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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 hypothesis in two areas with strong differences in economic development, sixteen OECD countries and sixteen Latin American economies. Applying panel cointegration and bootstrapping techniques that solve for cross-sectional dependence problems in the data, we find that the second stage of the hypothesis, which relates relative sector prices with the real exchange rate, only holds in the Latin American area. The failure of the latter in the OECD countries as a whole is reflected in departures from PPP in the tradable sectors, and is probably due to segmentation between national tradable markets.

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SUMMARY

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 panel techniques to overcome the problems of cross-sectional dependence in the data of our samples –especially in the case of the OECD countries–.

We find that while 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, is satisfied in each group of countries, the second stage, which relates the price differential with the real exchange rate, holds in the Latin American area, but not in the group of the OECD countries as a whole. Fulfilment of the second stage in LA countries is favoured by the large extent to which exchange rate variations pass-through on import prices in this group of countries – as emphasised by recent empirical studies -. In the OECD group, the failure is reflected in departures from PPP in the tradable sectors, and is probably due to transportation costs and non-competitive practices that still prevail in the countries of this area. Putting together the results for all the individual countries of our sample, it follows that the entire BS hypothesis clearly holds in five Latin American countries and perhaps in three OECD economies.

I INTRODUCTION[1]

The Balassa and Samuelson (BS) hypothesis (Balassa (1964), Samuelson (1964)) provides 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 apply 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 with respect to Germany since the early nineties (Halpern and Wyplosz (2001), Kovács (2002), Égert (2002a,b), Mihaljek and Klau (2004), Égert et al. (2002)). 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 countries analysed.

The empirical findings referring to economies that do not exhibit pronounced divergences in economic development between them, 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), 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 frank/US dollar exchange rate.

[1] We are grateful to the financial support by the Spanish Ministry of Education, Project SEJ 2006-15172. We also thank Enrique Alberola-Illa and Josep Lluís Carrion-i-Silvestre for their helpful comments on our empirical methodology.

Despite the fact that statistical significance of the empirical results seem sensitive to the level of economic development of the areas analysed, to our knowledge 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-1hereinafter) relates the difference in productivities with the difference in prices of the tradable and non-tradable sectors. The second (BS-2) establishes the link between the price differential and the real exchange rate measured with CPI deflators. This second relationship is immediately obtained by assuming that 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 Canzonery, Cumby and Diba (1999).

This paper presents two novelties that, in our opinion, contribute to improve the empirical results. First, we 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 calculate the variables of interest more accurately. Second, since cross-sectional dependence in 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 U.S. variable) – we apply unit root and cointegration tests, as well as bootstrapping techniques, which have been developed very recently to cope wit this problem. We use annual observations from the period 1991-2004.

As an advance of our findings, we obtain very satisfactory results for the first stage of the BS hypothesis in both groups of countries considered. 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. We do not find evidence to show that estimates of the BS-1 are better in one group than in the other. However, things look very different in the tests of the second part of the hypothesis (BS-2). Here we find that PPP holds for the tradable sectors of the Latin American countries as a whole but not for the group of OECD countries. Furthermore, when looking at individual members, we find that BS-2 is verified in more cases inside the Latin American group than in the OECD area.

The failure of PPP in the tradable sectors in developed areas is not surprising on theoretical or empirical grounds. The New Open Macroeconomics literature provides theoretical reasons, based on transportation costs, non-competitive practices and pricing-to-the market behaviour of exporters, and some empirical works have already found results along the same lines, as explained below in this paper. On the order hand, very recent empirical literature shows that acceptable fulfilment of PPP in the tradable sector of emerging market economies is a natural result, given the large extent to which variations in the nominal exchange rate are passed to import prices in these countries.

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.

II THEORETICAL FRAMEWORK

2.1 The Balassa and Samuelson model

To test the BS model for a pair of countries and check more easily what the causes of success or failure are, we follow the two-step procedure of Canzoneri, Cumby and Diba (1999). We then split the model in two parts, which we name BS-1 and BS-2. We analyse each part separately.

The **first part of the BS hypothesis** links the difference in *total* productivities with the difference in prices of tradable (T) and non-tradable (N) sectors. Under the usual assumptions of factors mobility and perfect competition, and assuming that sectoral aggregate productions are governed by Cobb-Douglas functions in each country, it is easy to derive[2]:

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

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 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 is 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 tested. The testable equation is:

$$dp = \left(alp_T - alp_T^*\right) - \left(alp_N - alp_N^*\right) \qquad (2)$$

[2] See, for instance, Égert et al. (2005)

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 two important peculiarities: first, the coefficients of the labour productivities are all equal to unity and, second, it is 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:

$$q = (e + p_T^* - p_T) - \lambda.dp \qquad (3)$$

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 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:

$$q = -\lambda.dp \qquad (4)$$

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**:

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

In terms of first differences, it would indicate 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. It is worth mentioning that, since the BS model is postulated in terms of real exchange rates, its main theoretical propositions apply in any nominal exchange-rate regime and may be tested in contexts of fixed and/or floating exchange rates.

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 provide evidence against PPP(T), using different statistical and econometric methods and different geographical samples.

To gain further insight into the sources of PPP(T) failure, we split the RER(T) into three components following a simple accounting procedure[3]:

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

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:

$$\tau = e + p_F^* - p_H \quad (7)$$

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))[4]. The stimulus in the demand for the tradables produced at home (García-Solanes, Sancho and Torrejón (2007)) also adds appreciating pressure on the terms of trade[5].

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 the same national tradable goods having the same price across markets. In terms of expression (6), this circumstance is reflected in that the third and fifth parentheses are significantly different from zero. Market segmentation may be due to two causes: a) *imperfect competition*, which frequently gives rise to "pricing-to-market" practices[6] (Krugman (1987)), and b) *arbitrage frictions*, created notably by transportation costs (Rogoff (1996)), information costs and non-tariff barriers.

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[3] See García-Solanes, Sancho and Torrejón (2008) for a detailed derivation.
[4] 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.
[5] 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)).
[6] Goldberg and Knetter (1997) survey the sources of "pricing-to-market" policies.

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 low 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[7]. Maier (2004) stressed the fact that the width of the non-arbitrage bands increases with exchange rate variability.

From the preceding paragraphs, it is easy to understand that quality variations and market segmentation inflict different trajectories to the RER(T). If continuous quality improvements coupled with demand pressures on tradables are the guiding 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.

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

III 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. All the series are transformed into natural logarithms and then converted into indices, with the first year of the sample being the base. The panel data set covers two groups of countries: 16 Latin 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 USA economy as the benchmark foreign country, since all the countries mentioned have substantial economic exchanges with this economy[8].

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

[7] 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.

[8] Although in the group of OECD countries mutual trade flows and economic relations are predominant, the adoption of an external country –the USA- as the reference can be justified by the fact that the OECD economies do not share the same economic business cycle and maintain differentiated trade exchanges with the North American economy.

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. The task is not straightforward because no consensus exists on this issue. In the tradable sector we include all the tradable economic activities specified in the official statistics, excluding agriculture. As in many other empirical analyses, we exclude agricultural activities from the classification 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 countries of this group 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 amount. Our decision adheres to the fact that 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 the 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.

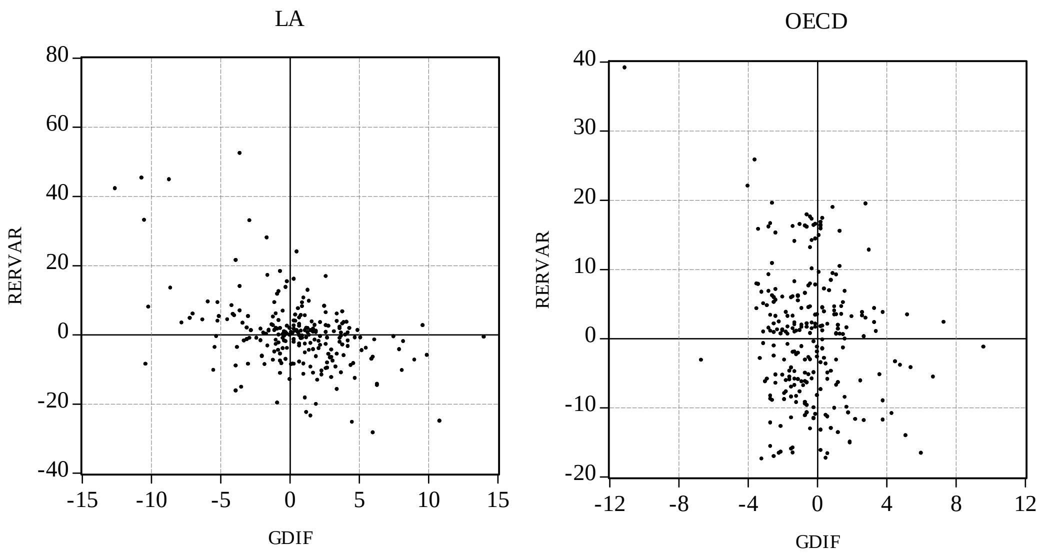
We define the relative price of non-tradable goods with respect to tradable ones as the ratio of the two corresponding sector GDP deflators. To obtain deflator indices we measured the aggregate production, that is, the value added, of each sector in both nominal and real terms, using current prices and 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 is expected that countries growing faster will tend to experience real exchange-rate appreciations with respect to other, slow growing economies. To verify this in a simple and descriptive way, in Graph 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 in 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. 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.

GRAPH 1
Growth differential (GDIF) and variations in the real exchange rate (RERVAR) in the two groups of countries. Annual observations (1991-2005)



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

IV ECONOMETRIC ANALYSIS

In this section, we apply *panel* stationary and cointegration techniques to test the two stages of the BS hypothesis in the two areas under study, since this methodology, based on pooled observations, increases the accuracy in the inference of the existence of unit roots and cointegration (Banerjee (1999)). Moreover, they raise the reliability of the estimates of the cointegration vector, especially when the observed period is relatively short.

However, the correct application of these techniques depends crucially on the assumption that individual time series are cross-sectional independently distributed. Since this assumption is not likely to be satisfied within groups of countries that maintain important economic links, new tests and statistics (Jönsson (2004), Bai and Kao (2005), Pesaran (2007), Banerjee and Carrión-i-Silvestre (2007), Westerlund (2007)) have been developed recently to handle this type of data dependence[9]. In fact, in the presence of cross-sectional dependence, the traditional panel unit roots, stationarity and cointegration tests tend to over-reject the null hypothesis.

As a first approximation to evaluate the degree of cross-sectional dependence in our data, we estimated individual ADF(p) regressions for p = 1, 2, 3, 4 and computed pair-wise cross section correlations coefficients of the residuals from these regressions. The simple mean of these correlation coefficients together with the associate cross-sectional dependence test statistic proposed in Pesaran (2004) showed some degree of cross-sectional dependence in all series, since we usually rejected the null hypothesis of zero cross-sectional dependence. In the LA countries correlations were 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.

Before performing the cointegration tests, we applied both panel unit-root and stationary tests that are well suited to handle with cross-sectional dependence (Pesaran (2007) and Jönsson (2004), respectively) to the following variables: $dp = \left(p_N - p_N^*\right) - \left(p_T - p_T^*\right)$, $da_T = alp_T - alp_T^*$, $da_N = alp_N - alp_N^*$, e and $dp_T = p_T - p_T^*$.

The results suggest that each of the variables contains one unit root in the two panels of our study[10], which justifies further investigation into whether the variables maintain the long run relationships derived from our model. 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 dp, da_T and da_N. 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, consequently, estimate the relationship between the compound variable $(dp + da_N)$ and da_T. The equation to be tested is:

$$\left(dp + da_N\right)_{it} = \theta_{oi} + \theta_{Ti}(da_T)_{it} + \varepsilon_{it} \tag{8}$$

[9] Wagner (2005), Westerlund and Basher(2006) and García-Solanes, Sancho and Torrejón (2007) applied these new econometric techniques to test PPP in the presence of cross-sectional dependence in several groups of countries.

[10] We do not present here these results for reasons of space, but they are available upon request.

We apply the test described in Banerjee and Carrion-i-Silvestre (2007) to check the null hypothesis of no cointegration. This test generalises the class of panel cointegration tests proposed by Pedroni (1999, 2004) to allow for cross-sectional dependence, which is treated using common factors like in Bai and Ng (2002) [11]. Table 1 shows the results from this test for both groups of countries, taking into account the possible existence of a trend in the long run relationship.

The results allow us to reject the null hypothesis of no cointegration between the series in all cases (constant and trend) in the two groups of countries. The p-values are, indeed, highly statistically significant. Consequently, we may assert that there is a long-run relationship between the variables dp , da_N and da_T in each group of countries.

TABLE 1
First stage of the BS hypothesis
Cointegration test of Banerjee and Carrion-i-Silvestre (2007)

$$\left(dp + dal_N\right)_{it} = \theta_{oi} + \theta_{Ti}(dal_T)_{it} + \varepsilon_{it}$$

(1991-2004)

	LA		OECD	
Statistics	**Constant**	**Trend**	**Constant**	**Trend**
$\dfrac{N^{-1/2} Z_{\hat{t}_{NT}}^{\hat{e}^j} - \Theta_2^{\hat{e}^j}\sqrt{N}}{\sqrt{\Psi_2^{\hat{e}^j}}}$	-5.969 [0.000]	-8.938 [0.000]	-5.074 [0.000]	-8.577 [0.000]
Number of factors	1	1	2	2

1. The null hypothesis is no cointegration. p-values are shown within brackets.
2. $Z_{\hat{t}_{NT}}^{\hat{e}^j}$ corresponds to the un-normalised cointegration test. The values for $(\Theta_2^{\hat{e}^j}, \Psi_2^{\hat{e}^j})$ are obtained from Banerjee and Carrion-i-Silvestre (2007), Table 3.
3. Critical values are given for the left tail of the normal distribution
4. To select the number of factors, we follow Bai and Ng (2002).

Given that there is a cointegration relationship between the series, we estimate the cointegration vector in the presence of cross-sectional dependence between the units in the panel. We consider two alternative cases: we first assume that each individual country has its own (differentiated) parameters, which will be revealed by the estimation results (heterogeneous model); then we assume that all panel members share the same parameters (homogeneous model).

4.2 The first stage of the BS hypothesis. Cointegration vectors

4.2.1 The heterogeneous model

Table 2 offers the individual estimates of the parameter θ_{Ti} and two complementary tests to ascertain whether it is significantly different from zero and, subsequently, different from unity. Columns 1 and 3 report, for individual countries of each group, the estimates obtained with the FMOLS method proposed by Phillips and Hansen (1990), whose *t* statistics tends to an asymptotically normal distribution. Estimations of the parameter θ_{Ti} are presented in the first row of the two columns, and the value of the *t* statistics appears in parentheses in the second row. As can be seen, estimations are

[11] As in Banerjee and Carrion-i-Silvestre (2007), we focus only on the parametric statistics.

TABLE 2
Estimation of the cointegration vector. Heterogeneous model:

$$\left(dp + da_N\right)_{it} = \theta_{oi} + \theta_{Ti}\left(da_T\right)_{it} + \varepsilon_{it}$$

(1991-2004)

$\hat{\theta}_{Ti}$	LA		OECD		$\hat{\theta}_{Ti}$
	FMOLS	**FMOLS(B)**	**FMOLS**	**FMOLS(B)**	
	3.045	-	*0.948*	-	
Argentina	(10.548)*	(5.276)*	(8.463)*	(5.172)*	Australia
	(7.084)*	(5.101)*	(-0.468)	(-3.284)	
	1.372	-	*1.048*	-	
Bolivia	(4.780)*	(3.333)**	(0.983)	(3.442)	Austria
	(1.295)	(3.270)	-	-	
	1.762	-	*0.667*	-	
Brazil	(12.458)*	(5.639)*	(1.762)***	(3.285)	Belgium
	(5.388)*	(4.517)*	(-0.880)	(-3.304)	
	-0.418	-	*1.910*	-	
Chile	(-0.308)	(-2.886)	(8.285)*	(3.227)**	Canada
	-	-	(3.948)*	(3.249)**	
	1.307	-	*-0.339*	-	
Colombia	(3.500)*	(3.234)**	(-0.200)	(-3.074)	Denmark
	(0.823)	(3.113)	-	-	
	0.669	-	*0.992*	-	
Costa Rica	(0.510)	(3.594)	(1.517)	(3.415)	France
	-	-	-	-	
	0.135	-	*1.942*	-	
Dominican Rep.	(0.427)	(3.440)	(3.542)*	(3.320)**	Germany
	-	-	(1.718)***	(3.423)	
	1.201	-	*0.528*	-	
Ecuador	(4.807)*	(3.645)**	(1.486)	(2.977)	Ireland
	(0.805)	(3.386)	-	-	
	1.989	-	*1.126*	-	
El Salvador	(4.577)*	(3.355)**	(8.771)*	(3.300)**	Italy
	(2.276)**	(3.466)	(0.979)	(3.267)	
	1.253	-	*0.175*	-	
Jamaica	(3.387)*	(2.824)***	(1.924)***	(2.967)	Korea
	(0.684)	(3.304)	(-9.056)*	(-4.441)*	
	1.381	-	*1.411*	-	
México	(11.242)*	(4.734)*	(4.040)*	(3.269)**	The Netherlands
	(3.103)*	(2.571)***	(1.177)	(3.271)	
	1.438	-	*3.771*	-	
Nicaragua	(3.366)*	(3.129)**	(2.041)**	(3.337)	Norway
	(1.025)	(3.113)	(1.500)	(3.172)	
	0.884	-	*0.876*	-	
Panama	(3.463)*	(3.323)**	(3.394)*	(2.805)***	Portugal
	(-0.453)	(-3.248)	(-0.481)	(-3.433)	
	0.478	-	*1.067*	-	
Paraguay	(1.439)	(3.354)	(6.664)*	(5.048)*	Spain
	-	-	(0.421)	(3.084)	
	1.018	-	*0.770*	-	
Uruguay	(1.012)	(3.585)	(5.576)*	(5.299)*	Switzerland
	-	-	(-1.670)***	(-3.719)	
	1.648	-	*0.284*	-	
Venezuela	(8.820)*	(4.829)*	(1.294)	(3.387)	United Kingdom
	(3.469)*	(3.104)**	-	-	

1. The first row shows the estimated values of θ_{Ti} by FMOLS. The second row reports – within parentheses – the value of the *t* statistics ($\hat{t}$) under the null hypothesis $H_0 : \theta_{Ti} = 0$, obtained with the two alternative methodologies, FMOLS and FMOLS(B). The numbers in parentheses in the third row show the *t* statistics under the null hypothesis $H_0 : \theta_{Ti} = 1$ obtained with FMOLS and FMOLS(B). The FMOLS *t* statistics follows a normal distribution. The critical values for FMOLS are: +/-2.575(*), +/-1.960(**) and +/-1.645(***) for significance levels of 1%, 5% and 10%, respectively.

2. The FMOLS(B) methodology uses the lower and upper critical values, t_L^* and t_R^*, respectively, of the t^* *bootstrap* distribution generally at 5%, generated with 5000 resamples for the FMOLS estimator under $H_0 : \theta_{Ti} = \theta_0$, for $\theta_0 = 0$ or 1 . The null hypothesis is rejected in either of the two cases: $\hat{t} < t_L^*$ or $\hat{t} > t_R^*$.

3. Values of the *t* statistics under $H_0 : \theta_{Ti} = 1$ for FMOLS are not reported when they are not significant.

significantly different from zero in 11 LA countries of the 16 members of the sample, and in 11 OECD countries of the 16 components of this group.

These results, however, are affected – and probably biased - by the assumption of cross-sectional independence in the data. To unravel the true significance of θ_{Ti} in this context, under the null hypothesis $H_0 : \theta_{Ti} = 0$, we apply bootstrapping inference to the FMOLS estimator.

The method consists of the following steps: with the help of the *t* statistics that were previously obtained with FMOLS ($\hat{t}$), 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 resamples. We then calculate the bilateral critical values of this distribution at the $\alpha / 2$ level of significance. Following Li and Maddala (1997) and Li and Xiao (2003), we adopted the value $\alpha = 0.05$. The two critical values are designed t_L^* (the left one) and t_R^* (the right one). Finally, we reject the null hypothesis in any of the two following circumstances: $\hat{t} < t_L^*$, or $\hat{t} > t_R^*$. The critical values of the t^* *bootstrap* distribution at 10%, 5% and 1% levels of significance, under the null hypothesis that: $H_0 : \theta_{Ti} = 0$ are shown in parentheses in the second row of columns 2 and 4 under the heading FMOLS(B). In the case of Argentina, for example, since $\hat{t} = 10.548$, and $t_R^* = 5.276$, we may reject the null hypothesis and accept that θ_{Ti} is significantly different from zero.

The estimated parameter is statistically different from zero, at least at the 10% level, in **eleven LA** countries and **eight OECD** economies.

Columns 2 and 4 also report, in the third row in parenthesis, the critical values of the t^* *bootstrap* distribution at the 5% level of significance, under the null hypothesis that $\theta_{Ti} = 1$. The null hypothesis is not rejected in seven LA countries (Bolivia, Colombia, Ecuador, El Salvador, Jamaica, Nicaragua and Panama) and seven OECD members (Australia, Germany, Italy, The Netherlands, Portugal, Spain and Switzerland). All these countries satisfy the *strong* version of the first step of the BS hypothesis. We could say that countries for which θ_{Ti} is significantly positive, but not necessarily equal to unity, satisfy a weak version of BS-1.

4.2.2 The homogeneous model

We estimate now the parameter θ_T under the assumption that it is shared by all members of the same panel and we report the results in Table 3. The values presented in the first row of columns 1 and 3 show that the point estimate is higher than unity in the LA countries and lower than unity in the OECD group, which indicates that the sensitivity of the relative price differential to the relative productivity increase in the tradable sector is higher in the first area than in the second one. According to the *t* statistics provided in parentheses in the second row, the null hypothesis that $\theta_T = 0$ is clearly rejected –at the 1% significance level– in both groups of countries, even when the presence of cross-sectional dependence in the data is taken into consideration with bootstrapping inference - FMOLS(B) results presented in columns 2 and 4.

Row 3 reports the *t* statistics for the null hypothesis $H_0 : \theta_T = 1$. As can be seen, the null hypothesis can be rejected in the LA panel, but not in the OECD group, with both estimation techniques. The estimations in the latter have a relatively large standard error within a wide confidence interval, which increases the probability of accepting the null hypothesis.

TABLE 3
Estimation of the cointegration vector
Homogeneous model: $(dp + da_N)_{it} = \theta_{0i} + \theta_T da_T + \varepsilon_{it}$
(1991-2004)

	LAT		OECD		
	FMOLS	**FMOLS(B)**	**FMOLS**	**FMOLS(B)**	
$\hat{\theta}_T$	*1.402*	-	*0.866*	-	$\hat{\theta}_T$
	(12.155)*	(6.726)*	(7.504)*	(5.313)*	
	(3.488)*	(2.945)*	(-1.163)	(-1.913)	

1. The first row shows the estimated values of θ_T by the FMOLS method designed by Kao and Chiang (2000). The second row reports – in parentheses– the value of the *t* statistics ($\hat{t}$) under the null hypothesis $H_0 : \theta_T = 0$, obtained with the two alternative methodologies, FMOLS and FMOLS(B), respectively. The numbers in parentheses in the third row show the *t* statistics under the null hypothesis $H_0 : \theta_T = 1$ obtained with FMOLS and FMOLS(B). The FMOLS *t* statistics follows a normal distribution. The critical values for FMOLS are: +/-2.575(*), +/-1.960(**) and +/-1.645(***) for significance levels of 1%, 5% and 10%, respectively.
2. The FMOLS(B) methodology uses the lower and upper critical values, t_L^* and t_R^*, respectively, of the t^* *bootstrap* distribution generally at 5%, generated with 5000 resamples for the FMOLS estimator under $H_0 : \theta_T = \theta_0$, for $\theta_0 = 0$ or 1. The null hypothesis is rejected in either of the two cases: $\hat{t} < t_L^*$ or $\hat{t} > t_R^*$.
3. Values of the *t* statistics under $H_0 : \theta_{Ti} = 1$ for FMOLS are not reported when they are not significant.

To sum up, we may assert that the first part of the BS hypothesis holds, in the homogeneous version, in both areas. Moreover, the parameter estimated $\hat{\theta}$ tends to be higher in the group of LA countries. At an individual level, we found favourable evidence in 11 out of 16 LA countries – in 7 of them the point estimate does not differ statistically from unity – and in 8 out of 16 OECD members, with 7 of them presenting a point estimate not statistically different from unity.

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 verify whether this relationship is satisfied, we apply panel cointegration tests 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:

$$e_{it} = \gamma_{0i} + \gamma_{pi} dp_{T_{it}} + \varepsilon_{it} \qquad (9)$$

Since in this equation it is assumed that the nominal exchange rate – the dependent variable – adjusts to variations in the price differential of tradable sectors, 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 4 shows the Banerjee and Carrion-i-Silvestre (2007) cointegration statistics. In both areas, the null hypothesis of non-cointegration may be rejected, at least at the 1% significance, with each of the models encountered. Consequently, we may assert that there is a cointegration relationship between the price differential of tradable goods and the nominal exchange rate in each group of countries.

TABLE 4
The first stage of the BS Hypothesis
Cointegration test of Banerjee and Carrion-i-Silvestre (2007)

$$e_{it} = \gamma_0 + \gamma_{pi} dp_{T_{it}} + \varepsilon_{it}$$

(1991-2004)

	LAT		OECD	
Statistics	**Constant**	**Trend**	**Constant**	**Trend**
$\frac{N^{-1/2} Z_{\hat{t}_{NT}}^{\tilde{e}^j} - \Theta_2^{\tilde{e}^j}\sqrt{N}}{\sqrt{\Psi_2^{\tilde{e}^j}}}$	-5.179 [0.000]	-4.654 [0.000]	-4.182 [0.000]	-5.701 [0.000]
Number of factors	3	3	2	2

1. The null hypothesis is no cointegration. p-values are shown within brackets.
2. $Z_{\hat{t}_{NT}}^{\tilde{e}^j}$ corresponds to the un-normalised cointegration test. The values for $(\Theta_2^{\tilde{e}^j}, \Psi_2^{\tilde{e}^j})$ are obtained from Banerjee and Carrion-i-Silvestre (2007), Table 3.
3. Critical values are given for the left tail of the normal distribution
4. To select the number of factors, we follow Bai and Ng (2002).

Since it seemed that a long run relationship exists between the two variables in each group of economies, we decided to estimate the cointegration vector in each model, and then to test the PPP relationship in the tradable sectors.

4.4 The second stage of the BS hypothesis. Cointegration vector

4.4.1 The heterogeneous model

Following the same procedure that we adopted when testing BS-1, we applied both FMOLS and FMOLS(B) techniques in each group of countries to estimate the cointegration vectors of the $e_{it} = \gamma_0 + \gamma_{pi} dp_{T_{it}} + \varepsilon_{it}$. The results are presented in Table 5.

As far as the LA group is concerned, the estimated values of the parameter γ_{pi} with the FMOLS methodology have the correct sign and are statistically significant in each country of the sample. Moreover, they are very close to unity in all countries, with the only exception of Jamaica. Referring to the OECD sample, the coefficients estimated with FMOLS are statistically significant in 9 out of 16 countries of the sample. What is more striking, in most cases the estimated values are either far from unity or have an incorrect sign (Switzerland and the UK).

TABLE 5
Estimation of the cointegration vector
Heterogeneous model: $e_{it} = \gamma_0 + \gamma_{pi} dp_{T_{it}} + \varepsilon_{it}$
(1991-2004)

$\hat{\gamma}_{pi}$	LAT		OECD		$\hat{\gamma}_{pi}$
	FMOLS	**FMOLS(B)**	**FMOLS**	**FMOLS(B)**	
Bolivia	*0.986* (14.005)* (-0.192)	- (4.859)* (-3.140)	*0.386* (0.624) -	- (3.161) -	Australia
Brazil	*0.976* (33.436)* (-0.812)	- (5.370)* (-3.137)	*1.404* (1.224) -	- (3.298) -	Austria
Chile	*0.783* (4.567)* (-1.265)	- (3.376)** (-3.194)	*0.922* (0.726) -	- (3.082) -	Belgium
Colombia	*0.925* (11.056)* (-0.892)	- (4.758)* (-3.258)	*0.403* (2.180)** (-3.234)*	- (3.137) (-2.646)***	Canada
Costa Rica	*0.918* (17.477)* (-1.551)	- (4.561)* (-3.269)	*1.004* (1.884)** (0.007)	- (3.441) (3.373)	Denmark
Dominican Rep.	*1.269* (12.362)* (2.621)*	- (6.575)* (3.894)	*0.909* (0.227) -	- (3.318) -	France
Jamaica	*0.644* (11.870)* (-6.559)*	- (4.522)* (-4.438)*	*2.709* (1.284) -	- (3.181) -	Germany
Mexico	*0.889* (11.362)* (-1.420)	- (4.751)* (-2.905)	*0.874* (4.065)* (-0.587)	- (3.557)** (-3.278)	Ireland
Nicaragua	*1.120* (12.378)* (1.328)	- (4.969)* (3.148)	*0.968* (2.664)* (-0.089)	- (2.569)*** (-3.276)	Italy
Paraguay	*1.164* (9.908)* (1.396)	- (5.014)* (3.054)	*1.265* (1.968)** (0.413)	- (2.912) (3.069)	Korea
Uruguay	*1.079* (12.575)* (0.923)	- (4.863)* (3.078)	*0.784* (0.996) -	- (3.347) -	The Netherlands
Venezuela	*0.837* (38.797)* (-7.530)*	- (5.059)* (-4.735)*	*0.303* (2.084)** (-4.806)*	- (3.398) (-3.453)**	Norway
			1.346 (2.988)** (0.768)	- (2.594)*** (3.078)	Portugal
			1.186 (2.661)** (0.418)	- (2.556)* (3.135)	Spain
			-5.711 (-2.273)** (-2.671)*	- (-3.598) (-2.670)***	Switzerland
			-0.215 (-0.574) -	- (-3.225) -	United Kingdom

1. See the explanations of Table 2.

The results that we obtain by applying bootstrapping inference to account for cross-sectional dependence are reported in columns 2 and 4 under the heading FMOLS(B). To refer first to the LA group, it is worth mentioning that the null hypothesis that γ_{pi} is equal to zero can be rejected in each country of the sample (on the lines of the values of *t* and *t** reported between parentheses in row 2). Furthermore, the null hypothesis $H_0 : \gamma_{pi} = 1$ cannot be rejected in 10 countries of the group (*Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Mexico, Nicaragua, Paraguay, and Uruguay*) in accordance with the t^* bootstrap statistics presented between parentheses in row 3.

As regards the OECD panel, the null $H_0 : \gamma_{pi} = 0$ can be rejected in only four countries, *Ireland, Italy, Portugal and Spain.* In these countries, the null $H_0 : \gamma_{pi} = 1$ can be accepted, which means that they satisfy PPP(T).

4.4.2 The homogeneous model

The results are reported in Table 6. For the Latin American group, the estimated value $\hat{\gamma}_p$ is very close to unity. Moreover, the estimated bootstrap *t** statistics indicate that the null $H_0 : \gamma_p = 0$ can be clearly rejected at 1% significance, and that the null $H_0 : \gamma_p = 1$ should be accepted for the whole panel. Consequently, we may assert that PPP(T) holds in the set of LA countries that did not adopt hard pegs against the US dollar during the period of analysis. This result agrees with the empirical findings of Burstein and Eichenbaum (2005), according to which PPP is a reasonable description of the behaviour of import and export prices in developing countries that experienced large devaluations during the 1990s and 2000s. The reason lies on the high exchange rate pass-through to the tradable prices that the authors detect in these countries.

For the OECD group, bootstrapping inference leads to the conclusion that the null $H_0 : \gamma_p = 0$ can be rejected at 5% significance, and that the null $H_0 : \gamma_p = 1$ may be accepted, despite the fact that the estimated parameter is relatively far from unity (0.717).

To have a more rigorous assessment of the results under cross-sectional dependence we applied the bias adjusted OLS method (BAOLS) suggested by Westerlund (2007) to each group of countries. As can be verified with the estimates presented in columns 3 and 6 of Table 6, the results are maintained in the LA group, but they vary considerably in the OECD panel with respect to those derived with the FMOLS(B) methodology. In the last panel, the null $H_0 : \gamma_p = 0$ can not be rejected on the basis of the BAOLS *t* statistics. Consequently, PPP(T) can not be accepted in the OECD area.

The latter results are in agreement with previous findings in this respect. Thus, SØndergaard (2001) detected disequilibria in the relative prices of the tradable goods of a group of OECD countries, and attributed them to monopolistic competition between firms. Engel (2002) also found that the variations in the RER in a set of OECD economies were almost exclusively caused by deviations from PPP in the tradable sectors, due not only to transportation costs, but also to the pricing-to-the-market behaviour of firms. García-Solanes, Sancho and Torrejón (2007), taking Germany as a benchmark, found very similar results in a group of six EU-15 countries.

TABLE 6
Estimation of the cointegration vector
Homogeneous model: $e_{it} = \gamma_{0i} + \gamma_p dp_{T_{it}} + \varepsilon_{it}$
(1991-2004)

	LAT			OECD			
	FMOLS	**FMOLS(B)**	**BAOLS**	**FMOLS**	**FMOLS(B)**	**BAOLS**	
$\hat{\gamma}_p$	*0.929*		*0.952*	*0.717*	-	*0.730*	$\hat{\gamma}_p$
	(6.975)*	(4.475)*	(7.293)*	(6.216)*	(5.330)**	(0.463)	
	(-0.530)	(-0.649)	(-0.368)	(-2.450)**	(-4.371)	-	

1. See the explanations provided in Table 4.
2. BAOLS is the Bias Adjusted OLS suggested by Westerlund (2007). The *t statistics* of BAOLS follows a normal distribution, and its critical values are: +/-2.576(*), +/-1.960(**) y +/-1.645(***). They correspond to the 1%, 5% and 10% level of significance, respectively.

Consequently, the main explanation of the variations of RER(T) that preclude the fulfilment of PPP(T)in the group of OECD is **market segmentation**. The fact that most of the estimated values of parameter γ_p – and also the point estimate of the whole area – are relatively far from unity in the OECD countries is consistent with the existence of two non-arbitrage bands within which the RER(T) may evolve randomly.

TABLE 7
The two stages of the Balassa and Samuelson hypothesis
(1991-2004)

	LAT					OECD					
	First part			Second part		First part			Second part		
	θ_{Ni}	$\hat{\theta}_{Ti}$	**Fulfils**	$\hat{\gamma}_{\rho i}$	**Fulfils**	θ_{Ni}	$\hat{\theta}_{Ti}$	**Fulfils**	$\hat{\gamma}_{\rho i}$	**Fulfils**	
Argentina	1	3.045	*Yes*	---	---	1	0.948	*Yes*	n.s.	*No*	Australia
Bolivia	1	1.372	*Yes*	0.986	*Yes*	1	n.s.	*No*	n.s.	*No*	Austria
Brazil	1	1.762	*Yes*	0.976	*Yes*	1	n.s.	*No*	n.s.	*No*	Belgium
Chile	1	n.s.	*No*	0.783	*Yes*	1	1.910	*Yes*	n.s.	*No*	Canada
Colombia	1	1.307	*Yes*	0.925	*Yes*	1	n.s.	*No*	n.s.	*No*	Denmark
Costa Rica	1	n.s.	*No*	0.918	*Yes*	1	n.s.	*No*	n.s.	*No*	France
Dominic Rep.	1	n.s.	*No*	1.269	*Yes*	1	1.942	*Yes*	n.s.	*No*	Germany
Ecuador	1	1.201	*Yes*	---	---	1	n.s.	*No*	0.874	*Yes*	Ireland
El Salvador	1	1.989	*Yes*	---	---	1	1.126	*Yes*	0.968	*Yes*	Italy
Jamaica	1	1.253	*Yes*	0.644	*No*	1	n.s.	*No*	n.s.	*No*	Korea
Mexico	1	1.381	*Yes*	0.889	*Yes*	1	n.s.	*Yes*	n.s.	*No*	The Netherlands
Nicaragua	1	1.438	*Yes*	1.120	*Yes*	1	3.771	*No*	n.s.	*No*	Norway
Panama	1	0.884	*Yes*	---	---	1	0.876	*Yes*	1.346	*Yes*	Portugal
Paraguay	1	n.s.	*No*	1.164	*Yes*	1	1.067	*Yes*	1.186	*Yes*	Spain
Uruguay	1	n.s.	*No*	1.079	*Yes*	1	0.770	*Yes*	n.s.	*No*	Switzerland
Venezuela	1	1.648	*Yes*	0.837	*No*	1	n.s.	*No*	n.s.	*No*	United Kingdom
Panel	**1**	**1.402**	***Yes***	**0.929**	***Yes***	**1**	**0.866**	***Yes***	**0.717**	***No***	**Panel**

As a synthesis of the empirical part of this paper, Table 7 summarises the results of our tests applied to the first and second stage of the BS hypothesis. Looking at the simultaneous fulfilment of the two BS stages, we find that in the Latin American group the hypothesis holds in the area as a whole and in five individual countries, 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 holds only in Italy, Portugal and Spain but not in the whole area due to PPP deviations 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 3.3.

V 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 use pooled observations and apply recent panel techniques to overcome the problems of insufficient time series data in many countries and cross-sectional dependence in the data of our samples –especially in the case of the OECD countries–.

We find that while 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, is satisfied in each group of countries, the second stage, which relates the price differential with the real exchange rate, holds in the Latin American area, but not in the group of the OECD countries as a whole, nor in most of their individual members included in our sample. The failure is reflected in departures from PPP in the tradable sectors, and is probably due to transportation costs and non-competitive practices that still prevail in the countries of this area. Putting together the results for all the individual countries of our sample, it follows that the entire BS hypothesis clearly holds in five Latin American countries and perhaps in three OECD economies.

Fulfilment of the BS hypothesis in the whole Latin American sample 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[12]. Difficulties are particularly severe in

[12] Apart from the – equilibrium – long term adjustments imposed by BS effects and other permanent real factors, the RER experiences short-run fluctuations as a result of nominal rigidities. According to Calderón and Schmidt-

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 Yves-Yeyati (2003)[13]. 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.

Hebbel (2003), deviations from equilibrium RERs are persistent in LA countries, with a median half-life of 2.6 years.

[13] As outlined above 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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