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, taking the USA 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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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 compared with Germany since the early nineties (Halpern and Wyplosz (2001), Kovács (2002), Égert (2002a,b), Mihaljek and Klau (2004), Égert et al. (2002)).

Some recent papers have investigated the fulfilment of the BS effect in groups of developing countries tacking the USA as the reference external country. Thus, Drine and Rault (2003), tested the BS hypothesis using annual data of 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 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 the context of this area.

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

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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.

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, 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) 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 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 Canzoneri, Cumby and Diba (1999).

This paper presents two novelties that, in our opinion, 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 is 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 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 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 some 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 verdict must be interpreted in very relative terms 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, things look very different in the tests of the second part of the hypothesis (BS-2). Here 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 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.

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 et al. (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,** 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 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:

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

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.

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

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:

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

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**:

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

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.

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[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]:

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

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:

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

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

[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. Søndergaard (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 that takes 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.

[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)).

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[7] (Krugman (1987)), and b) *arbitrage frictions*, created notably by 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 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[8]. 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. 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 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

[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.

take the USA economy as the benchmark foreign country, since all the countries mentioned have substantial economic exchanges 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. 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.

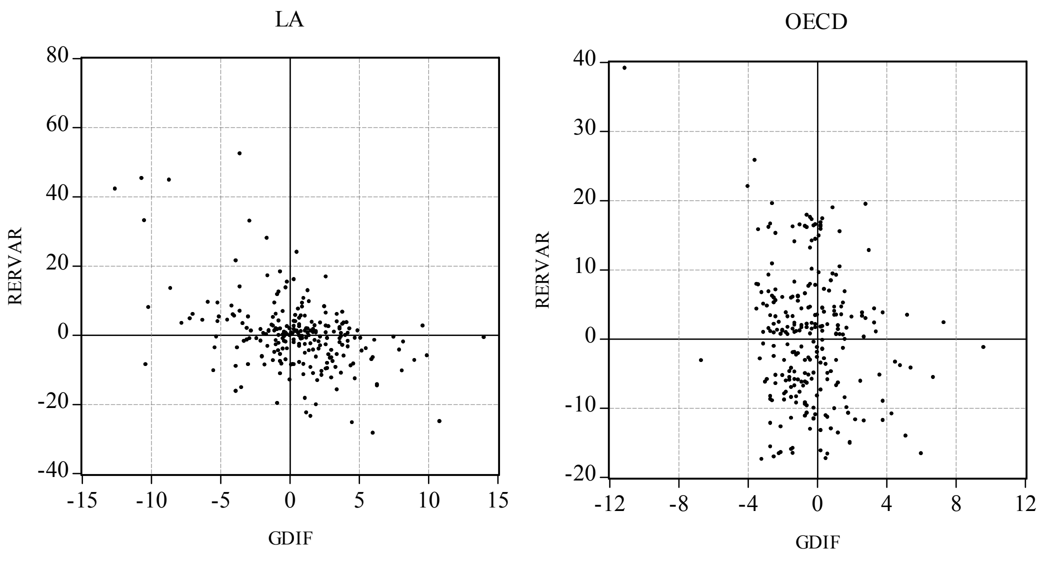
We define the relative price of non-tradable goods with respect to tradable ones as the ratio of the two corresponding sector GDP deflators. To obtain deflator indices we first measured the aggregate production, that is, the value added, in each sector in both nominal and real terms, using current prices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

GRAPH 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.

IV. 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 investigate 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 to infinity, as discussed in Wagner and Hlouskova (2004) and Wagner (2005). 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.

The presence of *Cross-Sectional Dependence* in the data has led some analysts to develop new econometric tests to deal with this problem in panel data samples. As far as unit root analysis are concerned, Jhönson (2004) and Pesaran (2007) built new unit root and stationary tests, respectively, to overcome the drawbacks created by this particular phenomenon. Furthermore, Jhönson (2004) proved that the statistics of his test works well with short time series ($T<30$), and Pesaran (2007) modified the critical values of his test in order to reach correct inferences in small sample and time dimensions ($10<N<200$ and $10<T<200$). As regards cointegration analysis, Banerjee and Carrion-i-Silvestre (2006) suggested a transformed Pedroni *Group-t* statistics, which under the null of not cointegration, fits an asymptotically normal distribution. Wagner and Hlouskova (2004), Wagner (2005), Westerlund and Basher (2006) and García-Solanes, Sancho and Torrejón (2008) applied new cointegration techniques to test the BS effect and PPP in the presence of cross-sectional dependence in several groups of European countries.

Although the above mentioned contributions correctly solve the problems raised by cross-sectional dependence, potentially wrong conclusions are still likely if the *dimension of the time series is relatively small*. The reason is that with short time series, the empirical distribution of the statistics could be biased towards either the left or the right of the theoretical one, leading then to wrong diagnosis about the null hypothesis. In order to solve both problems, cross-sectional dependence and short time series, some authors suggest applying bootstrap techniques in both unit root and cointegration analysis. These techniques generate empirical distributions after multiple re-sample of the data, from which the critical values are derived. Li and Maddala (1997) applied three bootstrap procedures to analyse cointegration relationships between individual time series, and Li and Maddala (1996) and Chang (2004) designed bootstrapping methodologies for panel unit root tests in samples affected by cross-sectional dependence. Westerlund and Edgerton (2007) applied bootstrap analysis to test the null of cointegration in panels with cross-sectional dependence. Finally, Wagner and Hlouskova (2004) suggested three bootstrap methods to deal with both cross-sectional dependence and small sample dimensions: a) Parametric bootstrap, b) Non-parametric bootstrap, and c) Residual-based-block-bootstrap. All of them take into consideration the small dimension of the panel and preserve from certain degree of cross-sectional dependence in the panel.

To begin with our econometric analysis of the two BS stages, 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 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 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.

Furthermore, since the time series of our panel are not sufficiently large to guarantee the absence of biased results, in the cointegration analysis we performed bootstrap inference 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. 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.

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 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[9]. 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:

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

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

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 (spurious regression); 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 non-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.

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[9] 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 **(-1.916)**	3.038 (0.733)	0.584 (-24.777)	-1.109 (-12.517)	4.688 (2.644)	-5.930* (-44.488)	-2.616* (-12.880)
(8')	-2.118 **(-2.581)**	2.428 (-0.613)	-0.243 (-25.007)	-4.186* (-13.305)	3.757 (1.469)	-2.168* (-39.527)	-4.709* (-13.737)
				OECD			
(8)	-2.206 (-1.634)	3.938 (0.593)	-1.183 (-26.680)	-4.561* (-13.047)	4.521 (2.382)	-8.757* (-48.242)	-3.591* (-14.098)
(8')	-1.164 **(-1.988)**	1.471 (-0.694)	-2.903* (-26.879)	-7.296* (-13.849)	2.658 (1.549)	-3.952* (-40.016)	-6.941* (-13.595)

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.

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, 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')).

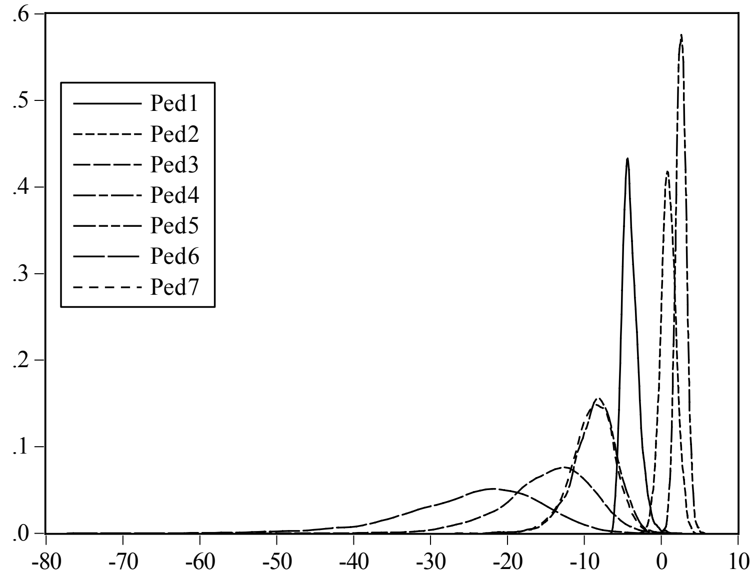
GRAPH 2

First stage of the BS hypothesis

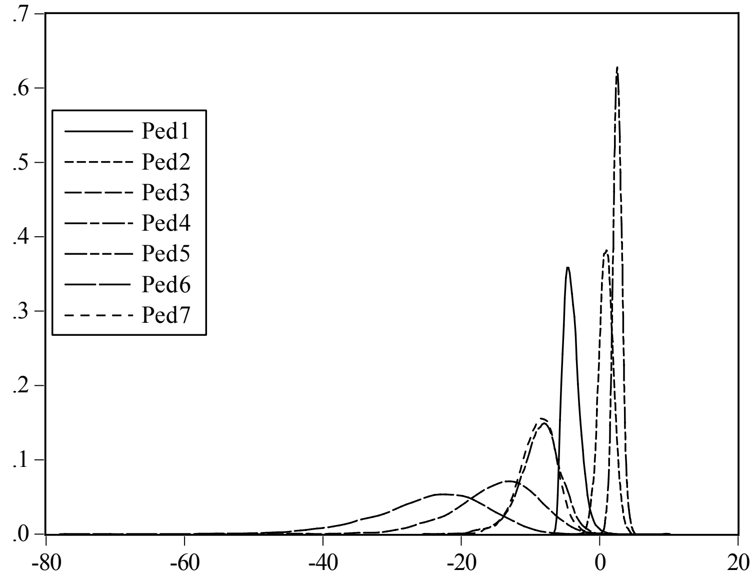
Empirical bootstrap distributions for the cointegration test of Pedroni (1999, 2004)

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

LA



OECD



1. The results are obtained with 5000 re-samples.
2. The estimations for the two equations include fixed effects.

Graph 2 shows the empirical bootstrap distribution for each of the Pedroni (1999, 2004) tests. As can be seen, most of them are biased to the left with respect to the theoretical normal distribution. For this reason, the possibility of rejecting the null hypothesis decreases with the application of the bootstrapping methodology. In fact, bootstrapping restricts this outcome to the results from $T^2 N^{3/2}\left(Z_{\hat{v}_{NT}}\right)$(Panel-$v$ statistic), and in the case of the OECD group the possibility of rejection shrinks further to the restricted version of the testable equation.

However, since there is some empirical basis to reject the null hypothesis of no cointegration in the restricted version of the model in each panel –which in turn matches well with the impression drawn from Graph 1–, the possibility of a long run relationship between the price differential and average labour productivities, as stipulated by the Baumol and Bowen (1966) effect, exists in each group of countries. In order to further explore this possibility on the basis of the bootstrap results reported in column 1 of Table 1, we estimated the vector error correction (VEC) for the restricted equation in each panel. We assumed that if the equilibrium error triggers a statistically significant dynamic adjustment of the explained variable towards its long-run equilibrium value, this is a good symptom of a long-run equilibrium relationship between the explanatory and explained variables. Results obtained with OLS regressions are presented in Table 2, where $y_{it} = \left(dp + da_N\right)_{it}$ and d stands for the differencial operator.

TABLE 2
First stage of the BS hypothesis
VEC estimation for the restricted equation

$$\mathrm{d}(y_{it}) = \delta_{0,i} + \lambda_T (y_{it-1} - \theta_{0,i} - \theta_T da_{T,it-1}) + \delta_1 \mathrm{d}(y_{it-1}) + \delta_2 \mathrm{d}(da_{T,it-1}) + \upsilon_{it}$$

	LA	OCDE
λ_T	-0.183 (-2.759)*	-0.095 (-2.127)**

1. Estimations include fixed effects in both VEC and cointegration equations.
2. Asterisks (*) and (**) indicate rejection of the null hypothesis at 1%, and 5% confidence levels, respectively.

As can be verified, the adjustment coefficient has the correct sign and is statistically significant in both groups of countries (at 1% and 5% significance in LA and OECD, respectively). 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 3 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.

Let us first describe the main properties of the BAOLS methodology and the results that we obtain by applying it to the two panels of our study. This 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 $\left(\hat{\beta}\right)$ and, subsequently, he calculated the bias of this estimation $\left(\hat{b}_{NT}\right)$ taking into account the optimal number of common factors as suggested by Bai and Kao (2002, 2005). Then he defined the BAOLS estimator as $\beta^{+} = \hat{\beta} - \hat{b}_{NT}$, which follows an asymptotically normal distribution.

The results obtained with **BAOLS** are reported in columns 1 and 3. 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 higher than unity in the LA countries (1.462) and lower than unity in the OECD group (0.855), 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 $\theta_T = 0$ is rejected at the 10% significance level in LA but it is not rejected in OECD. The 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 each panel.

TABLE 3
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)**	
$\hat{\theta}_T$	*1.462* (1.851)*** (0.585)	- (0.802)* (0.332)**	*0.855* (0.480) (-0.081)	- (0.431)** (-0.224)	$\hat{\theta}_T$

1. The first row shows the estimated values of θ_T by the BMOLS 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.

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 BMOLS estimator. We use the *Moving Block Bootstrap* algorithm proposed by Li and Maddala (1997) and Li and Xiao (2003), following the Westerlund (2007) methodology, because it preserves against the degree of cross-section dependence in the data, permitting then to infer correctly the significance of the parameters.

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 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 are shown in columns 2 and 4 under the heading **BMOLS(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 different from zero is rejected in both groups of countries: at 1% level in LA and at 5% in OECD. Furthermore, the hypothesis $\theta_T = 1$ is rejected in LA, confirming the value considerably higher than 1 reported in column 1, but not in the OECD.

To sum up, we find some evidence that the first part of the BS hypothesis can not be rejected in each area, by applying recent econometric methods that account for, and solve, the problems created by both cross-sectional dependence and relatively small size of the samples. 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 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 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 results for both groups of countries. For each panel, the numbers in the first row correspond to the estimation with fixed effects, and those of the second row include both fixed effects and one deterministic trend. Application of the simple Pedroni (1999, 2004) test points out that the null hypothesis of non cointegration is rejected in both panels with statistics $N^{-1/2}\tilde{Z}_{t_{NT}}$ and $N^{-1/2}\tilde{Z}^{*}_{t_{NT}}$. In the OECD group, rejection comes from each version of the estimated equation. Furthermore, in this panel, rejection also derives from $Z^{*}_{t_{NT}}$ using the fixed effects version. The Non-parametric bootstrapping methodology of Wagner and Hlouskova (2004) also provides some evidence of rejection, but it is restricted to the two estimating versions with the first statistic, $T^2 N^{3/2}\left(Z_{\hat{v}_{NT}}\right)$

TABLE 4
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			
No Trend	-0.425 **(-1.979)**	1.767 (-0.529)	0.264 (-21.348)	-0.864 (-11.659)	2.178 (1.307)	-2.732* (-35.017)	-2.205* (-11.827)
Trend	0.843 **(0.736)**	0.224 (-3.271)	0.202 (-17.442)	0.199 (-11.599)	1.365 (-0.980)	0.533 (-28.733)	0.018 (-13.342)
				OECD			
No Trend	-0.499 **(-1.201)**	2.304 (-0.895)	-0.727 (-28.064)	-4.251* (-14.517)	2.988 (1.404)	-2.039* (-42.289)	-4.106* (-14.124)
Trend	1.738* **(1.669)**	-0.139 (-3.968)	-0.891 (-19.109)	-1.512 (-13.304)	0.617 (-1.204)	-2.995* (-31.170)	-4.851* (-14.760)

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

For the same reasons explained in the cointegration test of the first part of BS, we also estimate here the VEC corresponding to probably long-run equilibrium relationship suggested by the results of column 1 of table 4. The estimations obtained with OLS regressions are presented in table 5, where $y_{it} = (dp + da_N)_{it}$ and **d** stands for the differential operator.

TABLE 5
Second stage of the BS hypothesis
VEC estimation for the restricted equation

$$\mathrm{d}(e_{it}) = \delta_{0,i} + \lambda_p (e_{it-1} - \gamma_{0,i} - \gamma_p dp_{T,it-1}) + \delta_1 \mathrm{d}(e_{it-1}) + \delta_2 \mathrm{d}(dp_{T,it-1}) + \upsilon_{it}$$

	LA	OCDE
λ_p	-0.392 (-2.132)*	-0.511 (-8.433)**

1. Estimations include fixed effects in both VEC and cointegration equations.
2. Asterisks (*) and (**) indicate rejection of the null hypothesis at 1%, and 5% confidence levels, respectively.

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. 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.

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 BMOLS and bootstrapping technique to it, BMOLS(B), in each group of countries to estimate the cointegration vectors of the equation $e_{it} = \gamma_{0,i} + \gamma_p dp_{T,it} + \varepsilon_{it}$.

TABLE 6
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)**	
$\hat{\gamma}_p$	*0.952* (7.293)* (-0.368)	- (0.716)* (-0.370)	*0.730* (0.463) -	- (0.706) -	$\hat{\gamma}_p$

1. See the explanations provided in Table 2.
2. The BMOLS(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 $H_0 : \gamma_p = 0$ or 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.

The results are reported in Table 6, using the same presentation than 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 at 1% significance by both BAOLS and BAOLS(B), and that the null $H_0 : \gamma_p = 1$ can not be rejected, for the whole panel and with each econometric methodology. 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 several recent empirical studies, which find that the exchange rate pass-through (ERPT) ratio to the prices of domestic tradables, and particularly to import prices, reach 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) and Burstein and Eichenbaum (2005) confirm these general results, remarking that ERPT is complete (100%) when import prices are measured at the docks[10]. It is worth noting that fulfilment of PPP(T), as a result of almost complete pass-through, is due to certain pricing behaviour of trading firms, and could be compatible with the existence of transportation costs.

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 clear long-run relationship between the nominal exchange rate and the price differential in the tradable sectors of the area. The direct implication of this finding is that the confidence interval of the parameter γ_p is very large and that, consequently, there are no grounds for testing the null hypothesis $H_0 : \gamma_p = 1$. Consequently, 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 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 (2008), taking Germany as a benchmark, found very similar results in a group of six EU-15 countries.

Taking into account those results and the theoretical discussion presented in section 2.2 of this paper, we derive that in the OECD area variations in the real exchange rate of tradables and, consequently, the failure of PPP(T), obeys to **market segmentation** likely created by both imperfect competition and arbitrage frictions.

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[10] The explanation of these empirical results lies on the facts that: a) emerging market economies are usually price takers in international goods markets, and that b) foreign firms set prices in its own currency, or preponderantly in US dollar (producer currency pricing). Consequently, foreign firms do not modify their prices expressed in their own currency after nominal exchange rate variations, which this leads to complete ERPT to importing prices in emerging market economies. This pricing behaviour allows the foreign firms to maintain their mark-up constant.

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).

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 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 some 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. However, contrary to previous studies that reach unanimous and definitively conclusions on this respect with the application of conventional cointegration analysis, our findings must be interpreted just as non-excluding the positive result. 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 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 graph 1, and corroborated by some of 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[11]. 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 Yves-Yeyati (2003)[12]. 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.

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

[12] 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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