

Appendix C. Panel Cointegration Test Statistics

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 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:

$$\hat{\epsilon}_{i,t} = \gamma_i \hat{\epsilon}_{i,t-1} + v_{i,t} \quad (C.3)$$

where the $\hat{e}_{i,t}$ are the estimated residuals from Equation 7. The null hypothesis of no cointegration is, as in the EG test,

$$H_0 : \gamma_i = 1, H_A : \gamma_i = \gamma < 1, \forall i \tag{C.4}$$

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

$$DF_\rho = \frac{\sqrt{NT}(\hat{\rho} - 1) + 3\sqrt{3}}{\sqrt{10.2}} \tag{C.5}$$

and

$$DF_\rho^* = \frac{\sqrt{NT}(\hat{\rho} - 1) + \frac{3\sqrt{N}\hat{\sigma}_v^2}{\hat{\sigma}_{0v}^2}}{\sqrt{3 + \frac{7.2\hat{\sigma}_v^4}{\hat{\sigma}_{0v}^4}}} \tag{C.6}$$

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

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 the residuals of the i th cross-unit, Equation 3. The first set of tests restricts this coefficient to be equal across all i units, as in Equation 4, which is similar to the restriction on the Levin and Lin (1993) AR parameter. The second set of tests relax the restriction on γ along the lines of the Im et al. (2003) panel unit root test:

$$H_0 : \gamma_i = 1, H_A : \gamma_i < 1, \forall i. \tag{C.7}$$

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 4, of the Pedroni panel cointegration statistics, the panel ν -statistic, PC_ν

$$PC_\nu = T^2 N^{\frac{3}{2}} \left(\sum_{i=1}^N \sum_{t=1}^T \hat{L}_{lli}^{-2} \hat{e}_{i,t-1}^2 \right)^{-1} \tag{C.8}$$

and ρ -statistic, PC_ρ

$$PC_\rho = T \sqrt{NZ} \hat{\rho}_{N,T} \left(\sum_{i=1}^N \sum_{t=1}^T \hat{L}_{lli}^{-2} \hat{e}_{i,t-1}^2 \right)^{-1} \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{lli}^{-2} (\hat{e}_{i,t-1} \Delta \hat{e}_{i,t} - \hat{\lambda}_i) \tag{C.9}$$

and the group ρ statistic, GC_ρ , based on the less restrictive group model, Equation 7,

$$GC_\rho = \frac{T}{\sqrt{N}} \sum_{i=1}^N \left(\sum_{t=1}^T \hat{e}_{i,t-1}^2 \right)^{-1} \sum_{t=1}^T (\hat{e}_{i,t-1} \Delta \hat{e}_{i,t} - \hat{\lambda}_i) \tag{C.10}$$

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 $\hat{u}_{i,t} = \hat{e}_{i,t} - \hat{\gamma}_i \hat{e}_{i,t-1}$. $\hat{L}_{lli}^{-2} = \hat{\Omega}_{21,i} \hat{\Omega}_{22,i}^{-1} \hat{\Omega}'_{21,i}$ is the multivariate estimate of the long-run covariance matrix, Ω_γ .



© 2014 The Author(s). This open access article is distributed under a Creative Commons Attribution (CC-BY) 3.0 license.

You are free to:

Share — copy and redistribute the material in any medium or format

Adapt — remix, transform, and build upon the material for any purpose, even commercially.

The licensor cannot revoke these freedoms as long as you follow the license terms.

Under the following terms:

Attribution — You must give appropriate credit, provide a link to the license, and indicate if changes were made.

You may do so in any reasonable manner, but not in any way that suggests the licensor endorses you or your use.

No additional restrictions

You may not apply legal terms or technological measures that legally restrict others from doing anything the license permits.



***Cogent Economics & Finance* (ISSN: 2332-2039) is published by Cogent OA, part of Taylor & Francis Group.**

Publishing with Cogent OA ensures:

- Immediate, universal access to your article on publication
- High visibility and discoverability via the Cogent OA website as well as Taylor & Francis Online
- Download and citation statistics for your article
- Rapid online publication
- Input from, and dialog with, expert editors and editorial boards
- Retention of full copyright of your article
- Guaranteed legacy preservation of your article
- Discounts and waivers for authors in developing regions

Submit your manuscript to a Cogent OA journal at www.CogentOA.com

