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The Balassa–Samuelson Hypothesis in Developed Countries and Emerging Market Economies: Different Outcomes Explained

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Abstract

This paper studies the Balassa–Samuelson effects in two areas with strong differences in economic development, sixteen OECD countries and sixteen Latin American economies. The USA is taken as a benchmark. Applying recent panel cointegration and bootstrapping techniques that solve for cross-sectional dependence and small panel size problems, we find some evidence for not rejecting the whole hypothesis in the LA area. In the context of OECD group, the second stage of the BS hypothesis, which relates relative sector prices with the real exchange rate, does not hold, probably because national markets remain to some extent segmented, as reflected in departures from PPP in the tradable sectors.

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1 Introduction[1]

The Balassa and Samuelson (BS) hypothesis (Balassa (1964), Samuelson (1964)) provides a theoretical explanation of the long run trends in the real exchange rates (RER). Its central tenet is that countries with faster productivity growth in their tradable sector –compared to growth in the non-tradable one– will experience an equilibrium real appreciation of their currency. Since improvements in the tradable sector productivity are normally linked to economic growth, a correlation between relative economic development and the real exchange rate is also postulated. Thus, it is expected that countries growing faster will tend to experience real exchange rate appreciations with respect to other, slow growing economies. The BS hypothesis has important implications for exchange rate policy and for the trade-off that many countries face between inflation targets and exchange-rate stability.

The empirical evidence obtained so far regarding the BS effect indeed indicates that the best results occur in the context of economies that grow at very divergent speeds, such as Japan compared to the USA in the post World War II period (see, for instance, Hsieh (1982) and Marston (1987)), and transition countries that need to grow very fast if they are to catch up with the standards of living of their developed neighbours. This is the situation in some Southern East Asian countries (Ito, Isard and Symansky (1997)) with respect to Japan during the seventies and eighties, and in Central and Eastern European countries compared with Germany since the early nineties (Halpern and Wyplosz (2001), Kovács (2002), Égert (2002a,b), Mihaljek and Klau (2003), Égert et al. (2002)).

Some recent papers have investigated the fulfilment of the BS effect in groups of developing countries taking the USA as the reference external country. Thus, Drine and Rault (2003), tested the BS hypothesis using annual data from the period 1990-1999 for twenty Latin American countries, and found that the hypothesis holds not only for the whole area, but also for Central American and South American groups of countries when considered separately. Calderon and Schmidt-Hebbel (2003) found that for five sub-periods that span the 1990's, RER changes predicted by productivity growth are in the same direction as actual changes in 13 of the 18 Latin American countries analysed. Finally, Choudhri and Khan (2004) tested the hypothesis with a panel data composed of 18 years (1976-1994) and 16 developing countries and obtained strong verification of BS effects in this area.

The empirical findings referring to economies that do not exhibit pronounced divergences in economic development between one another, such as groups of countries in the OECD, are not unanimous. For example, whereas Alberola and Tyrväinen (1998), Chinn and Johnston (1999) and MacDonald and Ricci (2001) obtained positive results for the whole general BS proposition, Canzoneri, Cumby and Diba (1999) found favourable evidence only for that part of the hypothesis that links the productive differential with the relative price of the tradable and non-tradable sectors. Heston, Nuxoll and Summers (1994) found that the difference between tradable and non-tradable prices moved with the income levels of OECD countries, which is consistent with the results of Canzoneri, Cumby and Diba (1999). According to Tille (2001),

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productivity developments accounted for 2/3 of the US dollar's appreciations against the Euro and 3/4 of its appreciation against the Japanese yen in the nineties. Lothian and Taylor (2006) derived a 40 percent effect in the case of the Sterling pound/US dollar real exchange rate during the very long period 1820-2001, after allowing for non-linear adjustments and volatility shifts across the exchange rate regimes. However, these authors did not find empirical support for the BS effect in the case of the French franc/US dollar exchange rate.

Despite the fact that the magnitude and the statistical significance of the empirical results seem sensitive to the level of economic development of the areas analysed, no empirical study attempts to compare the fulfilment of the BS hypothesis in two areas which exhibit sharp differences in standards of living and growth with respect to a common, foreign, developed country. To fill this gap, we undertake such a comparative analysis in the context of sixteen OECD countries and sixteen Latin American economies. We take the USA as the benchmark to calculate productivity and price differentials, as well as real exchange rates, and use the same sample period, and identical theoretical and empirical approaches in both cases.

The BS hypothesis is, in fact, composed of two stages. The first (denoted BS-1 hereafter) is known in the literature as the Baumol and Bowen (1966) effect and relates the difference in productivities with the difference in prices of the non-tradable and tradable sectors. The second (BS-2) establishes the link between the non-tradable/tradable price differential and the real exchange rate measured with CPI deflators. This second relationship is immediately obtained if PPP holds in the tradable sector. In order to look at the BS hypothesis more closely and detect the origin of the failure when the results for the entire BS hypothesis are poor, we test each part of the hypothesis separately, using the same procedure as Canzoneri, Cumby and Diba (1999).

This paper presents two novelties that contribute to improve the empirical results. First, we move into a sectoral dimension to classify the branches of activity into tradables and non-tradables according to the disaggregated methodology of the United Nations, which is a more rigorous approach than previously used. This allows us to obtain a direct and accurate measure of the labour productivity differential, which is more precise and reliable than the proxy GDP per capita used in previous studies such as Drine and Rault (2003) and Lane and Milesi-Ferretti (2004). For this task, we use a number of different sources to assemble data sets for two groups of countries, which are then utilised to build the relevant variables. Our statistical sources are the OECD (*National Accounts of the OECD countries*) and EUROSTAT, national banks, national statistic institutes, CEPAL (Economic Commission for Latin America and Caribbean of the United Nations), and the ILO (International Labour Organisation). We use annual observations from the period 1991-2004.

Second, since cross-sectional dependence on the panel data is usually present in countries with important economic links – especially where all variables are defined relative to a common denominator (the relevant USA variable) – we apply panel unit root and cointegration tests developed recently to cope with this problem. Furthermore, we apply non-parametric bootstrapping techniques to eliminate the bias created by series that are relatively short. This is a clear methodological improvement with respect to previous works that apply only conventional panel data unit root and cointegration tests, such as Drine and Rault (2003) and Choudhri and Khan (2004).

As an advance of our findings, we obtain evidence that the first stage of the BS effect can not be rejected in both groups of countries. The coefficient of the productivity differential has the correct sign, and its absolute value lies in the range established by the theoretical model in all cases. Nonetheless, contrary to previous studies that boast of unanimous positive results – derived with conventional cointegration techniques - our result must be interpreted with caution since many of our tests reject the cointegration hypothesis. We do not find evidence to show that estimates of the BS-1 are better in one group than in the other. However, for the second part of the hypothesis (BS-2) we find that PPP may be accepted for the tradable sectors of the Latin American countries as a whole but not for the group of OECD countries.

The failure of PPP in the tradable sectors in developed areas is not surprising either on theoretical or on empirical grounds. As we explain below, theoretical explanations may be found in pricing-to-market practices (Krugman(1987)), transportation costs (Rogoff (1996)), and imperfect knowledge of markets participants (Frydman and Goldberg (2007) and Frydman et al. (2008)). Empirical evidence against PPP in the tradable sectors of developed countries is provided, for instance, by Canzoneri, Cumby and Diba (1999), Sǿndergaard (2001) and Égert et al (2002).

The remainder of this paper is organised as follows. In Section 2 we derive the two parts of the BS hypothesis. In Section 3 we explain the composition of our tradable and non-tradable sectors and the way in which the variables of interest are measured. This section also includes a descriptive analysis of the main relationships that will be tested econometrically and discussed in Section 4. Finally, in Section 5 we summarise the main findings and derive some policy implications.

2 Theoretical Framework

2.1 The Balassa and Samuelson Model

To test the BS model for two sets of countries and to investigate the causes of success or failure, we follow the two-step procedure of Canzoneri et al. (1999). We then split the analysis in two parts, which we name BS-1 and BS-2. We analyse each part separately.

The **first part of the BS hypothesis,** known as the Baumol and Bowen (1966) effect, links the difference in *total* productivities with the difference in prices of tradable (T) and non-tradable (N) sectors. Under the usual assumption of factors mobility and perfect competition, and Cobb-Douglas production technology in each country, it is easy to derive[2]:

$$dp = \frac{\beta}{\alpha}\left(a_T - a_T^*\right) - \left(a_N - a_N^*\right) \tag{1}$$

The price differential (dp) is defined as: $dp = \left(p_N - p_N^*\right) - \left(p_T - p_T^*\right)$.

Variables a_T and a_N are the logs of *total* factor productivity in the tradable and non tradable sector, respectively. Coefficients β and α stand for the intensity of labour in

[2] See, for instance, Wagner and Hlouskova (2004), and Égert et al. (2005)

the production function of sectors N and T, respectively. Finally, p_T and p_N are the logs of the price index of each sector. Superscript (*) refers to the foreign country.

This traditional version of the model poses important empirical problems because most countries lack reliable data on capital stocks, which are necessary to compute total factor productivities. For this reason, some authors, for example Kohler (2000) and Sarno and Taylor (2001), suggest an adapted version of the BS model in terms of average *labour* productivities (*alp)*, which can be readily measured. The testable equation is:

$$dp = \left(alp_T - alp_T^*\right) - \left(alp_N - alp_N^*\right) \tag{2}$$

Equation (2) establishes that the price differential is determined by the difference between the relative labour productivities of the tradable and non-tradable sectors of the two countries. Compared with the traditional version in terms of total factor productivities, Equation (2) has an important implication: the coefficients of the labour productivities are all equal to unity and are directly testable.

Expressed in terms of first differences, Equation (2) would indicate that economies with particularly high increases in tradable labour productivity relative to non-tradable labour productivity will exhibit relatively high increases in the relative price of non-tradables, everything else constant.

The **second stage of the BS hypothesis** establishes a relationship between the price differential (dp) and the log of the real exchange rate measured with CPI indices *(* q), as indicated by the following expression:

$$q = \left(e + p_T^* - p_T\right) - \lambda dp \tag{3}$$

Where e is the natural log of the nominal exchange rate defined as the price of the foreign currency in terms of the domestic one, and the value of q is given by the expression $q = e + p^* - p$, in such a way that a decrease (increase) in q indicates a real appreciation (depreciation) of the domestic currency. The coefficient λ is the weight of non-tradable goods in the consumer's basket, and it is assumed to be identical in the two countries. The first parenthesis in expression (3) stands for the natural log of the RER calculated with the prices of tradable goods, and is known as the external RER (RER(T)). By assuming that PPP holds in sectors T, as is usually accepted, this parenthesis is equal to zero, and the second part of the BS may be written as:

$$q = -\lambda dp \tag{4}$$

According to (4), there is a negative relationship between the difference in the relative price ratios and the CPI-deflated real exchange rate: an increase in the price differential causes a RER appreciation, which is more pronounced the bigger the weight of N goods in the consumers' basket. It is worth noting that the second part of the BS hypothesis, as presented in Equation (4), relies crucially on the fulfilment of PPP in the tradable sector.

Joining the two BS parts we obtain the **complete BS hypothesis**:

$$q = -\lambda\left[\left(alp_T - alp_T^*\right) - \left(alp_N - alp_N^*\right)\right] \tag{5}$$

It suggests that the real appreciation in the exchange rate should be equal to the increase of the productivity differential transmitted to the CPI via the non-tradable inflation pass-through.

2.2 Failure of PPP in the Tradable Sector: the Quality Bias and Market Segmentation

As explained above, PPP in the tradable sector (PPP(T)) is an important pillar of the second stage of the BS hypothesis. Several studies, using different statistical and econometric methods and different geographical samples, provide evidence against PPP(T)[3].

To gain further insight into the sources of PPP(T) failure, we split the RER(T) into three components following a simple accounting procedure[4]:

$$q_T = \left(\delta + \delta^* - 1\right)\tau + \left(1 - \delta^*\right)\left(e + p_H^* - p_H\right) + \left(1 - \delta\right)\left(e + p_F^* - p_F\right) \tag{6}$$

Where q_T stands for the RER(T), δ, (δ^*) is the share of domestic (foreign) tradable goods within the tradable basket of domestic (foreign) consumers, and p_H, (p_F^*) is the price index of the tradable goods produced in the domestic (foreign) country, measured in the own currency. Parameter τ represents the terms of trade, and is defined as:

$$\tau = e + p_F^* - p_H \tag{7}$$

If, as pointed out by Obstfeld and Rogoff (2000), consumers of each country prefer home produced tradables compared to those produced abroad (home bias), both parameters, δ and δ^* will be greater than ½ and the first parenthesis of the Equation (6) will be unambiguously positive.

Equation (6) indicates that there are two broad factors that cause variations in the external real exchange rate. The first operates through **variations in the terms of trade** when home produced and foreign produced tradables are not homogeneous. Improvements in the relative quality of domestic tradables, for example, appreciate the terms of trade, which in turn appreciate the RER(T) (Cincibuch and Podpiera (2006))[5].

[3] Canzoneri, Cumby and Diba (1999) found large deviations from PPP(T) when looking at US dollar exchange rates in a group of fourteen OECD countries, and García Solanes, Sancho and Torrejón (2008) rejected BS-2 in a group of six EU-15 economies. Wu (1996) rejected PPP(T) with data of Taiwan, and Ito et al. (1997) and Chinn (1997) obtained similar results using data of several groups of Asian countries. Sǿndergaard (2001) showed that the RER(T) of ten OECD countries exhibited movements that were linked to cross-country growth differences in traded sector unit labour costs. Finally, Égert (2002a), Blaszkierwicz et al. (2004) and Égert et al (2002) also found unfavourable evidence for this relationship in a group of nine Central and Eastern European countries when taking the EU as a benchmark.

[4] See García-Solanes, Sancho and Torrejón (2008) for a detailed derivation.

[5] For this connection between variations in quality of tradable goods and variations in the RER, it is necessary that the statistical bodies do not correctly reflect the incidence of quality on the evolution of the CPI.

The stimulus in the demand for home tradables (García-Solanes, Sancho and Torrejón (2008)) also adds appreciating pressure on the terms of trade[6].

The second group of factors that may cause variations in the RER(T) arises as a result of **market segmentation,** since the lack of perfect integration between regional and/or national markets precludes that the same national tradable goods have the same price across markets. This would imply that the third and fifth parentheses are significantly different from zero in (6). Market segmentation may be due to two causes: a) *imperfect competition*, which may give rise to "pricing-to-market" practices[7] (Krugman (1987)), and b) *arbitrage frictions*, as a result of transportation costs (Rogoff (1996)), information costs and non-tariff barriers.

Market segmentation creates a band within which differences in prices of identical goods sold in two countries can move without triggering arbitrage transactions. In that case, adjustment towards the law of one price (LOOP), which lies at the centre of the band, is slow. However, when prices drift outside the range, arbitrage profits emerge and the ensuing transactions push prices quickly back towards the LOOP[8]. Maier (2004) stressed the fact that the width of the non-arbitrage bands increases with exchange rate variability.

The preceding discussion has shown that quality variations and market segmentation may inflict different trajectories to the RER(T). If continuous quality improvements coupled with demand pressure on tradables are the driving force, the result is an appreciating trend in the RER(T). However, when market segmentation is the factor that causes variations in tradable prices, the likely results are random adjustments in the RER(T) within two non-arbitrage bands. Finally, if the width of the non-arbitrage bands is large and/or suffers random variations, the RER(T) will exhibit non stationary movements. As we shall see in Section 4.4, speculative forces in the foreign exchange market can generate long swings in the RER(T), which may persist during the time that the above mentioned factors are in place.

In the following two sections we perform an empirical analysis of what has been discussed in this section.

3 Construction of Variables and Descriptive Analysis

3.1 Data Sources and Measurement of Variables

The data set used in this study consists of annual data from the period 1991 to 2004. We calculate average productivities of labour, sectoral prices, and real exchange rates. After transforming all series into indices, taking the first year of the sample as the base, we compute natural logarithms. The panel data set covers two groups of countries: 16 Latin

[6] Other sources of terms of trade appreciation are: increases in the regulated prices, improvements in the distribution sector (MacDonald and Ricci (2001)), and the presence of non-tradable components in tradable goods (Rawdanowicz (2004)).

[7] Goldberg and Knetter (1997) survey the sources of "pricing-to-market" policies.

[8] In a study on nine Central and Eastern European countries, Sarno and Taylor (2001) showed that short-term movements of real exchange rates – against the Deutsche mark - follow non-linear adjustments around their trend paths. The speed of adjustments is higher outside the bands than within bands.

American countries (Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominic Republic, Ecuador, El Salvador, Jamaica, Mexico, Nicaragua, Panama, Paraguay, Uruguay and Venezuela) on the one hand, and 16 OECD members (Australia, Austria, Belgium, Canada, Denmark, France, Germany, Ireland, Italy, Korea, The Netherlands, Norway, Portugal, Spain, Switzerland and the United Kingdom) on the other. For each country we take the US economy as the benchmark foreign country, since all the countries mentioned have substantial trade with this economy.

The data sources for constructing the price and productivity indices for the developed countries are the databases of the OECD "National Accounts of OECD Countries Detailed Tables Volume II (2006)" and "National Accounts of OECD Countries Detailed Tables Volume II (2003)". In addition, valuable information from the EUROSTAT of those countries was required to complete some series. The sources for the group of Latin American countries were CEPAL (Economic Commission for Latin America and Caribbean countries of the United Nations), and the ILO (International Labour Organisation). The IMF database was used for the nominal exchange rates of each country in both groups of countries.

In order to calculate productivity and relative prices it is crucial to correctly classify the economic branches into tradable (open) and non-tradable (sheltered) sectors. As in many other empirical analyses, we exclude agricultural activities from the tradable sector in both groups of countries, although for different reasons. In the case of the OECD area, the explanation is twofold; first, the bulk of exports correspond to industrial goods, and second many OECD countries apply protectionist and subsidy policies that distort the volumes of agricultural goods exchanged between them and third countries. In the case of the Latin American area, the exclusion is less evident since the share of agricultural products in total exports of these countries is a far from negligible. However, data on employment in Latin American countries correspond almost exclusively to urban activities, and exclude agricultural work.

Public sector activities were also excluded from the tradable sector in all countries because they are not performed under conditions of free competition, and producers do not behave as profit maximisers. As a result, the components of the **tradable sector** are Manufacturing, Transportation, Storage and Communications and Mining and Quarrying - the last activity includes oil and natural gas extraction. The inclusion of the last branch seems very important in the case of the Latin American countries, which have traditionally been producers and exporters of raw materials. The **non-tradable sector** includes Construction and five categories of private services (Electricity, Gas and Water Supply, Wholesale and Retail Trade, Hotels and Restaurants, Financial Intermediation and Real Estate) and excludes public services because of the lack of data on production and/or employment for those activities.

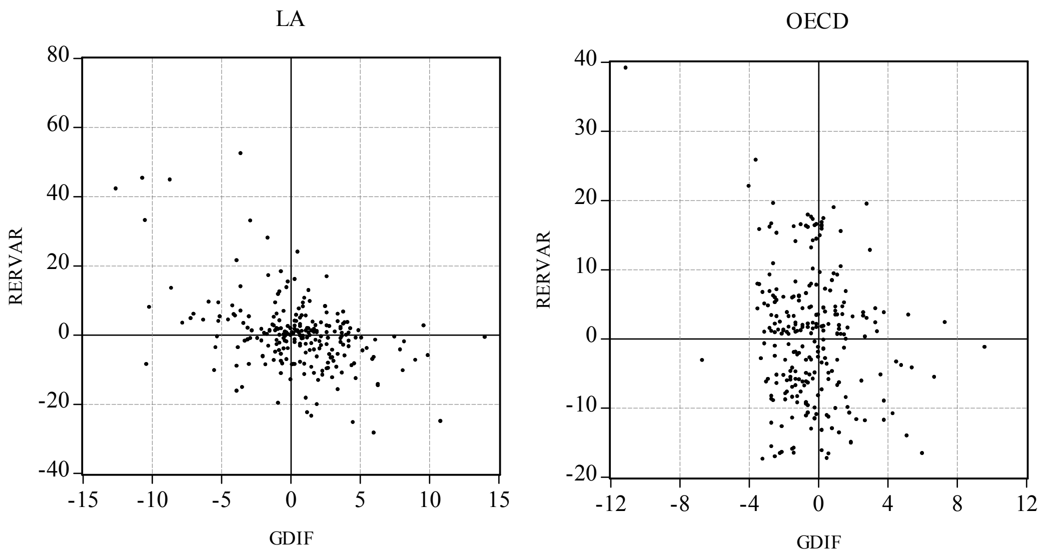
We define the relative price of non-tradable goods with respect to tradable ones as the ratio of the two corresponding sector GDP deflators. To obtain deflator indices we first measured the aggregate production, that is, the value added, in each sector in both nominal and real terms, using current prices and the prices of the base year respectively, and then we calculated the price deflators, P_T and P_N.

To obtain the average productivities of labour we first computed total labour employment in each sector, and then we calculated average labour productivities.

3.2 Descriptive Analysis

As explained above, it can be expected that fast growing countries will show a tendency for real exchange-rate appreciations compared to other, slow growing economies. To verify this in a simple and descriptive way, in Figure 1 we plot for each panel the difference in GDP growth and the variation of the CPI real exchange rate of each country with respect to the USA during the period covered by this study. Differences in growth rates (GDIF) are measured on the X-axis, and variations in the real exchange rates (RERVAR) are measured on the Y-axis.

Figure 1: Growth Differential (GDIF) and Variations in the Real Exchange Rate (RERVAR) in the Two Groups of Countries. Annual Observations (1991-2005)



Taking into account the definition of the real exchange rate that we use, fulfilment of the (complete) BS effect requires that increases in GDIF go with decreases in RERVAR. As can be verified in the graph, this condition probably holds more easily within the group of Latin American countries than in the set of OECD economies covered by our analysis. In fact, while in the group of OECD countries the set of points are grouped probably around a vertical line, in the group of LA economies the set of points may adjust to a negatively sloped line.

In the following section we perform econometric analysis to test rigorously the BS hypothesis and to ascertain whether our first impressions are confirmed.

4 Econometric Analysis

In this section, we apply *panel* unit root and cointegration tests, and non-parametric *bootstrapping* techniques to test the two stages of the BS hypothesis in the two areas under study. Pooling observations is a necessary strategy to raise the reliability of the estimates when the observed period is relatively short (Banerjee, 1999). The panel approach to investigating the BS hypothesis has already been applied by Halpern and Wyplosz (2001), De Broeck and Slok (2001) and Égert et al. (2002) in the context of Central and Eastern European transition countries, by Drine and Rault (2003) using data from a large group of Latin American countries, and by Choundhri and Khan (2004) in a sample of 16 developing economies.

However, the correct application of these techniques depends crucially on the assumptions that: a) individual time series are cross-sectional independent, and that b) the distribution of the test statistics tends towards an asymptotical normal distribution when both the sample (N) and the time (T) dimensions approach infinity. In cases where these assumptions are not satisfied, conventional panel techniques, such as Levin, Lin and Chu (2002) and Im, Pesaran and Shin (2003) for unit root tests, and Pedroni (1999 and 2004) for cointegration tests, may lead to biased results and, generally, to over-rejection of the null hypothesis. To deal with these potential problems, we adopted the econometric steps that we explain in turn.

As a preliminary econometric analysis we evaluated the degree of cross-sectional dependence in our data, by estimating individual ADF(p) regressions for $p = 1, 2, 3$, and computed pair-wise cross section correlations coefficients of the residuals from these regressions. The simple mean of these correlation coefficients, together with the associated cross-sectional dependence test statistic proposed in Pesaran (2004) showed some degree of cross-sectional dependence in all series. In the LA countries, correlations are usually between 0.04 and 0.16, while the OECD countries presented significantly higher correlation coefficients (between 0.12 and 0.60). These results call for the application of the new tests that take into account dependence across countries, especially in the case of the OECD set of countries[9].

Furthermore, since the time series of our panel are not sufficiently large to guarantee the absence of small sample bias, we performed bootstrap inference in the cointegration analysis with the Non-parametric bootstrapping algorithm suggested by Wagner and Hlouskova (2004), which is specially designed to cope with the problems raised by both small samples and cross-sectional dependence[10]. We applied this algorithm to the Levin, Lin and Chu (2002) and Im, Pesaran and Shin (2003) unit root tests, and to the Pedroni (2004) cointegration test. To calculate cointegraion vectors we applied bootstrapping techniques to the test suggested by Westerlund (2007), which takes into considertation cross-sectional dependence in the data.

[9] For example, Jönson (2004) and Pesaran (2007) developed new unit root and stationary tests, respectively, for this situation. Banerjee and Carrion-i-Silvestre (2006) suggested a transformed Pedroni *Group-t* statistics to solve this problem. Their statistics fits an asymptotically normal distribution under the null of no cointegration.

[10] Chang (2004) designed bootstrapping methodologies for panel unit root tests in samples affected by cross-sectional dependence, and Westerlund and Edgerton (2007) applied bootstrap analysis to test the null of cointegration in panels with cross-sectional dependence.

Before starting the cointegration analysis we verified that the following variables have one unit root:

$$dp = \left(p_N - p_N^*\right) - \left(p_T - p_T^*\right),\ da_T = alp_T - alp_T^*,\ da_N = alp_N - alp_N^*,\ e \text{ and } dp_T = p_T - p_T^*.$$

After applying the above-mentioned tests that deal with cross-sectional dependence, the results suggest that each of the variables contains one unit root in the two panels of our study, which justifies further investigation into whether the variables maintain the long run relationships derived from our model[11]. In the following lines we apply cointegration tests and estimate the cointegration vectors when justified.

4.1 The First Stage of the BS Hypothesis. Cointegration Tests

In this first stage we test for a cointegration relationship between variables dp, da_T and da_N. The unrestricted version is represented by Equation (2). However, given that the theoretical model postulates that the coefficient of da_N, is equal to minus one, we construct the variable $(dp + da_N)$ and also estimate the relationship between the compound variable $(dp + da_N)$ and da_T. The equations to be tested are:

$$dp_{it} = \theta_{0,i} + \theta_{T,i}\left(da_T\right)_{it} + \theta_{N,i}\left(da_N\right)_{it} + \varepsilon_{it} \qquad i = 1,\ldots N;\ t = 1,\ldots T \tag{8}$$

$$\left(dp + da_N\right)_{it} = \theta_{o,i} + \theta_{T,i}\left(da_T\right)_{it} + \varepsilon_{it} \tag{8'}$$

As explained above, we apply here the Wagner and Hlouskova (2004) Non-parametric bootstrapping technique to the tests built by Pedroni (1999, 2004). The methodology is composed of the following steps: first, estimate the possible cointegration relationship between the relevant variables; second, resample the residuals of this relationship 5000 times, by applying the Non-parametric bootstrapping of Wagner and Hlouskova (2004), and use them to calculate the values of the seven statistics of the Pedroni (1999, 2004) test. Third, build the empirical distributions of these values, which will be used to recover the non-biased critical values. Finally, test the null hypothesis of no cointegration with the help of these critical values.

Table 1 shows the results from this test for the two equations applied to the panel data of both groups of countries.

For each panel, the table has seven columns that report the results for the seven statistics of Pedroni (1999, 2004): $T^2N^{3/2}\left(Z_{\hat{v}_{NT}}\right)$, $T\sqrt{N}(Z_{\hat{\rho}_{NT}-1})$, $Z_{t_{NT}}$, $Z_{t_{NT}}^*$, $TN^{-1/2}(\tilde{Z}_{\hat{\rho}_{NT}-1})$, $N^{-1/2}\tilde{Z}_{t_{NT}}$, $N^{-1/2}\tilde{Z}_{t_{NT}}^*$, respectively, and two rows, one for each tested Equation, (8) and (8'), respectively. In each cell, the number in the upper position is the critical value of the corresponding statistic, whereas the number in parenthesis in the lower position is the critical value of the corresponding bootstrapping distribution for significance levels of 5%. Rejection of the null hypothesis is indicated by asterisks for the Pedroni (1999,

[11] To reinforce our conclusions, we also performed the Pesaran (2007) unit root test and the Jönsson (2004) stationary test and derived very similar results. They are not presented here for reasons of space, but are available upon request.

Table 1: First Stage of the BS Hypothesis Cointegration Test of Pedroni (1999, 2004) and Non-Parametric Bootstrapping Applied to this Test

$$dp_{it} = \theta_{0,i} + \theta_{T,i}\left(da_T\right)_{it} + \theta_{N,i}\left(da_N\right)_{it} + \varepsilon_{it} \quad (8)$$
$$\left(dp + da_N\right)_{it} = \theta_{0,i} + \theta_{T,i}\left(da_T\right)_{it} + \varepsilon_{it} \quad (8`)$$

(1991-2004)

	$T^2N^{3/2}\left(Z_{\hat{v}_{NT}}\right)$	$T\sqrt{N}(Z_{\hat{\rho}_{NT}-1})$	$Z_{t_{NT}}$	$Z^*_{t_{NT}}$	$TN^{-1/2}(\tilde{Z}_{\hat{\rho}_{NT}-1})$	$N^{-1/2}\tilde{Z}_{t_{NT}}$	$N^{-1/2}\tilde{Z}^*_{t_{NT}}$
LA							
(8)	-1.316	3.038	0.584	-1.109	4.688	-5.930*	-2.616*
	(-1.027)	(0.227)	(-30.247)	(-14.685)	(2.314)	(-56.001)	(-14.683)
	(-1.916)	(0.733)	(-24.777)	(-12.517)	(2.644)	(-44.488)	(-12.880)
	(-2.340)	(0.995)	(-21.954)	(-11.351)	(2.843)	(-39.912)	(-11.914)
(8')	-2.118	2.428	-0.243	-4.186*	3.757	-2.168*	-4.709*
	(-2.089)	(-1.207)	(-31.537)	(-16.065)	(1.050)	(-51.688)	(-16.382)
	(-2.581)	(-0.613)	(-25.007)	(-13.305)	(1.469)	(-39.527)	(-13.737)
	(-3.051)	(-0.257)	(-22.164)	(-12.114)	(1.718)	(-35.662)	(-12.340)
OECD							
(8)	-2.206	3.938	-1.183	-4.561*	4.521	-8.757*	-3.591*
	(-0.640)	(0.028)	(-33.693)	(-15.611)	(1.888)	(-62.469)	(-17.233)
	(-1.634)	(0.593)	(-26.680)	(-13.047)	(2.382)	(-48.242)	(-14.098)
	(-2.141)	(0.885)	(-23.639)	(-11.774)	(2.664)	(-42.359)	(-12.597)
(8')	-1.164	1.471	-2.903*	-7.296*	2.658	-3.952*	-6.941*
	(-0.618)	(-1.427)	(-34.524)	(-17.063)	(1.136)	(-50.469)	(-16.214)
	(-1.988)	(-0.694)	(-26.879)	(-13.849)	(1.549)	(-40.016)	(-13.595)
	(-2.552)	(-0.320)	(-23.492)	(-12.507)	(1.789)	(-35.526)	(-12.329)

1. The null hypothesis, Ho, is no cointegration.

2. The seven Pedroni (2004) statistics fit a typical normal distribution. The first one is tailed to the right, and the others are tailed to the left. The critical values, at 5% of significance, are 1,645, for the first statistic, and -1.645 for the other ones.

3. The asterisk (*) indicates that the null is rejected at the 5% level of significance.

4. The numbers in parenthesis report the critical values of the bootstrapping distributions at the 5% level of significance, with 5000 resamples. The bold writing indicates that the null hypothesis is rejected under bootstrapping distributions at the 5% of level of significance.

5. The estimations for the two equations include fixed effects.

2004) test, and by bold numbers in the bootstrapping test. As can be seen, in the group of Latin American countries, the null is rejected by three of the traditional Pedroni (1999, 2004) statistics ($N^{-1/2}\tilde{Z}_{t_{NT}}$, $N^{-1/2}\tilde{Z}^*_{t_{NT}}$ and $Z^*_{t_{NT}}$, for the restricted version), and by the bootstrap results applied to $T^2N^{3/2}\left(Z_{\hat{v}_{NT}}\right)$. In the OECD group, the null is rejected even more frequently with the Pedroni (1999, 2004) test ($Z^*_{t_{NT}}$, $N^{-1/2}\tilde{Z}_{t_{NT}}$, $N^{-1/2}\tilde{Z}^*_{t_{NT}}$ and $T\sqrt{N}(Z_{\hat{\rho}_{NT}-1})$ for the restricted version), but only by the bootstrapping test combined with $T^2N^{3/2}\left(Z_{\hat{v}_{NT}}\right)$ and for the restricted version of the model (Equation (8')).

Although the probability of rejecting the null hypothesis decreases with the application of the bootstrapping methodology, the results of Table 1 provide reasonable support for the Baumol-Bowen hypothesis, under the restricted version of the testable equation, since five out seven tests reject the null of no cointegration in each panel[12]. For this reason, we proceed with the estimation of the cointegration vector, corresponding to the restricted equation, in the presence of cross-sectional dependence between the units in the panel. We will work with the homogeneous model, which assumes that all panel members share the same parameters.

4.2 The First Stage of the BS Hypothesis. Cointegration Vectors

Table 2 offers the panel estimates of the parameter θ_T for the restricted equation (Equation (8')), under the assumption that it is shared by all members of the same panel, and two complementary tests to ascertain whether it is significantly different from zero and, subsequently, different from unity. Columns 1 and 3 report the results obtained with the **BAOLS** (*"Bias Adjusted Estimator"*) method designed by Westerlund (2007) and columns 2 and 4 offer, under the heading BAOLS(B), the statistical significance results that we obtain with bootstrapping inference applied to the Westerlund (2007)' method[13].

The first row offers the estimated values of the parameter θ_T, and the value of the *t* statistics appears in parenthesis in the second row. As can be seen, the point estimate of θ_T is statistically significant only in the LA countries, with a value clearly above unity (1.462). Row 3 provides in parenthesis the critical values of the *t* statistics for the null hypothesis $\theta_T = 1$. They indicate that the null can not be rejected in LA.

These results, however, may be affected –and probably biased– by the fact that the sample is relatively small. To unravel the true significance of θ_T in this context, under the null hypothesis $H_0 : \theta_T = 0$, we apply bootstrapping inference to the BAOLS estimator. We use the *Moving Block Bootstrap* algorithm proposed by Li and Maddala (1997) and Li and Xiao (2003), following the Westerlund (2007) methodology, which permits correct inference of the parameters' significance in cases of cross-sectional dependence[14].

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12 In order to reinforce the evidence that there is a long-run relationship between the price differential and average labour productivities in each group of countries, we estimated the error correction for the restricted equation in each panel. The idea is that if the equilibrium error triggers a statistically significant dynamic adjustment in the explained variable towards its long-run equilibrium value, this may be interpreted as evidence of a long-run equilibrium relationship between the explanatory and the explained variables. Results show that the adjustment coefficient has the correct sign and is statistically significant in both groups of countries.

13 The BAOLS method eliminates the bias generated by cross-sectional dependence by using the number of common factors across the members of the panel. To derive the BAOLS estimator, Westerlund (2007) estimated first the cointegration vector with OLS regressions $(\hat{\beta})$ and, subsequently, he calculated the bias of this estimation (b_{NT}) taking into account the optimal number of common factors as suggested by Bai and Ng (2002). Then he defined the BAOLS estimator as $\beta^{+} = \hat{\beta} - \hat{b}_{NT}$, which follows an asymptotically normal distribution.

14 The method consists of the following steps: with the help of the *t* statistics that were previously obtained with BAOLS ($\hat{t}$), we first derive the distribution of the t^* *bootstrap* by applying the *moving block bootstrap* method under the null hypothesis. The t^* bootstrap is obtained after 5000 re-samples. We then calculate the bilateral critical values of this distribution at the $\alpha/2$ significance level. Following Li and Maddala (1997), Li and Xiao (2003), and Westerlund (2007) we calculated critical values for $\alpha = 0.10$, 0.05 and 0.01. The two critical values are designed t_L^* (the left one) and t_R^* (the

Table 2: Estimation of the Cointegration Vector

Homogeneous Model: $\left(dp + da_N\right)_{it} = \theta_{0,i} + \theta_T da_{T,it} + \varepsilon_{it}$

(1991-2004)

	LA		OCDE		
	BAOLS	**BAOLS(B)**	**BAOLS**	**BAOLS(B)**	
	1.462	-	*0.855*	-	
$\hat{\theta}_T$	(1.851)***	(0.802)*	(0.480)	(0.431)**	$\hat{\theta}_T$
	(0.585)	(0.332)**	-	(-0.224)	

1. The first row shows the estimated values of θ_T by the BAOLS method designed by Westerlund (2007). The second row reports –in parenthesis– the value of the *t* statistics ($\hat{t}$) under the null hypothesis $H_0 : \theta_T = 0$, obtained with two alternative methodologies: BAOLS and bootstrapping technique applied to it, BAOLS(B), respectively. The numbers in parentheses in the third row show the *t* statistics under the null hypothesis $H_0 : \theta_T = 1$ using the same two methods. BAOLS *t* statistics follows a normal distribution. The critical values for BAOLS are: +/-2.575(*), +/-1.960(**) and +/-1.645(***) for significance levels of 1%, 5% and 10%, respectively.

2. The BAOLS(B) methodology uses the lower and upper critical values, t_L^* and t_R^*, respectively, of the t^* *bootstrap* distribution, generated with 5000 re-samples for the BAOLS estimator under $H_0 : \theta_T = \theta_0$, for $\theta_0 = 0$ or 1 .

3.for each estimation, the number of common factors is two.

The critical values of the t^* *bootstrap* distribution are shown in columns 2 and 4 under the heading **BAOLS(B)**. Row 2 offers the bootstrapping inferences for the null hypothesis $H_0 : \theta_T = 0$, and row 3 provides the inferences for the null hypothesis $H_0 : \theta_T = 1$. According to these results, the hypothesis that the estimated parameter is statistically equal to zero is rejected in both groups of countries. Furthermore, the hypothesis $\theta_T = 1$ is rejected in LA, confirming the considerably higher than 1 value reported in column 1, but not in the OECD. The fact that the null $H_0 : \theta_T = 0$ is rejected in the OECD panel with the bootstrapping methodology and not with BAOLS deserves a brief explanation. The bootstrapping methodology, which improves the reliability of the tests based on the *t* statistics, generates distributions, which variability may vary from that of the normal distribution. In our case, bootstrapping delivers a *t* statistics empirical distribution in the OECD area with lower variability that leads us to reject the null hypothesis.

To sum up, using recent econometric methods that deal with both cross-sectional dependence and relatively small size of the samples, we find evidence that the first part of the BS hypothesis can not be rejected in each area. Moreover, the estimated parameter θ_T is higher in the group of LA countries than in the OECD economies.

4.3. The Second Stage of the BS Hypothesis. Cointegration Tests

The second stage of the BS hypothesis establishes a relationship between the price differential and the real exchange rate (see Equation 4). Moreover, as explained above, the PPP in the tradable sector (PPP(T)) is the corner stone of this stage. In order to

right one). Finally, we reject the null hypothesis in any of the two following circumstances: $\hat{t} < t_L^*$, or $\hat{t} > t_R^*$.

verify whether this relationship is satisfied, we test here for cointegration relationships by applying the same bootstrapping methodology that was explained and used in Section 4.2 to the equation that links the nominal exchange rate, e, with the price differential in the tradable sector, dp_T. Therefore, we tested this model:

$$e_{it} = \gamma_{0,i} + \gamma_{p,i} dp_{T,it} + \varepsilon_{it} \tag{9}$$

Since in this equation flexibility is assumed in the nominal exchange rate, we excluded from the data of the LA countries the observations for which the nominal exchange rate was fixed with respect to the US dollar. Consequently, to test the BS-2, we dropped the data of Argentina, because this country adopted a currency board with respect to the US dollar during a very large part of the sample (1991-2001), and El Salvador, Panama and Ecuador because these countries used the US dollar as their own currency.

Table 3 shows the results for both groups of countries, assuming the presence of fixed effects (one constant and no trend). Application of the simple Pedroni (1999, 2004) test indicates that the null hypothesis of no cointegration is rejected in both panels with statistics $N^{-1/2}\tilde{Z}_{t_{NT}}$ and $N^{-1/2}\tilde{Z}^*_{t_{NT}}$. Furthermore, in this panel, rejection also derives from $Z^*_{t_{NT}}$. The Non-parametric bootstrapping methodology of Wagner and Hlouskova (2004) also provides evidence of rejection with the first statistic, $T^2N^{3/2}\left(Z_{\hat{v}_{NT}}\right)$[15]. Consequently, we decided to estimate the cointegration vector for the homogeneous version in each panel, and then to test the PPP hypothesis in the tradable sectors.

Table 3: The Second Stage of the BS Hypothesis Cointegration Test of Pedroni (1999, 2004) and Non-Parametric Bootstrapping Applied to this Test

$$e_{it} = \gamma_{0,i} + \gamma_{p,i} dp_{T,it} + \varepsilon_{it}$$

(1991-2004)

	$T^2N^{3/2}\left(Z_{\hat{v}_{NT}}\right)$	$T\sqrt{N}(Z_{\hat{\rho}_{NT}-1})$	$Z_{t_{NT}}$	$Z^*_{t_{NT}}$	$TN^{-1/2}(\tilde{Z}_{\hat{\rho}_{NT}-1})$	$N^{-1/2}\tilde{Z}_{t_{NT}}$	$N^{-1/2}\tilde{Z}^*_{t_{NT}}$
LA							
(9)	-0.425	1.767	0.264	-0.864	2.178	-2.732*	-2.205*
	(-0.940)	(-1.038)	(-27.151)	(-13.859)	(0.940)	(-45.466)	(-14.239)
	(-1.979)	(-0.529)	(-21.348)	(-11.659)	(1.307)	(-35.017)	(-11.827)
	(-2.436)	(-0.191)	(-19.120)	(-10.557)	(1.529)	(-30.644)	(-10.737)
OECD							
(9)	-0.499	2.304	-0.727	-4.251*	2.988	-2.039*	-4.106*
	(0.400)	(-1.676)	(-36.785)	(-18.021)	(0.862)	(-54.663)	(-17.594)
	(-1.201)	(-0.895)	(-28.064)	(-14.517)	(1.405)	(-42.289)	(-14.124)
	(-1.943)	(-0.467)	(-24.554)	(-12.857)	(1.666)	(-36.846)	(-12.666)

1. See the explanations provided under Table 1.
2. Excluding the countries with rigid exchange rates in the LA group: Argentina, Panama, El Salvador and Ecuador.

15 For the same reasons as explained in the cointegration test of the first part of BS, we also estimated the ECM for each panel. Again, the results reveal that there is a statistically significant adjustment coefficient – with the correct sign - which, in turn, presupposes an underlying long-run equilibrium relationship between the nominal exchange rate and the price differential in the tradable sectors of each group of countries.

4.4. The Second Stage of the BS Hypothesis. Cointegration Vector

Following the same procedure that we adopted when testing BS-1 in Section 4.2, we apply here both BAOLS and bootstrapping technique, BAOLS(B), in each group of countries to estimate the cointegration vectors of the equation $e_{it} = \gamma_{0,i} + \gamma_p dp_{T,it} + \varepsilon_{it}$.

The results are reported in Table 4, using the same presentation as in Table 2. For the Latin American group, the estimated value $\hat{\gamma}_p$ is very close to unity (0.952). Moreover, the estimated bootstrap *t** statistics indicates that the null $H_0 : \gamma_p = 0$ can be clearly rejected and that the null $H_0 : \gamma_p = 1$ can not be rejected, with both BAOLS and BAOLS(B) methodologies. Consequently, there is a statistically significant cointegration vector between the nominal exchange rate and the tradables' price differential, which allows us to assert that PPP(T) holds in the set of LA countries with flexible exchange rates during the period of analysis.

Table 4: Estimation of the Cointegration Vector

Homogeneous Model: $e_{it} = \gamma_{0,i} + \gamma_p dp_{T,it} + \varepsilon_{it}$

(1991-2004)

	LA		OCDE		
	BAOLS	**BAOLS(B)**	**BAOLS**	**BAOLS(B)**	
	0.952	-	*0.730*	-	
$\hat{\gamma}_p$	(7.293)*	(0.716)*	(0.463)	(0.706)	$\hat{\gamma}_p$
	(-0.368)	(-0.370)	-	-	

1. See the explanations provided in Table 2.

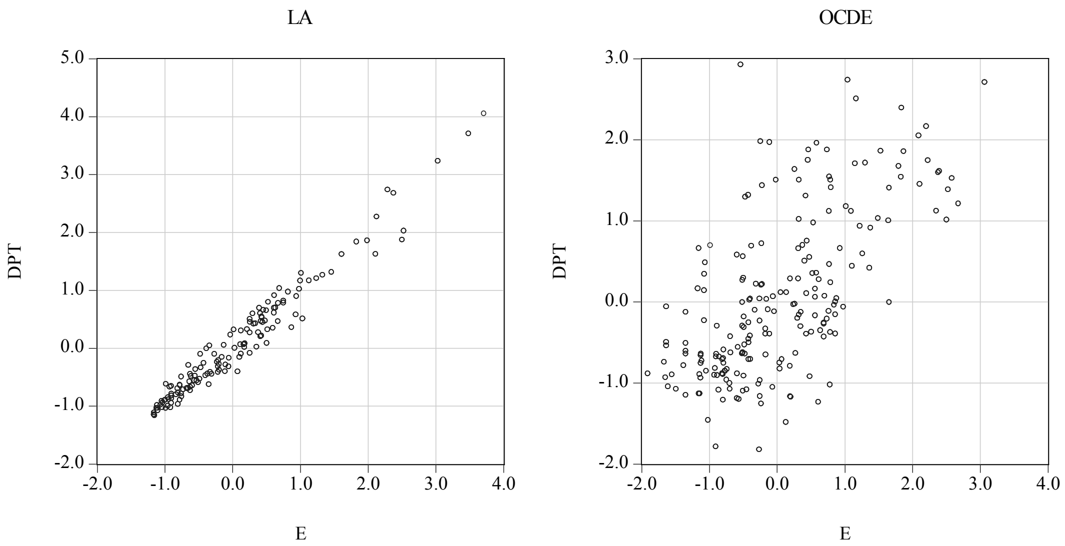
2. The BAOLS(B) methodology uses the lower and upper critical values, t^*_L and t^*_R, respectively, of the t^* *bootstrap* distribution, generated with 5000 resamples for the BAOLS estimator under $H_0 : \gamma_p = \gamma_0$, for alternatively $\gamma_0 = 0{,}1$. The null hypothesis is rejected in either of the two cases: $\hat{t} < t^*_L$ or $\hat{t} > t^*_R$

3. For both estimations the number of common factors is two.

For the OECD group, the null $H_0 : \gamma_p = 0$ can not be rejected on the basis of the statistics provided by each econometric methodology: BAOLS and BAOLS(B), which implies that there is no room for testing the null $H_0 : \gamma_p = 1$. Consequently, there is no clear long-run relationship between the nominal exchange rate and the price differential in the tradable sectors, and PPP(T) can not be accepted in the OECD panel. The fact that the estimated value of parameter γ_p is not significantly different from zero is consistent with a random behaviour of the RER(T) within the OECD group of countries.

The different results obtained for the two panels are backed by the two graphs presented in Figure 2. This Figure shows a cross-plot of the tradables price differential and the nominal rate variables using annual observations. The graph in the left plots the points of the LA countries, and the graph in the right represents the OECD observations. If PPP(T) holds, the cross-plot should be randomly scattered around the 45º line defining the equilibrium position. As can be seen, this condition is approximately fulfilled in the LA group, but it is clearly not satisfied in the OECD panel.

Figure 2: A Cross-Plot of the Tradables Price Differential (DPT) and the Nominal Exchange Rate (E) with Respect to the US in Both Groups of Countries (in logs) 1991-2004

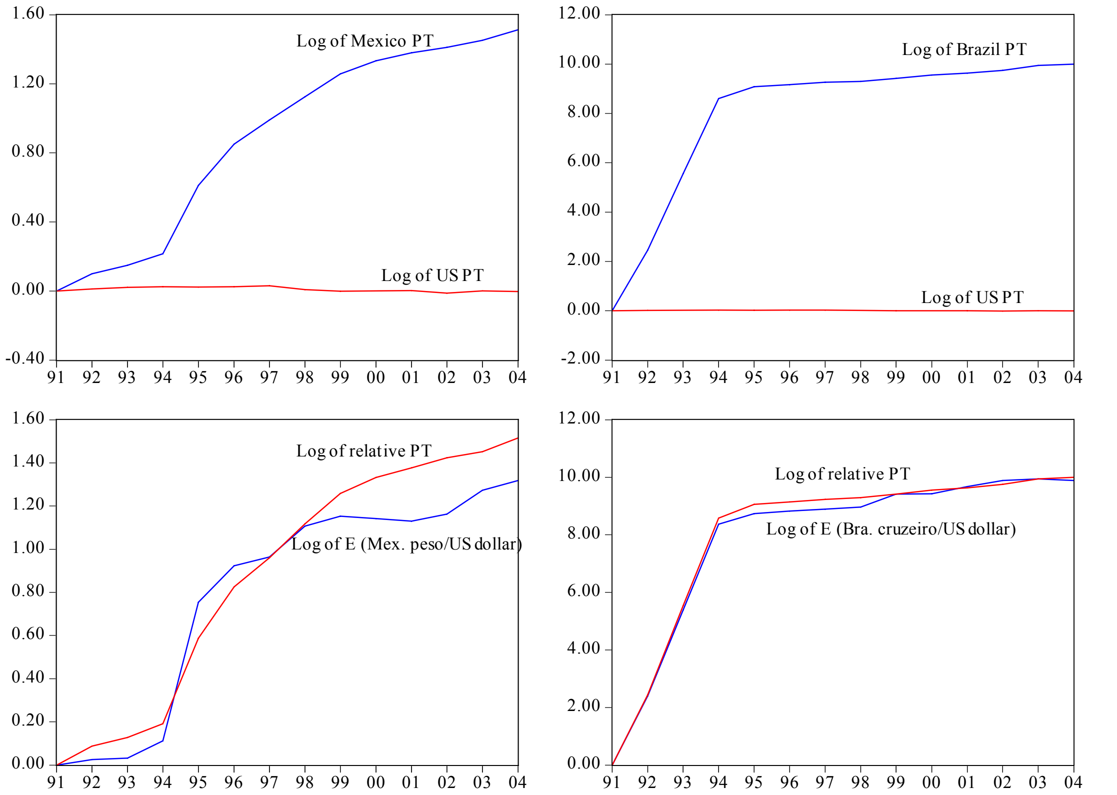


Let us now dig deeper in the analysis to find the factors that may account for the different outcomes in the two areas. We start by representing graphically the key variables of the tests for some representative countries of each panel. Due to reasons of space and only for illustrative purposes, we choose the two economically biggest countries of each group with flexible exchange rates with respect to the US dollar: Mexico and Brazil for the LA group, and the UK and Germany for the OECD panel. Thus, Figure 3 plots the time observations of tradable prices – domestic country and the US - in the upper panels and their relative prices and the nominal exchange rates in the lower panels. The graphs in the left side of this figure correspond to Mexico, and those in the right side represent Brazil. Figure 4 shows the same information for the two OECD representatives: the UK and Germany.

We first analyse the results in the two LA countries, depicted in Figure 3. The upper panels show that whilst the US P_T index has followed a rather stable trajectory, the two LA P_T indices have noticeably grown over the whole period, resulting in an upward sloping stochastic trend in relative tradable prices. The lower panels show that the nominal exchange rates reflect these upward sloping trends with no apparent long departures from them. The absence of long and lasting swings in the nominal exchange rates indicates that these variables have not been affected by cyclical speculative pressures. In any case, the close correspondence between variations in tradables' relative prices and in nominal exchange rates in each country explains why the estimated parameter γ_p of Equation (9) is very close to unity, and justifies the stationarity of the RER(T) in the LA panel.

Figure 3: The Time Graphs of Tradable Prices (Upper Panel) and their Relative Prices and Nominal Exchange Rates (Lower Panel)
(1991-2004)

Mexico **Brazil**

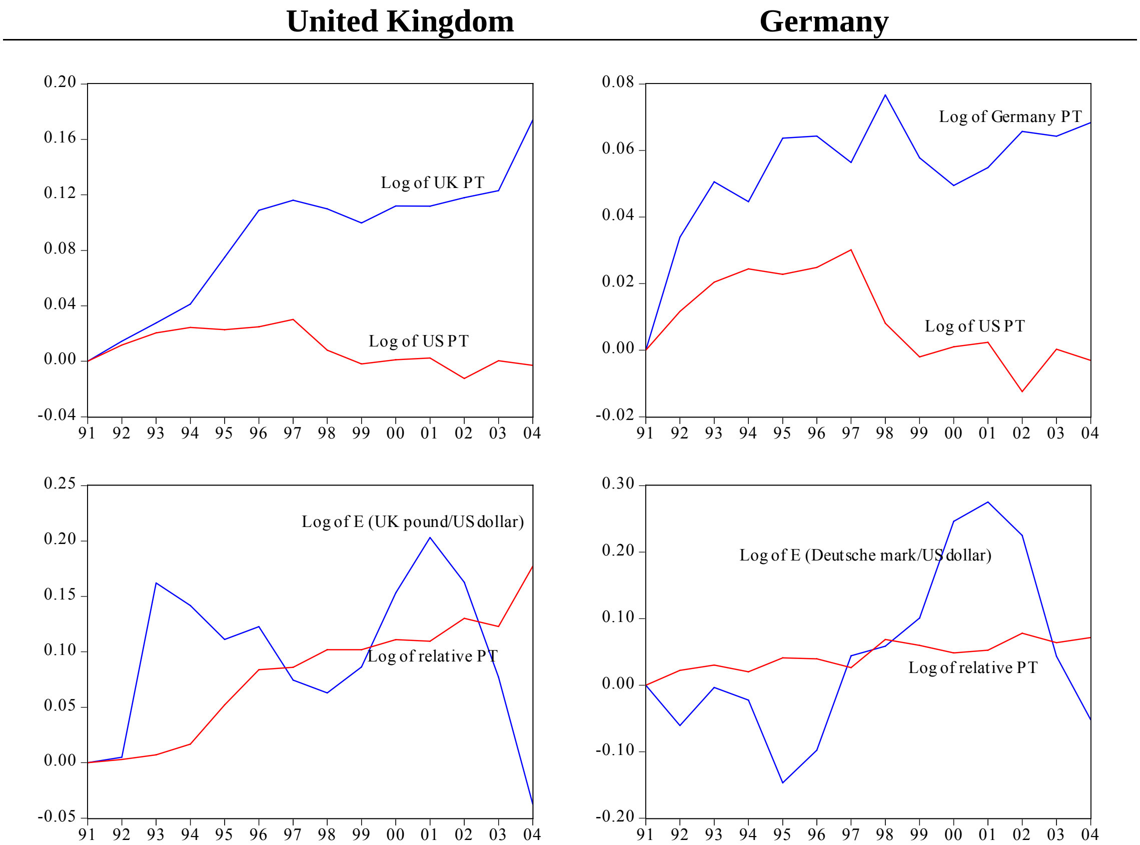


These results reveal that the LA economies, considered small countries in economic terms, behave as price takers – in foreign currency - in international markets of tradable goods. Correspondingly, variations in P_T are rapidly matched by variations in E, and adjustments in E are rapidly passed through onto variations in P_T. Indeed, a number of recent empirical studies find that the exchange rate pass-through (ERPT) ratio onto the prices of domestic tradables, and particularly to import prices, reaches high levels in emerging market economies. Barhoumi (2005), for instance, calculated that this ratio ranges between 77% and 83%, whilst Frankel, Parsley and Wei (2005) pointed out that the ratio in emerging market economies is almost four times as high as it is for developed countries. The contributions of Burstein, Eichenbaum and Rebelo (2002, 2005) confirm these general results, remarking that ERPT is complete (100%) when import prices are measured at the docks. It is worth noting that fulfilment of PPP(T), caused by an almost complete pass-through, is due to certain pricing behaviour of trading firms, and could be compatible with the presence of transportation costs[16].

[16] The explanation of these empirical results lies in the facts that: a) emerging market economies are usually price takers in international goods markets, and that b) foreign firms set prices in their own currency, or preponderantly in US dollar (producer currency pricing). Consequently, foreign firms do not modify their prices expressed in their own

Let us now examine the results of the two OECD countries, represented in Figure 4. The US P_T index exhibits stability compared with the upward trend of tradables' prices of the United Kingdom and Germany over the period. Consequently, the relative price line shows an upward trend, which is more pronounced in UK than in Germany. The difference with respect to the LA countries case is that nominal exchange rates of the OECD countries move in a contrary direction to relative prices for extended periods of time, describing swings of variable intensity and duration. This feature, which is very general in the OECD group of countries, explains why the data do not support PPP(T) in this area, as was also found in a number of previous studies, such as Søndergaard (2001), Engel (2002) and García-Solanes, Sancho and Torrejón (2008).

Figure 4: The Time Graphs of Tradable Prices (Upper Panel) and their Relative Prices and Nominal Exchange Rates (Lower Panel)
(1991-2004)



Frydman et al (2008) provide a convincing explanation of the factors that generate pronounced and persistent movements in the nominal exchange rates in industrial countries, which are ultimately responsible for the lasting departures from PPP. They attributed these movements to speculative transactions in the foreign exchange market,

currency after nominal exchange rate variations, which leads to complete ERPT to importing prices in emerging market economies. This pricing behaviour allows the foreign firms to maintain their mark-up constant.

created by agents' imperfect knowledge about the underlying model. Juselius (2008) analysed the PPP puzzle in the German mark/US exchange rate by grouping together components of similar persistence in the frame of a Cointegrated VAR approach, and found, indeed, that the German CPI and the nominal exchange rate, respond very slowly to shocks.

It seems reasonable to borrow the findings and arguments of Frydman et al (2008) and Juselius (2008) to account for the forces that feed the long swings of the nominal exchange rates in Figure 4. However, since we work with tradable prices, instead of consumer prices, we need some *complementary* factors to explain why RER(T) is not stationary - the PPP puzzle in the **tradable** sectors. In fact, long movements in the RER(T) indicate that potential profits which are easily visible and free of exchange rate risks are not realised by arbitrageurs and traders. Consequently, we believe that there is a complementary reason for the pronounced movements in the RER(T) not being rapidly eroded in the OECD panel. This reason is market segmentation of national tradable markets, probably created by both arbitrage frictions and imperfect competition.

As a synthesis of the empirical part of this paper, we may assert that in the Latin American group the BS hypothesis holds in the area as a whole despite the fact that capital is not completely mobile between countries of this area and the USA, as assumed by the BS model. By contrast, in the OECD group, the entire BS hypothesis does not hold in the whole area due to PPP failure in the tradable sectors of those countries with respect to the USA. These econometric results seem to confirm our first impressions derived from the descriptive analysis in Section 3.2, and they agree with the main findings of previous studies on BS effect devoted to LA countries, such as Drine and Rault (2003).

5 Concluding Remarks

The literature testing the Balassa and Samuelson hypothesis provides different results, depending on the degree of economic development of the countries analysed with respect to a foreign developed country. Thus, whereas some studies show that the BS hypothesis tends to be satisfied in groups of countries lagging considerably behind the USA, other works obtain very poor results in areas with similar standards of living to that country. In this paper we test the BS effect by looking at two areas differing substantially in development and growth: sixteen OECD countries, on the one hand, and sixteen Latin American economies, on the other hand. We take the U.S. as the benchmark country. In order to detect the origin of possible failures, we split the BS hypothesis into two parts and subject them to individual scrutiny. We use pooled observations and apply recent econometric panel techniques to overcome the problems of insufficient data in many countries and cross-sectional dependence in the data of our samples –especially in the case of the OECD countries.

We find evidence for not rejecting the first stage of the hypothesis, which links the difference between the productivities with the difference in prices of the tradable and non-tradable sectors, in each group of countries. The same conclusion applies in the Latin American area with respect to the second stage, which relates the price differential with the real exchange rate. Nonetheless, this stage is rejected in the group of the OECD countries as a whole. The failure is reflected in departures from PPP in the tradable sectors, and is probably due to two groups of factors: a) speculative behaviour of economic agents generally supported by imperfect knowledge of the underlying causing mechanisms, and b) non-competitive practices and arbitrage frictions that still prevail in the countries of this area.

The likely fulfilment of the BS hypothesis in the whole Latin American sample, suggested by our descriptive analysis, and corroborated by our empirical tests, has some exchange-rate-policy implications. Since the countries of this area are frequently hit by asymmetric shocks and their long-term economic growth experiences noticeable upheavals with respect to the USA economy, their equilibrium RER against the US dollar must adjust accordingly. If the nominal exchange rate is pegged to the US dollar or is maintained rigidly stable around this currency, the volatility in the RER will convey high variability in domestic CPI inflation rates[17]. Difficulties are particularly severe in cases where negative supply shocks and slow growth episodes impose disinflation efforts in the countries of the LA area. Under such situations, national authorities might feel compelled either to maintain very restrictive monetary and fiscal policies to beat down inflation, or to allow overvaluation in the real exchange rate. Both outcomes harm growth and employment. The solution to avoid these negative results would be to permit flexibility in the nominal exchange rate, as a weapon to absorb external shocks, as was emphasised by Edwards and Levy-Yeyati (2003)[18]. Very recent studies, such as Calderón and Schmidt-Hebbel (2003) and García-Solanes and Torrejón (2007) prove that the usefulness of flexible exchange-rate regimes is magnified in the LA area when the accompanying monetary policies are guided by inflation targeting strategies.

[17] Apart from the (equilibrium) long term adjustments imposed by BS effects and other permanent real factors, the RER may experience short-run fluctuations as a result of nominal rigidities.

[18] As outlined in this work, the ability of flexible exchange rates to absorb external shocks increases with the extent and speed at which fluctuations in these variables are passed through to prices of tradable goods compared to prices of non-tradable goods.

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