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**Testing the Balassa-Samuelson Hypothesis: Evidence
from 10 OECD Countries**

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ABSTRACT

This study tests the Balassa-Samuelson (BS) hypothesis for 10 OECD countries between 1975 and 2007 using the recent data sets and econometric methods. The study employs the Johansen cointegration approach in country-specific analysis. To investigate the existence of cointegration in panel series Pedroni, Kao and Johansen-Fisher tests of panel cointegration are used. The country-specific and panel cointegration test results confirm the existence of cointegration among the real effective exchange rate, relative productivity and terms of trade. And the country-specific coefficient estimations put evidence in favor of the BS hypothesis, except USA, and Japan. On the other hand, the panel estimation results could not confirm the validity of the BS hypothesis due to wrong sign of the BS effect. In a nutshell, the study's major findings suggest that the BS hypothesis still works well in OECD countries while explaining the long-run movements of the real effective exchange rates.

Key Words: Balassa-Samuelson Hypothesis, Real Effective Exchange Rate, Johansen Cointegration, Panel Cointegration.

To My Family

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LIST OF ABBREVIATIONS AND ACRONYMS

ADF	Augmented Dickey-Fuller
ARDL	Autoregressive Distributed Lag
BS	Balassa-Samuelson
CPI	Consumer Price Index
DOLS	Dynamic Ordinary Least Squares
ECM	Error Correction Mechanism
Eq.	Equation
EU	European Union
FMOLS	Fully modified Ordinary Least Squares
GDP	Gross Domestic Product
IC	Information Criterion
IFS	International Financial Statistics
IMF	International Monetary Fund
IPS	Im-Peseran-Shin
OECD	Organisation for Economic Co-operation and Development
PPP	Purchasing Power Parity
PRO	Relative Productivity
REER	Real Effective Exchange Rate
RER	Real Exchange Rate
TOT	Terms of Trade
UK	United Kingdom
USA	United States of America
VAR	Vector Autoregression
WDI	World Development Indicators

“The only relevant test of the validity of a hypothesis is comparison of its predictions with experience.”

Milton Friedman, 1953, p.8

1. INTRODUCTION

In 1920, Gustav Cassel, a famous Swedish economist proposed the Purchasing Power Parity (PPP) theorem which based on *law of one price*, and states that in the long-run exchange rates should be identical across countries. In other words, “the PPP theory predicts that, in the long-run, relative prices determine the exchange rate (i.e., $e=P/P^*$); and any deviation of relative prices from the equilibrium exchange rate will be transient and ultimately mean-reverting in the long-run” (Chowdhury, 2007:p.4). Nonetheless, empirical studies rejected this version of (absolute) PPP theorem.¹ Apart from reasons such as transaction costs, transportation costs and inefficient markets, the most convincing explanation came from Balassa (1964) and Samuelson (1964) -known as the Balassa-Samuelson hypothesis- which states that the productivity differences in tradables and nontradables sectors across countries lead to differentiation of wages, price levels and, hence the real exchange rates.

In particular, the BS hypothesis explains the two effects:

(i) The price level differences across countries (The Penn Effect) : According to the BS hypothesis, when the productivity level of tradables sector in home increases relative to foreign country’s tradables sector, home experiences higher price level due to increase in general wage level.² This is known as the Penn Effect, which is an explanation of the high price levels in rich (high per capita income) countries.

(ii) The real exchange rate differences across countries (The BS effect): The BS hypothesis claims that in a country where the productivity of tradables sector is higher than the other country, then the real exchange rate index (R) of this country will be higher. This increase in the real exchange rate index -which is known as real appreciation- stems from the definition of the real exchange rate in which price level of home (P) stands in the numerator and price level of foreign country (P^*) and nominal exchange rate (e) stand in the denominator.

¹ Antweiler (2008) and OECD (2005) give clear definitions of PPP. And see Pilbeam (2006:pp.135-139) for empirical studies which rejected the PPP theorem.

² Assume that the productivity levels of nontradables sectors are identical in two countries.

($R = P/e.P^*$). Throughout the study, we concentrate on and test the second effect of the BS hypothesis regarding the real exchange rate differences across countries.

The BS hypothesis, which explains the movements of the real exchange rates with changes in productivity levels across countries, became popular in the literature of empirical economics in the sense that it opened a new research avenue for the researchers who want to understand the failure of the PPP theorem. The main strength of the BS hypothesis stems from its validity when it is tested by different empirical methods. In this regard, the BS hypothesis has been used since 1964 in order to explain the long-run movements of the real exchange rates across countries.³

This study aims to test the BS hypothesis for 10 OECD countries between 1975 and 2007 using the recent data sets and econometric methods. The study differs from the others in four respects. First, the paper tests the BS hypothesis not only for individual countries but also for the panel (10 countries). Second, the OECD average of labor productivity is used as the benchmark to calculate the relative productivities. Third, in addition to the relative productivity explanatory variable, the “terms of trade” added as the second explanatory variable, which can be seen as a test of “the extended (unrestricted) BS model”.⁴ Fourth, the study employs the recent econometric tests and methods which are the Johansen cointegration, panel unit root tests of Im-Pesaran-Shin and Breitung, and three different panel cointegration tests of Pedroni, Kao and Johansen-Fisher.

The main conclusion of the study is that the BS hypothesis still keeps its importance in explaining the real exchange rate movements across OECD countries. The findings of the study mostly confirm the validity of the BS hypothesis, even though the results are country and model specific. All cointegration tests verify the long-run relation among the real effective exchange rate, relative productivity and terms of trade. In particular, the country-specific estimations indicate that the BS hypothesis is valid for 7 countries out of 10 when the terms of trade variable excluded (the original BS model). When the terms of trade included

³ Some authors take the beginning of the hypothesis as 1933 due to the study of Harrod (1933), and these authors prefer using the name of “Harrod-Balassa-Samuelson Hypothesis”. See for example Obstfeld and Rogoff (1996: p.210). However, throughout this study we prefer calling the “Balassa-Samuelson Hypothesis” by following the common convention. See the discussion in Tica and Druzic (2006:p.5).

⁴ See Alexius and Nilsson (2000) for a use of similar name.

(the extended BS model), the BS hypothesis is rejected only for 2 countries out of 10. In addition, the panel data analysis section of the study gives mixed results regarding the validity of the BS hypothesis. Although the cointegration relation has been found among the panel series of real effective exchange rate, relative productivity and terms of trade, the panel estimation of the BS model generated the wrong sign (negative) for the coefficient of relative productivity variable (so-called the BS effect).

The organization of the study is as follows. Section 2 gives a brief literature review in comparative perspective paired with a summary literature results table. Section 3 revisits the BS model formally in which the deterministic and empirical BS models are derived, and assumptions are explained. Section 4 describes the data sources and transformation and, analyses the data with three figures. Section 5 presents the unit root tests, cointegration tests, and estimation results both for individual countries and panel under two separate main subsections. Section 6 reviews the main findings of the study and concludes.

2. LITERATURE REVIEW

2.1 Some Factors Behind the Popularity of the Balassa-Samuelson Hypothesis

After two separate papers of Balassa and Samuelson in the same year (1964), in which they explain the real exchange rate differentials across countries with productivity differences, the popularity of the BS hypothesis has increased over time. According to a survey conducted by Tica and Druzic (2006:p.4); “In total, since it was (re)discovered in 1964, the theory has been tested 58 times in 98 countries in time series or panel analyses and in 142 countries in cross-country analyses. In these estimates, country-specific BS coefficients have been estimated 164 times in total, and at least once for 65 different countries”.

Some of the main reasons can be counted as follows why the BS hypothesis preserves its importance and popularity in empirical economics:

- (a) The economic importance of finding the determinants of real exchange rates. And the continuing desire of researchers to explain why *law of one price* fails across countries.
- (b) The desire of researchers to explain the high price levels in developed countries.
- (c) The invention of new econometric techniques and easy implementation of the techniques via new econometric software programs. This enables researchers to test the BS hypothesis by using various kinds of models and techniques.

(d) The theoretical contributions to the theorem by adding additional variables such as terms of trade, oil prices and openness.⁵

(e) The availability of new data sets throughout the time, especially the sectoral productivity databases enable scholars to test the BS hypothesis without assuming all sectors have the same productivity level within a country.

In addition to these global factors Tica and Druzic (2006:p.4) claims that “the enlargement process of the EU” has promoted research on the BS model in the sense that price level differences across countries in the era of passing to a common currency (Euro) raised the reputation of the BS model among researchers in EU region.

2.2 Some Benchmark Studies in the Literature of the BS Hypothesis

In addition to the previous survey study of Froot and Rogoff (1994), a recent survey study of the BS hypothesis was prepared by Tica and Druzic in 2006. This survey enables us to see not only the all results of previous studies in a nutshell, but also explains the evolution of the BS hypothesis in terms of theoretical and econometric methods. Furthermore, the survey gives us a chance to compare our findings both with recent and previous studies’ results.

In here, we discuss some selected empirical studies regarding the BS hypothesis in which authors used similar methods and variables with us. Then, we present the summary findings of some selected studies in Table 1.

In his pioneering study, Balassa (1964) employed OLS analysis in order to estimate the equation in which real exchange rates were used as the dependent variable and per capita income levels as the independent ones. Froot and Rogoff (1994:p.32) notes that “Balassa (1964) reports a regression for a cross-section of twelve industrial countries for the year 1960 in which the estimated BS effect was 0.51 with a positive intercept term.” This result implies that 1% increase in per capita income levels lead to 0.51% increase in real exchange rate levels.

⁵ See for example Rogoff (1992), and De Gregorio & Wolf (1994).

According to Tica and Druzic (2006:p.6); in addition to Rogoff (1992), and Obstfeld & Rogoff (1996:pp.214-216) the most important theoretical contributions to the BS hypothesis came from De Gregorio, *et.al* (1994), and Asea & Mendoza (1994). After dozens of published papers, in 1994 De Gregorio and Wolf integrated the “terms of trade” formally into the BS model. In their influential study, they develop a simple model of a small open economy producing exportable and nontradable goods and consuming importable and nontradable goods, and present empirical evidence for a sample of fourteen OECD countries. Clearly, they conclude that “The evidence from OECD countries broadly supports the predictions of the model, namely that faster productivity growth in the tradable relative to the nontradable sector and an improvement in the terms of trade induces a real appreciation.” (De Gregorio and Wolf (1994:p.i).⁶

In a benchmark article for our study, Alexius and Nilsson (2000) use the terms of trade and relative real GDP (as a proxy of productivity) to explain the real exchange rate movements in 15 OECD countries from 1960 to 1996. They use the Johansen cointegration approach in search of cointegration relation among variables and estimate the BS model by using FMOLS method. And they report the presence of cointegration among three series in all countries and estimate the correct sign (positive) for the BS effect in two thirds of the cases.

Egert (2002) investigates whether the BS effect holds for the Czech Republic, Hungary, Poland, Slovakia and Slovenia during the transition process by employing the Johansen cointegration approach. Egert (2002) uses the relative labor productivity in the industrial sector as the independent variable and the “real effective exchange rate” as the dependent one. In the paper, he clearly concludes that the results are in favor of the BS model. In addition, he notices that real appreciation has not been backed by productivity increases in all countries such as in Slovakia and Czech Republic. Apart from Egert (2002), some other important studies in which real effective exchange rate used as the dependent variable are Simon and Kovacs (1998), Fischer (2002), Drine and Rault (2005).

After Pedroni’s influential papers in 1996 and 1999 regarding the panel cointegration methods, testing the BS hypothesis by employing the Pedroni cointegration method has

⁶ Also Alexius and Nilsson (2000:p.383) accept that, De Gregorio and Wolf (1994) extended the BS model by including the terms of trade.

become popular in the literature. Drine and Rault (2005) use the Pedroni cointegration method for panel data and the Johansen cointegration method for time series data in order to test the BS hypothesis for 12 OECD countries. And they find out different results for time series and panel data. They conclude that “Whereas standard time series approach turns out to be unable to put in evidence a significant long-run relationship is largely accepted for all countries using recent advances in the econometrics of non-stationary dynamic panels methods. This result doesn’t mean however that the BS is uniformly supported by data for all OECD countries, since actually four of them (Australia, Belgium, Canada and USA) are proved not to follow the BS path.”

Candelon, *et al.* (2007) employ new panel-cointegration techniques such as Kao paired with DOLS estimation method to estimate the long-run determinants of real exchange rates in eight new EU member states, for 1993–2003 period. They use a wide range of regressors list consists of openness, government consumption, total consumption, private consumption and productivity. And mostly their findings are in favor of the BS hypothesis in which the estimated BS coefficient lies in the 0.64 - 0.85 range.

Last but not least, Choudhri and Khan (2005) examines the BS hypothesis, in which terms of trade together with non-traded and traded goods productivity differentials are used as the explanatory variables, to explain the real exchange rate movements of 16 countries in 1976-1994 period by employing DOLS method. According to the authors, the results of the study provide strong verification of the BS effects for developing countries.

Table 1. Summary of Some Selected Studies Regarding the BS Hypothesis

<u>Authors</u>	<u>Dependent Variables</u>	<u>Independent Variables</u>	<u>Method</u>	<u>Estimated BS effect</u>
Balassa (1964)	RER	Per capita income	OLS	0.51
Alexius and Nilsson (2000)	RER	Domestic and foreign GDP, terms of trade	Johansen cointegration and FMOLS	between -1 and 1.9
Egert (2002)	REER	Labor productivity growth	VAR, Johansen cointegration	between -0.3 and 2.4
De Broeck and Slok (2001)	REER	Productivity, openness, terms of trade	Pooled mean estimation	between 0 and 3.5
Choudhri and Khan (2005)	RER	Labor productivity differences, terms of trade	DOLS	between 0.9 and 1.2
Drine and Rault (2005)	REER, Price levels	Per capita GDP, productivity differences	VAR-ECM, Pedroni cointegration	between 0.6 and 1.5

3. REVISITING THE BALASSA-SAMUELSON MODEL

In this section, the BS model is presented formally and an empirical BS model is derived to use in our estimations. After Balassa (1964) and Samuelson (1964) and with theoretical contribution of De Gregorio and Wolf (1994) regarding the “terms of trade”, today widely accepted the BS model can be formalized to explain the real exchange rates as follows:

$$RER = f(\text{PRO}, \text{TOT}) \dots\dots\dots(1)$$

Simply, the BS hypothesis predicts that PRO and TOT variables are assumed to have a positive effect on the real exchange rate. In other words, an increase in productivity level and terms of trade relative to the numeraire country matched with an increase in the real exchange rate level of home country.⁷

3.1 The Basic Framework of the Balassa-Samuelson Model

Let there be 2-country and 2-sector world in which T stands for tradables and N stands for non-tradables sector. And * (asterisks) denotes the foreign country. Assume that labor (L) is immobile internationally whilst perfectly mobile within the country. And capital (K) is perfectly mobile internationally. Non-tradables sector produces the goods in a country which cannot be traded internationally. In contrast, tradables sector produces the goods which can be traded within country and internationally. In this framework, international trade assumed to be equalize prices of traded goods in two countries that P_T is normalized to 1, which necessarily implies $e = 1$ and $P_T^* = P_T = 1$.

Recall the real exchange rate definition;

$$RER = P / eP^* \dots\dots\dots(2)$$

In where, $P = P_T^\alpha \times P_N^{1-\alpha}$ (general price level equation in home country).....(3)

$$P^* = P_T^{*\alpha} \times P_N^{*1-\alpha}$$
 (general price level equation in foreign country).....(4)

(e = Nominal exchange rate, defined as the price of the domestic currency in terms of the foreign one.)

⁷ When the domestic currency per foreign one, nominal exchange rate definition (say, 8 SEK / 1\$) is used in calculation of the real exchange rate.

If we divide (3) with (4) and apply the assumption $P_T^* = P_T = 1$, equation (2) can be re-written as ⁸:
$$RER = \frac{P_N^{1-\alpha}}{P_N^*{}^{1-\alpha}} \dots\dots\dots(5)$$

$$Y = Y_T + Y_N; Y_T = A_T F(K_T, L_T); Y_N = A_N F(K_N, L_N) \text{ (production functions in home) } \dots\dots(6)$$

$$Y^* = Y_T^* + Y_N^*; Y_T^* = A_T^* G(K_T^*, L_T^*); Y_N^* = A_N^* G(K_N^*, L_N^*) \text{ (production functions in foreign) } \dots(7)$$

Now we turn to the production side of the economy, in the long-run as an effect of “perfect labor mobility” within the country, the real wage levels are identical in T and N sectors (i.e., $W = W_T/P = W_N/P$). And by using the first order conditions of equations (6) and (7) in profit maximization context, one may show that the marginal product of capital equals to the world interest rate (factor price of K), and the marginal product of labor equals to the nominal wage (factor price of L) in two sectors.⁹ That is to say, the nominal wages (W_T and W_N) are equalized across sectors within the country. Hence, we can write;

$$W = P_T \times A_T = P_N \times A_N ; \text{ since } P_T = 1 ; \rightarrow P_N = A_T / A_N \dots\dots\dots(8)$$

$$W^* = P_T^* \times A_T^* = P_N^* \times A_N^* ; \text{ since } P_T^* = 1 ; \rightarrow P_N^* = A_T^* / A_N^* \dots\dots\dots(9)$$

Under the assumption of “the shares of labor income in non-tradables and tradables sectors are equal”¹⁰; plugging (8) and (9) into equation (5) and taking the natural logarithms yield the deterministic BS model equation:

$$\ln RER = (1 - \alpha)[(\ln A_T - \ln A_T^*) - (\ln A_N - \ln A_N^*)] \dots\dots\dots(10)$$

With the assumption that $A_N = A_N^*$, then non-tradables sector vanishes from equation (10) and it becomes;

$$\ln RER = (1 - \alpha)[(\ln A_T - \ln A_T^*)] \dots\dots\dots(11)$$

⁸ In other words, “in a two-good world, the relative price level (the real exchange rate) between the countries depends solely on the relative price of non-tradables” (Muscatelli *et al.*, 2007:p.1405).

⁹ For a clear exposition of the derivations see for example, Dumitru and Jianu (2009:pp.885-886). And for a similar derivation of the BS model see Muscatelli, *et al.* (2007:pp.1405-1406), and Lothian & Taylor (2008:pp. 1745-1747).

¹⁰ If we do not make this assumption there should be an additional ratio in front of $(\ln A_T - \ln A_T^*)$ in the square brackets, which represents the shares of the labor income in non-tradables and tradables sectors. Obstfeld and Rogoff (1996:pp.208-212) explicitly shows the derivation of this type of equation. However, they also use the equation (10) in the sense that empirics showed that this ratio is almost one.

In equation (11), the coefficient $(1 - \alpha)$ is known as the BS effect and theoretically assumed to be positive. Thus the BS hypothesis suggests that, in home country an increase in the productivity of tradables relative to foreign country should associate with an increase in the real exchange rate level of home country. Put differently home country, whose productivity growth rate is higher than the foreign one in tradables, firstly experiences an increase in wage levels in both sectors due to factor price equalization across sectors within the country. Then the general price level, say CPI, goes up in home due to high wages level. Consequently, from the definition of the real exchange rate (eq.2), relatively high price level in home leads to an increase in RER, which is known as real appreciation.

3.2 The Empirical Balassa- Samuelson Model

In order to use the deterministic BS model (eq.11) in our estimations, we need to convert it into an empirical one. Before converting it into an empirical model, it is beneficial to explain the assumption that we make in equation (11) that “there is no productivity growth in non-tradables sector over time” ($A_N = A_N^*$). Although it seems a strong assumption, in empirical studies it has been widely employed.¹¹ Moreover, many studies such as Balassa (1964), Alexius and Nilsson (2000), which use relative GDP as a proxy of relative productivity of countries implicitly make this assumption. The rationale behind this assumption is two-fold for our study. First, there is no long time series data of sectoral productivity for cross-section analysis. Second, there is no consensus among scholars today, about the classification of some sectors whether they are tradables or not. In this respect, for the robustness of the BS hypothesis test of 10 countries, the assumption we make seems plausible in point of view of theory. In fact, this assumption implies that relative to tradables sector, non-tradables sector’s productivity growth rate can be negligible. As can be seen from Table 2, between 1995 and 2003, the annual changes in labor productivity index in manufacturing sector for 5 selected countries are always significantly higher than the annual changes in services sector. In this regard, the values in Table 2 support our simplifying assumption.

¹¹ Egert (2002) and Lothian & Taylor (2008) use the same assumption.

Table 2. Labor Productivity Indices for Selected Countries and Years

	Labor Productivity Indices					
	in Manufacturing			in Services		
	1995	2003	Average annual Change (%)	1995	2003	Average annual Change (%)
Canada	95	115	2.63	98	108	1.28
Denmark	100	123	2.88	100	109	1.13
Japan	100	138	4.75	100	106	0.75
Sweden	70	111	7.32	92	100	1.09
USA	77	120	6.98	91	105	1.92

Source: OECD_STAN Indicators Database ed.2005.

And if we write equation (11) as a stochastic model with an intercept term in natural logarithmic form and eliminate the T (sectoral) subscript, we get the basic empirical BS model ¹²;

$$\ln REER_{i,t} = \delta + \beta_1 * \ln \left(\frac{A_{i,t}}{A_t^*} \right) + u_{i,t} \dots \dots \dots (12)$$

And as we mentioned in section 2, especially after the study of De Gregorio and Wolf (1994) the terms of trade has been used widely as a second independent variable in the BS model.¹³ Thus, we add the terms of trade to equation (12) and get equation (13).

$$\ln REER_{i,t} = \delta + \beta_1 * \ln \left(\frac{A_{i,t}}{A_t^*} \right) + \beta_2 * \ln(TOT_{i,t}) + u_{i,t} \dots \dots \dots (13.i)$$

If we denote the relative productivity $\left(\frac{A_{i,t}}{A_t^*} \right)$ as “PRO”, simply we can write,

$$\ln REER_{i,t} = \delta + \beta_1 * \ln(PRO_{i,t}) + \beta_2 * \ln(TOT_{i,t}) + u_{i,t} \dots \dots \dots (13.ii)$$

In our estimations, the derived empirical BS model (eq.13.ii), which includes the terms of trade, is used as the benchmark equation (unrestricted model) and the model without the terms of trade (eq.12) is used as the restricted BS model.

¹² Note that we use REER in eq.12 instead of RER, since we use REER data in model estimations. See the next page for the definition of REER.

¹³ See for example, Alexius and Nilsson (2000), Choudhri & Khan (2005), and Jaunky (2007).

Before proceeding to data description section, it is crucial to define the variables and coefficients of the equations (12) and (13).

The definitions of the variables and expectations about the coefficients are as follows:

REER_{it} : According to IMF, “the real effective exchange rate is computed as the weighted geometric average of the price of the domestic country relative to the prices of its trade partners”. And CPI-based REER can be expressed as ¹⁴ :

$$REER = \prod_{j \neq i} \left[\frac{P_i R_i}{P_j R_j} \right]^{W_{ij}} \dots \dots \dots (14)$$

In where ; **P_i** : home country’s CPI index, **R_i** : nominal exchange rate of home country in US dollars, **P_j** : price index of country j (foreign), **R_j** : nominal exchange rate of country j’s currency in US dollars, **W_{ij}**: country j’s weight for home.

In this type of REER definition, an increase in the index denotes a real appreciation of the home currency, whereas a decrease implies a real depreciation (see e.g. CBRT (2008)).

PRO_{it} : Relative productivity, $\left(\frac{A_{i,t}}{A_t^*}\right)$ in where;

A_{it} : Labor productivity in manufacturing sector (tradables) in country i.

A_t^{*} : Unweighted mean of “Labor productivity in manufacturing sector (tradables) of 14 OECD countries”, simply “average productivity of OECD”.

TOT_{it} : Terms of trade (relative prices of country i’s export to import).

β₁ : The Balassa- Samuelson (BS) effect, theoretical *expected sign is positive*.

β₂ : The terms of trade (TOT) effect, theoretical *expected sign is positive*.

i : Canada, Denmark, Italy, Germany, Japan, Netherlands, Norway, Sweden, UK, USA.

t : 1975, 1976...2007.

¹⁴ For a detailed exposition of calculations and weights; see IMF (2008), CBRT (2008), Zanello and Dominique (1997).

The β_1 and β_2 coefficients can be interpreted as the elasticity of the PRO and TOT with respect to REER, since the independent and dependent variables are in natural logarithmic forms. Simply, the expected positive signs of the coefficients point out that; an (%) increase in “the relative productivity” and an (%) increase in “the terms of trade” in country i , should associate with an (%) increase in the real effective exchange rate. Thus, the main goal of the estimations in section 5 is to find out the sign and size of these coefficients for 10 countries to test the validity of the BS hypothesis.

4. DATA

4.1 Sources and Description of Data

To test the BS hypothesis, 10 OECD countries are selected for the 1975-2007 period. In other words, the panel dataset is constructed with 10 cross-section units and 33-year sample. Cross section units (countries) and data length are selected according to data availability. Nonetheless, the selected 10 countries have similar properties in the sense that all of them are OECD countries and developed countries in terms of per capita income by the end of 2007.

The real effective exchange rate data are extracted from World Development Indicators. WDI uses the IMF-IFS statistics database as the main source to extract the CPI-based real effective exchange rate data. More precisely, our CPI-based REER data set is taken from the IMF-IFS via WDI database which uses 2000 as the base year (2000=100). As mentioned in section 3, “the real effective exchange rate is computed as the weighted geometric average of the price of the domestic country relative to the prices of its trade partners”. In our model, the usage of “real effective exchange rate” data instead of “real exchange rate” data may better reflect the changes in relative productivity. Because, we do not use single-country labor productivity as a benchmark, instead we use the OECD average to calculate the relative productivity, as explained below. Put differently, the usage of the relative (OECD) productivity together with REER may increase the explanatory power of the BS model and robustness of the estimation results.

“Labor productivity (output per employed person) in manufacturing sector” is chosen as the proxy of productivity in tradables sector. It is worth mentioning that there is a long lasting debate about the best productivity proxy issue in literature. Although “labor productivity” has

been used in empirical studies widely, as mentioned in section 2, some authors use GDP as a proxy and some argue that total factor productivity (TFP) is a better proxy than the labor productivity. However, it is always mentioned that TFP is difficult to calculate for different type of sectors and countries, moreover comparable long time series data are not available for our study.¹⁵

The labor productivity data in manufacturing sector are taken from the publication of “U.S. Department of Labor, Bureau of Labor Statistics” in March 2009 which uses 1996 as a base year (1996=100). Before the calculation of the relative productivity data, we converted the base year 1996 into 2000 for consistency with two other variables’ base years. And to construct the relative labor productivity data the “unweighted mean of labor productivity of 14 OECD countries” is used as the benchmark.¹⁶ For example, to calculate the relative productivity ($A_{i,t}/A_t^*$) for Canada in 1980; first we calculate the “unweighted mean of labor productivity in manufacturing sector of 14 OECD countries” in 1980. Then Canada’s labor productivity index in 1980 is divided by this mean value to get the “PRO” (the relative productivity) data for Canada.

The second explanatory variable of the model “terms of trade” data are calculated from the IMF-IFS data base. The terms of trade is defined as the relative prices of a country's export to import. In this regard, export and import unit prices of 10 countries are extracted from the IMF-IFS data base in which 2000 is used as the base year (2000=100). By taking the ratio of two prices, we get the raw terms of trade data. Then, we multiply it by 100 in order to make the scale of the TOT index 100, as the REER and relative productivity indices.

4.2 Analysis of Data

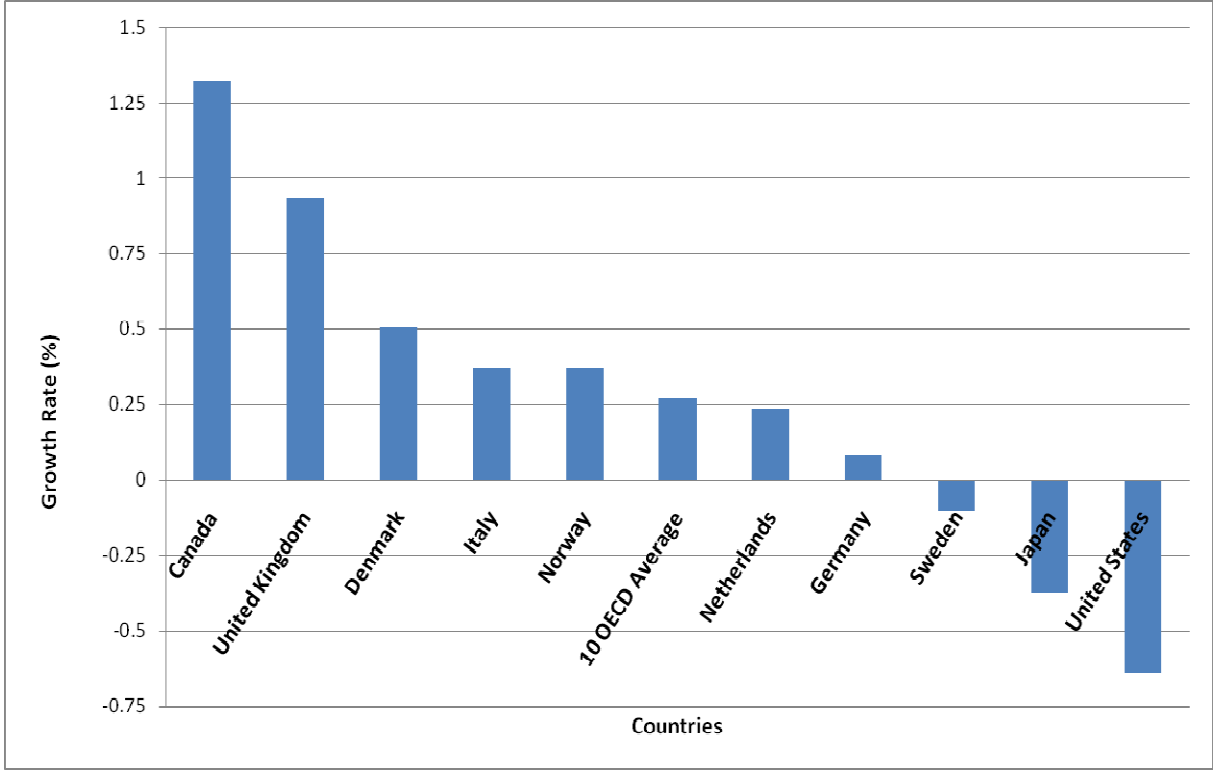
After the description of the variables and data, in here we present the figures of the average annual growth rates of three variables. In figures, we both present the country-specific results and “average of 10 OECD countries” in a descending order. By doing this, we can see the performance of each country’s relative to the average of 10 OECD countries.

¹⁵ See the discussion on this issue in Tica and Druzic (2006:p.11)

¹⁶ These 14 countries are Australia, Belgium, Canada, Denmark, France, Italy, Germany, Japan, Netherlands, Norway, Spain, Sweden, UK, and USA.

Figure 1 demonstrates the average annual REER growth rates of 10 OECD countries in 1975-2007 period. In this period, the average highest annual REER growth rate (1.3%) occurs in Canada, followed by UK, Denmark, Italy and Norway. These countries are above the “average of 10 OECD countries” which is 0.26 %. The average annual REER growth rates of Netherlands, Germany, Sweden, Japan and USA are below “the average of 10 OECD countries”. In particular, USA places at the bottom with -0.64 % average annual REER growth rate, which implies that USA experienced 0.64 % annual REER depreciation in 1975-2007 period.

**Figure 1. Average Annual REER Growth Rate (%) in 1975-2007 Period:
Evidence from 10 OECD Countries**

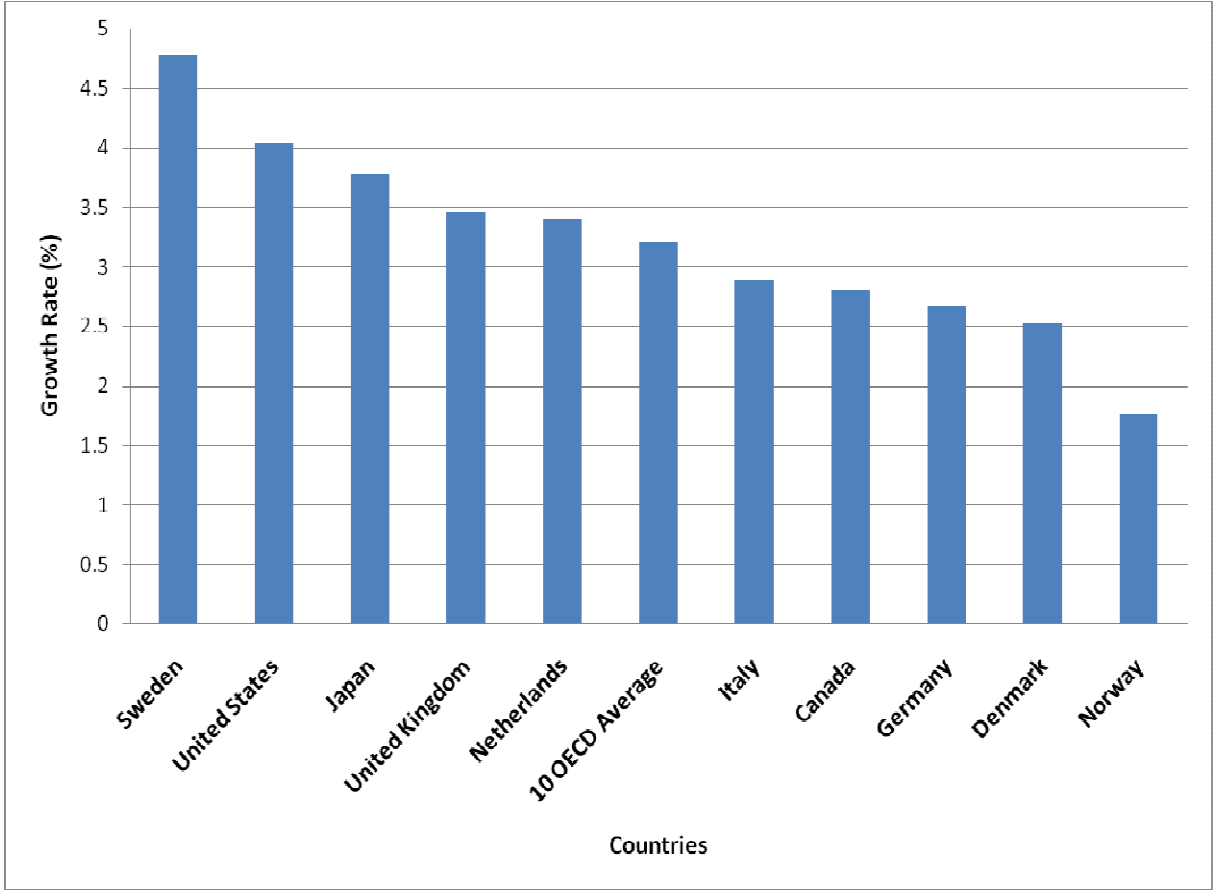


Source: World Development Indicators, The World Bank Group.

Figure 2 illustrates the average annual labor productivity growth rates of 10 OECD countries in 1975-2007 period. According to Figure 2, Sweden experienced the highest average annual labor productivity growth rate with 4.7 %. And USA, Japan, UK and Netherlands are the followers of Sweden in terms of the average annual labor productivity growth rate in 1975-2007 period. The “average of 10 OECD countries” is 3.2% in this period. Italy, Canada, Germany, Denmark and Norway are the relatively bad performer countries since their productivity growth rates are below the “average of 10 OECD countries”. Among 10

countries, with 1.75 % annual growth rate, Norway takes place at the bottom in terms of the average annual labor productivity growth rate.

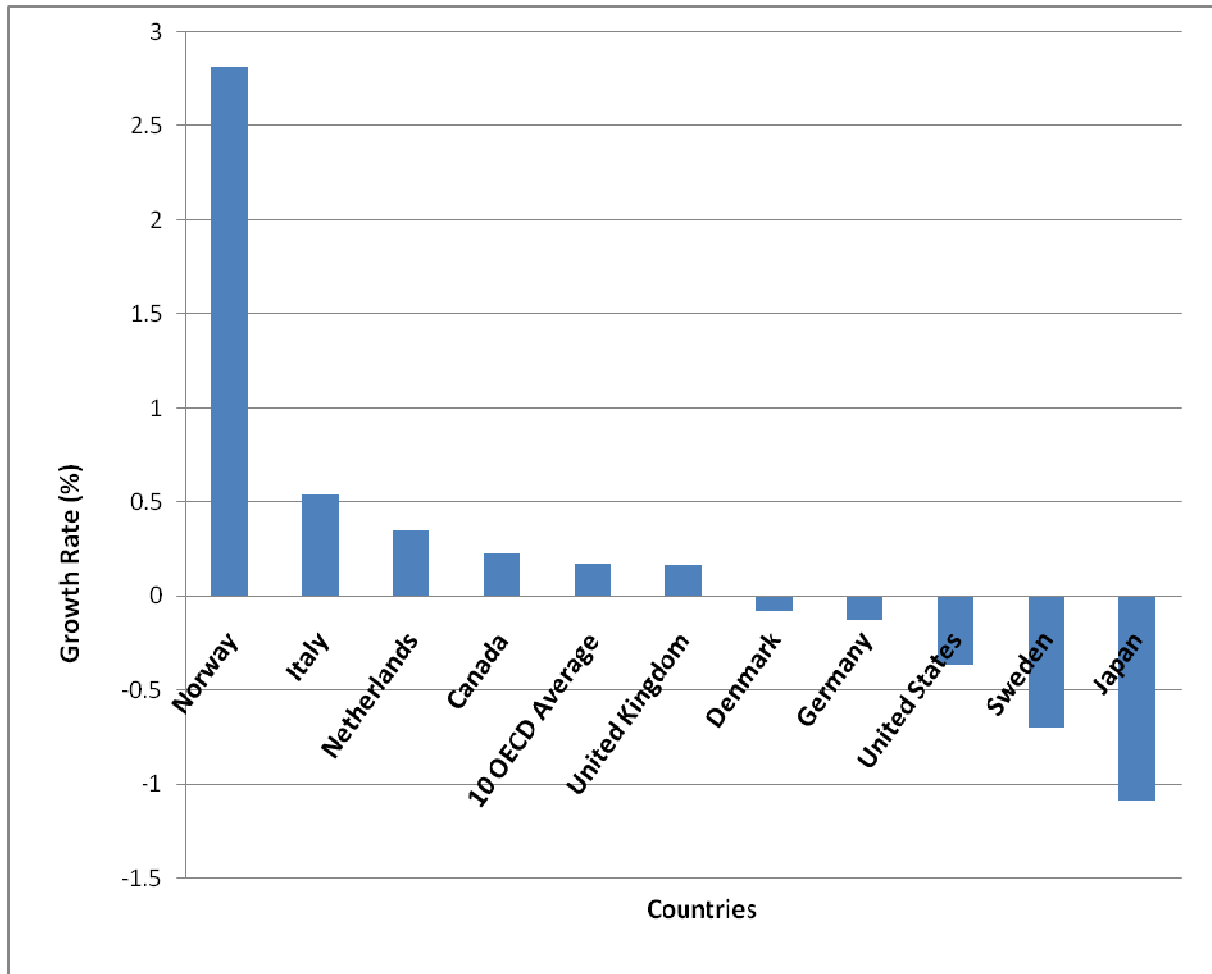
**Figure 2. Average Annual Labor Productivity Growth Rate (%) in 1975-2007 Period:
Evidence from 10 OECD Countries**



Source: U.S. Department of Labor, Bureau of Labor Statistics, 2009.

Figure 3 portrays the 10 OECD countries’ average annual TOT growth rates in 1975-2007 period. It can be revealed from Figure 3 that Norway is the outlier country among 10 countries with 2.8 % average annual TOT growth rate. The “average of 10 OECD countries” is 0.17 % in this period. Apart from Norway, Italy, Netherlands and Canada are placed above the average (0.17%). On the other hand Denmark, Germany, USA, Sweden, and Japan are the countries who experienced negative average annual TOT growth rates in this period. Among 10 countries, Japan is at the last rank with -1.08 % average annual TOT growth rate in 1975-2007 period.

**Figure 3. Average Annual TOT Growth Rate (%) in 1975-2007 Period:
Evidence from 10 OECD Countries**



Source: IMF-IFS Database.

Three basic results that can be revealed from the interpretation of the figures are as follows:

- Among 10 countries, Canada, UK, Denmark experienced significant annual REER appreciation in 1975-2007 period.
- Sweden, USA, Japan and UK are relatively successful countries in terms of the “average annual labor productivity growth rate” that they succeeded to hold their growth rates above the “average of 10 OECD countries”.
- TOT is a critical variable for Norway that it outweighs others in terms of the “average annual TOT growth rate”. TOT is also relatively important for Italy, Netherlands and Canada that their “average annual TOT growth rates” are above the “average of 10 OECD countries”.

5. METHODS AND ESTIMATION RESULTS

5.1 Methods

As mentioned in introduction, one of the distinguishing features of the study is the use of recent econometric tests and methods to test the BS hypothesis. Overall, we use eight different tests and methods to check the validity of the BS hypothesis in 10 OECD countries, as summarized in Table 3.

Table 3. Employed Tests and Methods

	Country-Specific Results	Panel Results
Unit Root Tests	ADF	IPS and Breitung
Cointegration Tests	Johansen	Pedroni, Kao, and Johansen-Fisher
Estimation Method	Johansen (Maximum Likelihood)	DOLS

After brief discussions on these methods in related sections, the results of these methods are presented in two sub-sections, which are country-specific and panel estimation sections. Before starting the cointegration tests and estimations, in section 5.2 we employ the country-specific and panel unit root tests in order to identify the stationarity of the series. Then, in sections 5.3 and 5.4, we follow a two-step approach to test the BS hypothesis for 10 OECD countries:

- First, we test whether or not; there is cointegration relation among REER, PRO and TOT variables.
- Second, we estimate the coefficients of the BS model (equations 12 and 13), and according to estimation results, we decide whether the BS hypothesis holds or not.

5.2 Unit Root Tests

A non-stationary series has a different mean at different points in time, and its variance increases with the sample size (see e.g. Harris and Sollis, 2005: p.29). In other words, it is not a mean reverting series; hence a shock (innovation) in the series does not die away. It is formulated as “non-stationary series have long memory”. In the econometrics literature, tests to detect the existence of stationarity (or non-stationarity) in the series are known as unit root tests. And if a series is becoming stationary after taking the first difference, it is known as integrated of order 1 or I (1).

In time-series econometrics, non-stationary series (unit root) phenomenon is crucial in the sense that linear combinations of these series generate spurious or nonsense regressions. In a spurious regression, although t values of the coefficients are significant, R-square and Durbin-Watson (d.w) values are low. More importantly, the estimated coefficients are biased. In this respect, detecting the existence of unit root in time-series becomes vital in order to get rid of spurious regression risk.¹⁷ One way to escape from spurious regression in existence of non-stationary series is looking for a **cointegration** relation among variables. Cointegration defined as:

“Cointegration means that despite being individually nonstationary, a linear combination of two or more time series can be stationary”. “Economically speaking, two variables will be cointegrated if they have a long-term, or equilibrium, relationship between them.”

(Gujarati, 2003:p.822 and 830)

In the cointegration literature, the Engle–Granger (EG) two-step approach is a famous one in order to test the existence of cointegration if there are only two series. In case of more than two series, the Johansen approach is the most applicable one, moreover it has some distinct advantages on the EG approach such as it is not a two-step approach. We employ the Johansen approach to test the existence of cointegration in country-specific series since we have three series. It is worth mentioning that in both approaches the necessary condition to search for cointegration is that “integration order of series should be the same”. Simply, all series should be I (d).

5.2.1 Country-Specific Unit Root Test Results

Before cointegration tests, our first task is to conduct the unit root tests in order to understand whether the series are stationary or not. We use the Augmented Dickey–Fuller (ADF) unit root test for individual series. And for the panel series, we employ the Im-Peseran-Shin (IPS) and Breitung panel unit root tests.

The ADF unit root test results of country-specific series are documented in Table 4. The ADF test uses the null hypothesis that “series has a unit root”. For example, the unit root test for Canada’s REER series is conducted as follows:

¹⁷ See the discussion in Gujarati (2003:p.806-807) on spurious regressions.

H_0 : REER (Canada) has a unit root.

H_1 : REER (Canada) does not have a unit root (stationary).

For the level (the original values of series) we accept the null. Since, -1.53 is not significant at 5%. However if we take the first difference t-value becomes -5.57 which is significant at 5%, and we reject the null. Thus, we conclude that REER of Canada is I (1).

Table 4. ADF Unit Root Test Results of Country-Specific Series

	VARIABLES					
	REER		PRO		TOT	
	Level	First difference	Level	First difference	Level	First difference
CANADA	-1.535167 (0.7950)	-5.570857 (0.0004)	-3.025691 (0.1417)	-4.618982 (0.0045)	-1.499314 (0.8088)	-4.735821 (0.0048)
DENMARK	-2.167158 (0.4908)	-4.487358 (0.0062)	-2.412536 (0.3666)	-5.086615 (0.0015)	-2.194638 (0.4763)	-3.921323 (0.0235)
GERMANY	-3.117222 (0.1195)	-5.785665 (0.0002)	-0.141317 (0.9917)	-6.160361 (0.0001)	-2.216767 (0.4648)	-4.511697 (0.0058)
ITALY	-2.141961 (0.5040)	-4.864695 (0.0024)	0.739853* (0.9911)	-4.287496* (0.0021)	-1.119237* (0.6959)	-4.739137* (0.0006)
JAPAN	-1.966596* (0.2992)	-4.184688 (0.0027)	-3.356682 (0.0760)	-4.759665 (0.0032)	-1.152128 (0.9035)	-4.287677 (0.0099)
NETHERLANDS	-2.248193 (0.4485)	-5.341529 (0.0007)	-2.508787* (0.1229)	-5.780036* (0.0001)	-1.421048** (0.8351)	-4.559828** (0.0052)
NORWAY	-2.397041 (0.3740)	-5.479582 (0.0005)	-2.389405 (0.3777)	-4.694513 (0.0037)	-0.891747 (0.9448)	-4.762738 (0.0033)
SWEDEN	-2.688790 (0.2474)	-4.963406 (0.0019)	-2.875087 (0.1833)	-4.043640 (0.0175)	-1.287855 (0.8729)	-3.631978 (0.0432)
UK	-2.080568 (0.5366)	-4.680097 (0.0039)	-2.383519* (0.1544)	-3.094322* (0.0374)	-2.248376 (0.4484)	-4.528129 (0.0056)
USA	-3.161792 (0.1106)	-4.423442 (0.0072)	-1.868569 (0.6472)	-4.239519 (0.0118)	-2.476182 (0.3368)	-5.036180 (0.0016)
DECISION	REER is I(1) for all series		PRO is I(1) for all series		TOT is I(1) for all series	

Notes: (1) Values without parentheses are t-statistics. (2) Probabilities are in parentheses. (3) **Bold** numbers denote that they are significant at 5% level. (4) All tests are conducted by including “intercept and trend”, except series with (*). (5) Tests are conducted for series with (*) by including “intercept”. (6) Values with (**) are Phillips-Perron t-statistics. (7) Automatic lag length selection (Schwarz) is used with maximum 8 lags.

In a similar fashion, when we conduct the unit root tests for all country-specific series, we conclude that all series are I (1).¹⁸ That is to say, we can search cointegration relation among the series.

¹⁸ Only TOT series of Netherlands was found I (2) by the ADF test. However, the Phillips-Perron unit root test showed that the series is I (1). For consistency, we reported the Phillips-Perron test result for the TOT series of Netherlands that we marked the values with (**) in Table 4. In addition, the Elliot-Rothenberg-Stock unit root test also confirmed that this series is I (1).

5.2.2 Panel Unit Root Test Results

Now we switch to the panel unit root tests. We conduct two different panel unit root tests to test whether panel of series are stationary or not. The Im-Peseran-Shin (IPS) and Breitung panel unit root tests results are reported in Table 5. To test the unit root in panel series, equation (15) is used:

$$\Delta y_{it} = \rho^* y_{i,t} + \sum_{L=1}^{p_i} \vartheta_{iL} \Delta y_{i,t-L} + u_{it} \dots \dots \dots (15)$$

The IPS tests the following hypotheses, where $\rho^* = \rho - 1$:

- H₀: $\rho^* = 0$ for all i (Panel has a unit root)
- H₁: $\rho^* < 0$ for at least one i. (At least, one series in panel does not have a unit root.)

For example, to test whether the “panel series of REER” has a unit root or not, we test the significance of -0.63. Since it is not significant at 5% level, we conclude that the “panel series of REER” has a unit root. However, it becomes stationary when we take the first difference of the series that -9.92 is significant at 5%. Thus, we accept the alternative. In a similar fashion, if we conduct the IPS test for all three panel series, it is found out that they are I (1).

When we accept the alternative hypothesis in the IPS test, this implies that at least one country’s series is stationary; however, we cannot be sure whether all series are stationary in the panel. Due to the alternative hypothesis that the IPS test uses, it takes place in the class of “individual unit root process” tests. In order to test whether “all series in panel are stationary”, we conduct “Breitung common individual” unit root test. Breitung tests the following hypotheses ¹⁹:

- H₀: $\rho_i^* = 0$ for all i. (Panel has a unit root)
- H₁: $\rho_i^* < 0$ for all i. (Panel does not have a unit root.)

¹⁹ Note that, the IPS test is less restrictive than Breitung in the sense that it allows for individual unit root processes, so that ρ_i^* may vary across cross-sections (Harris and Sollis, 2005:pp.193-200). Also see E-views (2005:pp.534-540) and Im *et al.* (1997). In addition, see Baltagi and Kao (2000) for an extensive review of panel unit root and cointegration tests.

For example, to test the unit root in the “panel series of TOT” (for level) we test the significance of 2.84, since it is not significant at 5%, we accept the null. When we take the first difference of the “panel series of TOT”, Breitung t-statistic becomes -6.86 which is significant at 5%, thus we accept the alternative. If we proceed similarly, Breitung test results show that all panel series are I (1).

As seen in Table 5, all panel unit root (individual and common) tests indicate that the panels of the REER, PRO, and TOT series are I (1) that we can seek panel cointegration among them.

Table 5. Panel Unit Root Test Results

	VARIABLES					
	REER		PRO		TOT	
	Level	First difference	Level	First difference	Level	First difference
PANEL (IPS test)	-0.63357 (0.2632)	-9.92457 (0.0000)	-1.22878 (0.1096)	-8.94550 (0.0000)	0.90109 (0.8162)	-7.17785 (0.0000)
PANEL (Breitung test)	0.78349 (0.7833)	-6.75561 (0.0000)	2.54816 (0.9946)	-3.88411 (0.0001)	2.84036 (0.9977)	-6.86105 (0.0000)
DECISION	REER is I(1) for panel		PRO is I(1) for panel		TOT is I(1) for panel	

Notes: (1) Values without parentheses in IPS test are “IPS-W-stats”. (2) Values without parentheses in Breitung test are “Breitung t-stats”. (3) **Bold** numbers denote that they are significant at 5% level. (4) Probabilities are in parentheses. (5) Tests are conducted by including “intercept and trend”. (6) Automatic lag length selection (Schwarz) is used.

5.3 Country-Specific Cointegration and Estimation Results

Since all country-specific series are found I (1), we can seek cointegration relation among REER, PRO, and TOT variables for each country using the Johansen approach. The aim of the cointegration tests is to understand whether or not, there is a long-run relation among REER, PRO and TOT variables as the BS hypothesis suggests.

The Johansen Multivariate Cointegration Approach

After the influential papers of Johansen (1988) and Johansen & Juselius (1990), the Johansen approach has become popular in the cointegration literature. Johansen (1988) expresses the following VAR system:

$$\Delta Z_t = \Gamma_1 \Delta Z_{t-1} + \dots + \Gamma_K \Delta Z_{t-K+1} + \pi Z_{t-K} + \varepsilon_t \dots \dots \dots (16)$$

“ π contains information about the long-run relationships among the variables included in vector Z . Johansen (1988) demonstrates that π could be decomposed into the product of two $p \times r$ matrices, i.e., $\pi = \alpha\beta'$ where the elements of β matrix form the long-run cointegrating coefficients and the elements of α matrix form the adjustment parameters. Number of cointegrating vectors r is determined by the rank of π .” (Bahmani-Oskooee and Economidou, 2009:p.195).

Johansen (1988) estimates equation (16) by using maximum likelihood method and extracts trace and maximum eigenvalue (λ -max) statistics to determine the number of cointegrating vectors. In addition, Johansen (1988) also reports the π matrix which contains the normalized α and β coefficients. However, there are two important things to decide before running the equation (16) for our model. These are:

- (i) optimal lag length of the VAR system (K in eq.16),
- (ii) decision about whether or not to add a trend (deterministic or quadratic) component to equation (16) and trend assumption in level data.

The specification of equation (16) regarding (i) and (ii) significantly affects the coefficients in π matrix. Put differently, the Johansen approach is sensitive to the specification of VAR-equation.²⁰ In order to decide the optimal lag length in our VAR model, we employ the “information criterion (IC)” approach. By using this approach we constructed Table 6, in which optimal lags are selected by three different information criteria. In the literature, there is no consensus among scholars which information criterion (IC) is the best one.²¹ In this regard, in the last column we stated the “preferred lags” as a conclusion of our search process. “Preferred lags” are the lags which give the best results in terms of “rank of cointegration” and “significance of estimated coefficients” in π matrix. As seen in Table 6, sometimes the preferred lags are not determined by any three information criteria. This is

²⁰ Stock and Watson (1993) shows that Johansen’s cointegration test is sensitive to the lag lengths used in the VAR models. Ahking (2002) and Turner (2007) find out that the Johansen approach is sensitive to the specification of the deterministic terms (trends).

²¹ For example, Harris and Sollis (2005:p.117) suggests the HQC, Khim and Liew (2004) recommends AIC as the best information criterion in determining the optimal lag length.

not surprising, that we only reported three information criterion results.²² In addition, this approach is used in other studies such as Alexius and Nilsson (2000).

Table 6. Lag Length and Model Selection

	AIC	BIC	HQC	Preferred lags	Preferred Model
CANADA	4	1	4	5	5
DENMARK	4	1	2	4	3
GERMANY	5	1	5	4	4
ITALY	5	1	5	5	4
JAPAN	5	1	5	4	5
NETHERLANDS	1	1	1	5	3
NORWAY	1	1	1	5	3
SWEDEN	1	1	1	5	4
UK	2	1	1	2	4
USA	5	4	5	5	5

Notes: Lag length selection is conducted by using Gretl - VAR lag selection command, with k-max=5. **AIC:** Akaike criterion, **BIC:** Schwartz Bayesian criterion, and **HQC:** Hannan-Quinn criterion. “Model numbers” are the same as in E-views 5.1-under Johansen cointegration command-. **Model 3:** Linear trends in level data, not in VAR. **Model 4:** Level data and VAR have linear trends. **Model 5:** Quadratic trends in level data, and linear trend in VAR (the less restrictive model).

There is no a systematic approach to decide whether or not to include a trend component into VAR models such as the IC approach for optimal lag length. Thus, determining the trend component and its type in VAR models is more problematic. Some authors, such as Ahking (2002) and Cheong (2005) prefer reporting VAR model estimation results with trend and without trend to tackle with this problem. But in here, we follow a similar “search process”, as we did in “lag length selection”.

In E-views there are five different trend specification options under the Johansen cointegration approach. We tried all five specifications with “preferred lags” in country-specific VAR models and reported the best results in the last column of Table 6 according to:

- Significance of estimated coefficients in π matrix and trend component.
- Consistency between the “rank of cointegration” test results of trace statistics and max-eigenvalue.

²² For example, we did not report FPE (Final prediction error) criterion since it is not available in Gretl and E-views software packages.

In most cases, the model 5 yielded the best results and in others at least a linear trend added to the model to capture the dynamic structure of the series.

Using the preferred lags and chosen model specifications, we run the VAR models for each country. Table 7 presents the trace statistics and max-eigenvalue results that we use to determine the “rank of cointegration” among three variables. For example, to find the rank of cointegration among three series of Canada, we follow two steps and test the following hypotheses:

Step 1:

$H_0: r = 0$ (no cointegration) ; $H_1: r \leq 1$ (at most one cointegration relation)

Using the trace statistic’s critical value at 5%, we reject the null. In other words, 86.23 is too significant that its probability is 0; hence we accept the alternative hypothesis.

Step 2:

$H_0: r \leq 1$ (at most one cointegration relation) ; $H_1: r \leq 2$ (at most two cointegration relations)

Since the probability of trace statistic is 0.87 that implies 5.98 is not significant at 5%. Thus we accept the null hypothesis. So far, according to trace rank test the rank of cointegration among Canada’s three series is 1.

We reach the same conclusion, if we apply the same steps by using maximum eigenvalues.

Step 1:

$H_0: r = 0$ (no cointegration) ; $H_1: r \leq 1$ (at most one cointegration relation)

At 5%, 80.25 is significant that its probability is 0. Hence, we accept the alternative hypothesis.

Step 2:

$H_0: r \leq 1$ (at most one cointegration relation); $H_1: r \leq 2$ (at most two cointegration relations)

Since the probability of maximum eigenvalue is 0.91, that implies 4.76 is not significant at 5%, thus we accept the null. So far, according to maximum eigenvalue rank test the rank of cointegration among Canada’s three series is 1.

If we employ the same steps for 10 countries, we see that the rank of cointegration among three series ranges between 1 and 3. Although the results of the rank tests are country-specific, for each country at least one cointegration relation has been found. That is to say, there is cointegration (long-run relation) among REER, PRO, and TOT variables in each country.

Table 7. Test Results For Cointegrating Rank

	Unrestricted Cointegration Trace Rank Test				Unrestricted Cointegration Maximum Eigenvalue Rank Test			
	tr (r = 0)	tr (r ≤ 1)	tr (r ≤ 2)	Rank	λ (r = 0)	λ (r ≤ 1)	λ (r ≤ 2)	Rank
CANADA	86.23475* (0.000)	5.981138 (0.8703)	1.219398 (0.2695)	1	80.25361* (0.0000)	4.761740 (0.9193)	1.219398 (0.2695)	1
DENMARK	42.65031* (0.0010)	11.66304 (0.1738)	0.699328 (0.4030)	1	30.98727* (0.0015)	10.96371 (0.1560)	0.699328 (0.4030)	1
GERMANY	98.63518* (0.0000)	32.78245* (0.0059)	11.74183 (0.0671)	2	65.85273* (0.0000)	21.04062* (0.0285)	11.74183 (0.0671)	2
ITALY	113.1335* (0.0000)	63.17209* (0.0000)	26.99207* (0.0001)	3	49.96144* (0.0000)	36.18001* (0.0001)	26.99207* (0.0001)	3
JAPAN	50.39815* (0.0006)	22.95751* (0.0107)	2.168101 (0.1409)	2	27.44064* (0.0183)	20.78941* (0.0141)	2.168101 (0.1409)	2
NETHERLANDS	64.27498* (0.0000)	17.626* (0.0235)	1.08409 (0.2978)	2	46.64898* (0.0000)	16.54191* (0.0215)	1.08409 (0.2978)	2
NORWAY	73.1332* (0.0000)	22.21885* (0.0042)	0.817101 (0.3660)	2	50.91435* (0.0000)	21.40175* (0.0032)	0.817101 (0.3660)	2
SWEDEN	62.55111* (0.0002)	33.13274* (0.0052)	13.00705* (0.0414)	3	29.41837* (0.0161)	20.12569* (0.0390)	13.00705* (0.0414)	3
UK	44.77427* (0.0322)	18.70152 (0.2988)	6.547679 (0.3941)	1	26.07275* (0.0464)	12.15384 (0.4010)	6.547679 (0.3941)	1
USA	55.87086* (0.0016)	27.62224* (0.0300)	3.961498 (0.7477)	2	28.24862* (0.0235)	23.66074* (0.0112)	3.961498 (0.7477)	2

Notes: (1) Values in parentheses are probabilities. (2) The null hypotheses are $r = 0$, $r \leq 1$, $r \leq 2$. (3) * denotes the rejection of the null at 5% level. (4) There is full consistency between rank test results of trace and maximum eigenvalue.

Now, the remaining task is to find the size and sign of this long-run relation among three variables. We do this by reporting the β elements of π matrix. If we take the REER as the dependent variable, the estimated normalized β coefficients show the size and sign of the relation among REER, PRO and TOT. To this end, we present Table 8, in which estimation results of the unrestricted (equation 13) and restricted models (equation 12) are reported. By interpreting the significance, size and sign of the normalized β coefficients, we decide whether the BS hypothesis holds or not. Moreover, by excluding the TOT variable from the unrestricted model, we can see whether it is important for the BS model.

Table 8. Country- Specific Estimation Results

	UNRESTRICTED MODEL (eq.13)			RESTRICTED MODEL (Restriction $\beta_2 = 0$) (eq.12)		
	REER	β_1 (PRO)	β_2 (TOT)	REER	β_1 (PRO)	β_2 (TOT)
CANADA	1.00	4.553164 (19.78)	1.042141 (5.55)	1.00	4.757827 (7.83)	0
DENMARK	1.00	0.542179 (13.5)	-0.22972 (-1.99)	1.00	0.515238 (10.2)	0
GERMANY	1.00	1.994009 (9.95)	1.318104 (8.73)	1.00	0.690840 (3.45)	0
ITALY	1.00	1.568617 (3.12)	-0.010976** (-0.012)	1.00	1.561349 (3.12)	0
JAPAN	1.00	-1.833648 (-2.25)	-1.037656 (-7.35)	1.00	-0.83088** (-0.674)	0
NETHERLANDS	1.00	0.882087* (1.71)	3.747938 (6.13)	1.00	4.584104 (3.27)	0
NORWAY	1.00	0.296982 (7.25)	0.240538 (4.01)	1.00	-0.02619** (-0.74)	0
SWEDEN	1.00	0.775681 (19.25)	0.115919 (2.2)	1.00	0.500234 (8.33)	0
UK	1.00	1.019265 (2.97)	-2.812291 (-4.07)	1.00	1.363578 (3.23)	0
USA	1.00	-1.740930 (-7.25)	0.814363* (1.84)	1.00	-2.104034 (-19.09)	0

Notes: (1) Values in parentheses are t-values. (2) All values are significant at 5% level, except the values with (*) and (**). (3) Values with (*) are significant at 10% level. (4) Values with (**) are insignificant.

The interpretations of country – specific estimation results are as follows:

- Canada is a country in which the BS hypothesis holds in both the restricted and unrestricted models due to positive and significant β_1 coefficients. Moreover, Canada has the biggest β_1 coefficient in both the restricted and unrestricted models among 10 countries with values of 4.55 and 4.75, respectively. These results imply that in Canada, 1% increase in PRO leads to an increase in REER around 4.5%. Put another way, the biggest effect of relative productivity on REER takes place in Canada among 10 countries. As we indicated in Figure 1, Canada was the leading country in “average annual REER growth rate”, the big size of β_1 might stem from this. In addition β_2 is significant and equals to 1 for Canada, which is consistent with theory. This implies

that 1% increase in the “terms of trade” associates with 1% increase in the real effective exchange rate of Canada.²³

- Denmark, Germany and Italy are three countries in which the BS hypothesis holds in both the restricted and unrestricted models due to positive and significant β_1 coefficients. In particular, the estimated β_1 coefficients are 0.54 in Denmark, 1.99 in Germany and 1.56 in Italy in the unrestricted model. In the restricted model the estimated β_1 coefficients of these countries are as follows; 0.51 in Denmark, 0.69 in Germany, and 1.56 in Italy. In both models, the estimated β_1 coefficients are significant at 5%. Hence, it is fair to conclude that the BS hypothesis performs well for these countries. Although the TOT effect (β_2) is 1.31 (consistent with prediction) in Germany, β_2 has negative sign in Denmark and Italy which is inconsistent with the prediction of the BS hypothesis. In addition, β_2 has been found insignificant for Italy, which implies the TOT variable has no effect on REER of Italy.
- Netherlands, Sweden and UK are other three countries in which the BS hypothesis holds in both the restricted and unrestricted models due to positive and significant β_1 coefficients. In the unrestricted model, the estimated β_1 coefficients are 0.88 in Netherlands, 0.77 in Sweden and 1 in UK. In the restricted model, the estimated β_1 coefficients are 4.58 in Netherlands, 0.5 in Sweden and 1.36 in UK. Furthermore, in the unrestricted model estimated β_2 coefficients are 3.74 in Netherlands, 0.11 in Sweden and -2.8 in UK and all are significant. It is worth noting that exclusion of TOT led to a significant increase in the BS effect of Netherlands. This is because of the relative importance of TOT for Netherlands. As seen in Table 8, the biggest estimated β_2 coefficient belongs to Netherlands among 10 countries. If we exclude the TOT variable, perhaps the TOT effect merges with the BS effect which leads to over-estimation of β_1 in the restricted model. Put another way, to explain the long-run REER movements in Netherlands, one should use the TOT variable as an explanatory variable which may help him to estimate a less-biased BS effect.

²³ Such a big BS effect seems over-estimated for Canada, thus it has to be used with a caution. A similar BS effect is estimated by Chowdhury (2007) for Australia. He estimated the BS effect 5.6 and he suspects that “the elasticity coefficient is over-estimated due to the exclusion of relevant explanatory variables” (Chowdhury, 2007:p.3). Thus, for the BS model of Canada inclusion of some relevant explanatory variables such as openness, government consumption, and total consumption may generate less-biased β_1 coefficient.

- Japan is a country in which the BS hypothesis fails in both the restricted and unrestricted models due to negative signs of β_1 coefficients. Moreover, in the restricted model β_1 coefficient is insignificant. The negative and significant (-1.83) β_1 coefficient in the unrestricted model implies that 1% increase in the relative productivity leads to -1.83% decrease in the REER of Japan, which is against the prediction of the BS hypothesis. Similarly, β_2 is found negative (-1.03) in the unrestricted model which is inconsistent with the prediction of the BS hypothesis. As indicated in Figure 3, Japan is at the last rank among 10 countries in terms of the “average annual TOT growth rate”. The relatively poor TOT performance of Japan might lead a negative β_2 coefficient.
- USA is a country in which the BS hypothesis fails in both the restricted and unrestricted models due to negative signs of β_1 coefficients. That is to say, the BS effect works in an opposite way that the relative productivity growth leads depreciation in the “REER of USA”, as in Japan. As presented in Figure 1, USA placed at the bottom among 10 countries in terms of the “average annual REER growth rate”. Thus, one may claim that because of its relatively poor performance in REER growth, the estimation generated the wrong sign for the BS effect.²⁴ Although the estimated BS effect of USA is negative, USA has a positive and significant TOT effect which is 0.81. In this respect, the unrestricted BS model performed partially well for USA since the TOT effect has a correct (positive) sign.
- Norway is the only country that the BS hypothesis holds in the unrestricted model and fails in the restricted one. In particular, the estimated β_1 is positive (0.29) and significant in the unrestricted model that the BS hypothesis holds. Nonetheless, when we exclude the TOT variable, β_1 turns out negative (-0.02) and insignificant which implies the rejection of the BS hypothesis for Norway in the restricted model. This result shows that TOT is a crucial variable in explaining the “REER of Norway”. Recall the conclusion that we drew from Figure 3; “in 1975-2007 period, Norway is the leading (outlier) country in terms of the “average annual TOT growth rate” among 10 countries which indicates the importance of TOT variable for Norway. In this regard, there is consistency between the findings of the data analysis and estimation results.

²⁴ Drine and Rault (2005) also rejected the BS hypothesis for USA due to negative sign of the BS effect.

Table 9 summarizes the country-specific estimation results regarding the validity of the BS hypothesis. According to Table 9, the BS hypothesis is valid 8 out of 10 countries when the unrestricted model is used and valid 7 out of 10 when the restricted model is employed. And the TOT effect has been found as predicted (positive and significant) in 6 out of 10 countries. In the unrestricted model, the average BS effect of 10 OECD countries is 1.6 and the average TOT effect of 10 OECD countries is 0.34. In the restricted model, the average BS effect of 10 OECD countries is 1.1, which is consistent with theory and previous studies' results.²⁵

Table 9. Summary of Country-Specific Estimation Results

	UNRESTRICTED MODEL (eq.13)	RESTRICTED MODEL (eq.12)
BS HOLDS	8 out of 10	7 out of 10
BS FAILS	2 out of 10	3 out of 10
AVERAGE BS EFFECT	1.6	1.1
AVERAGE TOT EFFECT	0.34	-

5.4 Panel Cointegration and Estimation Results

Panel cointegration tests have become popular especially after Pedroni’s influential papers in 1996 and 1999. Pedroni’s panel cointegration test fills an important gap in the literature that enables scholars to test whether there is long-run relation among panel series.

More technically, the Pedroni cointegration test uses the following equations in order to test the existence of cointegration in panels:

$$y_{it} = \alpha_i + \gamma_t t + \beta_{1i}x_{1it} + \beta_{2i}x_{2it} + \dots + \beta_{ki}x_{kit} + e_{it} \dots\dots\dots(17)$$

$$e_{it} = \rho_i e_{i,t-1} + \varepsilon_{it} \dots\dots\dots(18)$$

In the Pedroni panel cointegration test the value of ρ_i in equation (18) determines the result, as in the Engle-Granger two-step approach. Pedroni developed two-type of main statistics to make distinction within-dimension and between-dimension. Table 10 reports the results of the Pedroni panel cointegration test for the panel series of “REER, PRO and TOT”.

²⁵ See for example, Alexius and Nilsson (2000), and Egert (2002).

Table 10. Pedroni Panel Cointegration Test Results

<i>(Within dimension)</i>	Statistic	Prob.
Panel v-Statistic	1.798412*	0.0792
Panel rho-Statistic	-2.42544	0.0211
Panel PP-Statistic	-4.302275	0.0000
Panel ADF-Statistic	-4.367117	0.0000
<i>(Between dimension)</i>	Statistic	Prob.
Group rho-Statistic	-1.086104**	0.2212
Group PP-Statistic	-3.846245	0.0002
Group ADF-Statistic	-3.918545	0.0002

Notes: (1) All values are significant at 5% level except (*) and (**). (2) Value with (*) is significant at 10% level. (3) Value with (**) is insignificant. (4) Test is conducted by assuming “deterministic intercept and trend”, and choosing automatic AIC with a max lag of 7.

More explicitly, the within-dimension statistics test the following hypotheses:

$$H_0: \rho_i = 1 \text{ (No cointegration)}$$

$$H_1: \rho = \rho_i < 1 \text{ (All individuals in panel are cointegrated, and } \rho = \rho_i \text{ for all } i)$$

For example, by using “Panel rho-statistic” which is one of the four “within dimension” statistics, we reject the null hypothesis at 5%, since -2.42 is significant that its probability is 0.02. Hence, we accept the alternative hypothesis and conclude that all individuals in panel are cointegrated, with assuming $\rho = \rho_i$, for all i .

And the between-dimension statistics test the following hypotheses:

$$H_0: \rho_i = 1 \text{ (No cointegration)}$$

$$H_1: \rho_i < 1 \text{ (All individuals in panel are cointegrated for all } i)$$

For example, by using “Group PP-Statistic” which is one of the three “between dimension” statistics, we reject the null hypothesis at 5%, since -3.84 is significant that its probability is 0. Hence, we accept the alternative hypothesis and conclude that all individuals in panel are cointegrated, without assuming $\rho = \rho_i$, for all i .

In words of Harris and Sollis (2005:p.202) “pooling along the within-dimension amounts the effectively pooling ρ_i in equation (18) across the different (i) individuals in the panel such that $\rho_i = \rho$. The between-dimension estimators are based on averaging the individual

estimated values of ρ_i for each member i . Put differently, the between-dimension statistics (estimators) are less restrictive in the sense that they allow heterogeneity across individual members.” This difference is evident in their alternative hypotheses.²⁶

If we repeat the panel cointegration test by using all seven statistics in Table 10, we accept the alternative hypothesis six times and reject it only in use of “Group-rho statistic”. According to the values in Table 10, we accept the existence of cointegration relation among the whole panel series of “REER, PRO and TOT”.²⁷

To check the robustness of this result, we conduct two alternative panel cointegration tests and report their results in Table 11.

Table 11. Alternative Panel Cointegration Test Results

Kao Residual Cointegration Test								
	ADF t-Statistic	Prob.						
PANEL	-10.42*	0						
Johansen-Fisher Panel Cointegration Test								
	Fisher Stat. from Trace Test				Fisher Stat. from max-eigen Test			
	tr (r=0)	tr (r ≤ 1)	tr (r ≤ 2)	Rank	λ (r=0)	λ (r ≤ 1)	λ (r ≤ 2)	Rank
PANEL	236.9* (0.0)	62.18* (0.0)	23.22 (0.2782)	2	209.1* (0.0)	63.86* (0.0)	23.22 (0.2782)	2

Notes: (1) The Kao test uses the null of no cointegration and it is conducted by assuming “no deterministic trend” (default), and choosing automatic 8 lags by AIC with a max lag of 8. (2) The Johansen-Fisher test is conducted by assuming “linear deterministic trend” with 5 lags and probabilities are presented in parentheses. (3) Johansen-Fisher tests $r = 0$, $r \leq 1$, $r \leq 2$. (4) Values with (*) are significant at 5% level.

The Kao residual (panel) cointegration test uses the ADF test type t-statistic to test the same null hypothesis (no cointegration) as Pedroni. According to Table 11, the t-statistic of Kao for the whole panel is -10.42 which is significant at 5%. Hence, we reject the null and accept the existence of cointegration relation among the panel series of “REER, PRO, and TOT”.

The second alternative panel cointegration test that we conduct is the Johansen- Fisher panel cointegration test; which tests the null of no cointegration as Pedroni and uses a similar

²⁶ For details see; Harris and Sollis (2005:pp.200-206), Pedroni (1999), and E-views (2007:pp. 372-376).

²⁷ Six out of seven statistics confirmed the cointegration relation, thus rejection of the alternative hypothesis in use of Group rho-statistic can not affect the general result of “the existence of cointegration”.

approach as the Johansen cointegration. The Johansen- Fisher panel cointegration test reports trace statistics and maximum eigenvalues to test the “rank of cointegration” among the panel series. In other words, we can learn whether there is one or more than one cointegration relation among the panel series of “REER, PRO and TOT” by using the Johansen-Fisher test. In this respect, it has a distinct advantage over Kao and Pedroni. But its results are as sensitive as the Johansen approach to the lag-length and trend-specification, as we discussed in section 5.3. When we apply the same steps as the Johansen cointegration test, the Johansen-Fisher trace and maximum eigenvalue test results indicate that the “rank of cointegration” is two among the panel series of “REER, PRO and TOT”. Put another way, Johansen- Fisher test confirms that there is cointegration among the panel series. Moreover it says that the number of cointegrating vectors is two among three panel series.

There is consistency among the findings of three panel cointegration tests in that they all found out cointegration among the panel series of “REER, PRO and TOT”. In addition, these findings are in favor of the BS hypothesis in the sense that they all imply the existence of a long-run relation among three series.

After finding the existence of long-run relation among three panel series, now we should find the size and sign of this relation. In other words, we should estimate the β_1 and β_2 coefficients in order to understand whether the BS hypothesis holds or not for the panel of 10 OECD countries. In econometrics, there is no perfect estimation method in “estimation of panel-cointegrated series”. The determination of estimation method in panel-cointegrated series is an important issue for the robust and unbiased results. If the series are panel-cointegrated, OLS method usually generates biased coefficients.²⁸ Harris and Sollis (2005) suggest FMOLS and DOLS estimation methods to estimate less unbiased coefficients, which are widely used in empirical studies. However, there is no consensus among scholars concerning which method is better. “Although Pedroni (2000) found that group-means DOLS estimator has higher size distortions than group-means FMOLS estimator, Kao and Chiang (2000) showed that FMOLS may be more biased than DOLS” (Harris and Sollis, 2005:p.207). Apart from these two methods, ARDL method, which is based on OLS, is still in use in studies where panel-cointegrated series have been estimated. Panopoulou and Pittis (2004) found out

²⁸ See for example Kao and Chiang (2000:p.179) and Hayakawa and Kurozumi (2006:p.2).

that ARDL generates better results than both DOLS and FMOLS in the case of a serially correlated cointegration error; however, this is not a widely accepted result.

The Dynamic Ordinary Least Squares (DOLS)

By following Kao and Chiang (2000), who found out that “the DOLS outperforms both the OLS and FMOLS estimators”, we chose to employ the DOLS method to estimate the panel regressions of the BS model.

The dynamic OLS (DOLS) estimator of Phillips and Loretan (1991), Saikkonen (1991), and Stock & Watson (1993) is a parametric approach where leads and lags of the regressors are explicitly added to the model. As Hayakawa and Kurozumi (2006:p.2) note “the DOLS estimators are known to be asymptotically equivalent and efficient”. In addition, Hayakawa and Kurozumi (2006:p.4) demonstrate that “leads are not necessary in cointegrating regression models when the cointegrating regression error does not Granger-cause the first difference of the I (1) regressors.” We follow the common convention and use both “lead and lag” of the regressors in our DOLS estimations. By applying to “general to specific” approach, we decide to use “one lead and one lag” of the regressors since the higher degree leads and lags are found insignificant.²⁹

Table 12 presents the estimation results of the BS model (equations 12 and 13) for the panel of 10 OECD countries by using the DOLS method. According to the estimation results, the BS effect is found as -0.45 and significant at 5% in the unrestricted model. This implies that %1 increase in the “panel PRO series” leads to 0.45% decrease in the “panel REER series”. And when TOT is excluded (equation 12), the BS effect is estimated as -0.83 which is significant at 5%. These findings are against the predictions of the BS hypothesis. In other words, the panel estimation results reject the BS hypothesis since the β_1 is negative in both models. On the other hand, the TOT effect is found positive and significant as 0.52 in the unrestricted model, which is consistent with what the extended BS hypothesis suggests.

²⁹ Harris and Sollis (2005:pp.206-211) presents the basics of the DOLS method. See Mark and Sul (2003) for detailed exposition of the DOLS method with an empirical application. For a similar application to our study, see Choudhri and Khan (2005) that they estimated the BS model with DOLS.

Table 12. Panel Estimation Results

DEPENDENT VARIABLE: REER		
	Estimation Method: DOLS	
	Eq.13	Eq.12
δ (intercept)	2.31	4.74
t-value (probability)	2.63 (0.009)	230 (0.00)
β_1 (the BS effect)	-0.45	-0.83
t-value (probability)	-1.97 (0.0492)	-4.17 (0.00)
β_2 (the TOT effect)	0.52	
t-value (probability)	2.75 (0.0062)	
Adjusted R-Square	0.832	0.829
d.w	2	1.93
F-stat	82	97

Notes: (1) All coefficients are significant at 5%. **(2)** The estimations are conducted by choosing cross-section fixed effects and no-GLS effects with ordinary coefficient covariance method.

The rejection of the BS hypothesis in the panel estimations does not necessarily imply that the BS hypothesis is completely invalid. In contrast some important factors, which does not stem from the BS hypothesis itself, may lead the rejection of the BS hypothesis in panel series.

These are:

- (a) The usage of DOLS estimation method.
- (b) The usage of panel data comprising insufficient numbers of N and T; (N: Number of included countries to panel, and T: length of period).
- (c) The inclusion of some specific countries (USA, Norway, Japan) in which the BS hypothesis fails, as our country-specific results showed.³⁰

In sum, a researcher who uses a different dimension-panel, a benchmark country and an estimation method can find the validity of the BS hypothesis for the panel series. For example, Alexius and Nilsson (2000) found out the BS hypothesis is valid for a panel of 15 OECD countries by using USA as a benchmark country and employing FMOLS method.

Although we have found cointegration relation among three series in panel, the panel estimations could not confirm the validity of the BS hypothesis due to negative sign of β_1 .

³⁰ When we exclude Japan, Norway, and USA and re-run the BS model, β_1 remained negative in this reduced panel. However, a different reduced panel may generate a positive β_1 coefficient.

The debate on panel cointegration and estimation methods is still going on among econometricians. However, it is clear that by pooling the data of 10 countries, we lose the country-specific BS and TOT effects, moreover our panel findings are less reliable than our country-specific results due to aforementioned discussion on the panel estimation methods.

We are closing this section with a warning of Harris and Sollis (2005) on panel cointegration and estimation techniques:

“However, panel cointegration tests are still at a fairly early stage of development, mostly being based on a single equation approach. Moreover, tests for cointegration occurs are typically undertaken separately from estimating the cointegrating vector, and there are issues over which estimator is to be preferred [the non-parametric (FMOLS) versus parametric (DOLS)]. Still, there are likely to be significant advances in panel data technique, it is unlikely that unambiguous results will be obtained with panel data comprising moderate values of N and T.” Harris and Sollis (2005:p.212)

6. CONCLUSION

The BS hypothesis, which explains the real exchange rate movements with productivity differences between countries, still keeps its importance in empirical economics. Since 1964, dozens of published studies have tested the hypothesis by using different methods. The theoretical contributions such as the inclusion of “terms of trade” and improvements in econometric techniques such as in panel cointegration motivate scholars to test the validity of the BS hypothesis.

Our study, which tests the validity of the BS hypothesis in 10 OECD countries, uses two approaches (country-specific and panel) with employment of recent cointegration and estimation methods. The findings of the study are mostly in favor of the BS hypothesis. First of all, both country-specific and panel cointegration tests showed that there is long-run relation among REER, PRO and TOT variables. In the unrestricted model, country-specific coefficient estimations of the BS model confirmed that the BS effect is positive and significant in 8 out of 10 countries. And the BS effect has been found positive and significant in 7 out of 10 countries in the restricted model. Generally speaking, TOT is a critical variable

in the BS hypothesis that we found out positive (predicted) and significant coefficients in 6 out of 10 countries. In particular, it is relatively more important for Netherlands and Norway. In Netherlands, the exclusion of the TOT variable from the BS model led over-estimation of the BS effect and in Norway it led to the rejection of the BS hypothesis. Among 10 OECD countries, Japan and USA are two specific countries in which both the BS and TOT effects have been found negative probably due to country-specific reasons. It seems that for Japan and USA, inclusion of some additional explanatory variables such as total consumption, government consumption, and openness may help us to get better results regarding the BS effect. In addition, relaxing the assumption that we made about non-tradables sector may affect the results, especially for USA in where non-tradables (services) sector productivity growth is more evident.

In the study, more interesting results have been found in panel estimation section. Although panel cointegration tests confirmed the existence of cointegration, the panel estimation of the models rejected the validity of the BS hypothesis due to wrong signs of the BS effect. But, the TOT effect has been found positive and significant in panel estimation, as the BS hypothesis predicts. Briefly, the panel results revealed mixed results about the validity of the BS hypothesis. Thus, it is not fair to reject the BS hypothesis by only looking at the findings of the panel estimation results. As we discussed, the panel estimation and cointegration tests are in their early stages that their results should be interpreted carefully. Furthermore, as Garcia-Solanes *et al.* (2008) note, factors which distort the perfect competition, such as political interests, segmentation in tradables good market and transportation costs may affect the validity of the BS hypothesis.

Finally, today the BS hypothesis, which cannot be rejected widely in empirical studies (as in this study), still plays an important role in explaining why the PPP fails and the real exchange rates differ across countries. With the improvements in econometric methods and availability of new sectoral productivity data, it seems to keep its importance also in the future.

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