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TESTING THE BALASSA-SAMUELSON EFFECT:

IMPLICATIONS FOR GROWTH AND PPP

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Abstract

The derivation of the Balassa-Samuelson effect allows for different empirical specifications that may have important economic implications. Problems related to spurious regression could arise from the mixed order of integration of the series used and from the lack of a long run stable relationship among the variables of the model. This paper addresses these problems by using the bounds testing approach developed by Pesaran, Shin and Smith (1999). Our empirical results do not show supportive evidence for the Balassa-Samuelson effect in the long run. This seems to suggest that PPP holds. However, one of the implications of PPP is that the real exchange rate does not have any real impact on the economy. Further empirical analysis rejects this implication. In fact, the real exchange rate seems to have a long run impact on relative growth rates.

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1. Introduction

For most countries there has been a tendency for increases in the prices of non-traded goods to exceed increases in the prices of traded goods (e.g. Kravis and Lipsey, 1988). The Balassa-Samuelson model explains this phenomenon through the differential productivity growth between the traded and non-traded goods' sectors[1]. It is argued that the traded goods sector has a higher productivity growth than the non-traded goods sector. Therefore the relative slower rate of growth in the non-traded goods sector results in higher relative non-traded goods' prices[2]. One prediction of this model is that the tradable-nontradable price difference is lower for rich countries than for poor countries (see Heston *et al*, 1994). Another consequence is that if traded goods' productivity relative to non-traded goods' productivity is growing faster at home than abroad, then the home country should experience an appreciation of the real exchange rate[3].

The Balassa-Samuelson model rests on two assumptions: (1) labor markets are competitive within each country, thus labor mobility leads to wage equalization between the traded and non-traded goods' sectors; and (2) PPP holds only for tradable goods. Recent empirical work has focused on examining these assumptions[4]. Strauss (1998) rejects these

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[1] Actually the main motivation behind the model was to explain the persistent deviations from PPP (see Balassa, 1964, and Samuelson 1964, 1994).
[2] The other competing theories to explain the phenomenon are the relative-factor-endowments model associated with Bhagwati (1984) and the Linder-type hypothesis that stresses the role of demand factors (see, Bergstrand, 1991).
[3] See Rogoff (1996).
[4] Meanwhile, the theoretical work has derived the key propositions of the Balassa-Samuelson model in dynamic two-sector growth models, two-country general equilibrium models, and open economy models with imperfect competition (e.g. Asea and Mendoza, 1994; and Balvers and Bergstrand, 1997).

assumptions for the major industrial countries, and argues that the results are consistent with the presence of industry and/or sectoral-specific human capital. Canzoneri *et al* (1999) test these assumptions using a panel of OECD countries. The results indicate that relative prices reflect relative labor productivities in the long run. However, the evidence on PPP in traded goods is less favorable[5].

This paper employs a different method in that the Balassa-Samuelson effect is tested using a time-series approach. The main objective is to explain some problems that might appear from the theoretical model and its econometric implementation. Specifically, two issues are of interest. First, the theoretical derivation of the model allows us to use two different types of reduced form specification. This raises the problem of which reduced form is more appropriate. In order to choose the correct reduced form, the orders of integration of the time series are of extreme importance. The second issue is the examination of the long run statistical properties of the time series, since it is important to show whether or not they have a long run stable relationship. Therefore, testing for cointegration between the series is essential to address this issue. As we shall see, this will have consequences for the validity of the Balassa-Samuelson effect in the long run, bringing forward important implications for the PPP hypothesis and economic growth.

The paper is structured as follows. The next section presents the model. Section 3 describes the econometric methodology. The estimation results appear in section 4. Finally, section 5 offers some concluding remarks.

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[5] De Gregorio *et al* (1994) examine other factors beyond the productivity differential to explain the relative price between tradable and non-tradable goods, such as demand shifts toward non-tradable goods, and real wage pressures. The relevance of these factors is analyzed empirically for France, Germany, Italy, Spain, and the UK.

2. The Model

There are two countries (the foreign country is denoted with an asterisk) that use labor (L) to produce, under constant returns to scale technology, two types of goods, a traded good (T) and non-traded good (N)[6]:

$$Y_T = f(L_T) \qquad Y_N = g(L_N)$$

$$Y_T * = F(L_T *) \qquad Y_N * = G(L_N *)$$

The labor market is competitive and labor is perfectly mobile within each country but not between countries. As a consequence the nominal wage is the same in both sectors for each country:

$$P_T \ f'(L_T) = w = P_N \ g'(L_N) \tag{1}$$

$$P_T * F'(L_T *) = w* = P_N * G'(L_N *) \tag{2}$$

where the prime after the function denotes the marginal productivity of labor. The second assumption of the Balassa-Samuelson model is that PPP holds only for tradable goods:

$$P_T = e P_T * \tag{3}$$

where e denotes the nominal exchange rate.

The price levels are defined as weighted geometric averages of prices in both sectors:

$$P = P_T^{1-i} \ P_N^i \tag{4}$$

$$P* = P*_T^{1-j} \ P*_N^j \tag{5}$$

Without loss of generality, to simplify matters we can make the price of tradable goods equal to one: $P_T = P_T * = 1$. By equation (3), this implies that the nominal exchange rate is also equal to one: $e = 1$. Equations (4) and (5) can thus be rewritten as:

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[6] See Obstfeld and Rogoff (1996) for a model with capital. It should be stressed that a model with capital must assume perfect international capital mobility in order to eliminate the role of demand side factors in the determination of relative prices.

$$P = P_N^i \qquad (4')$$

$$P^* = P^{*j}_N \qquad (5')$$

Similarly, from equations (1) and (2):

$$P_N = f'(L_T)/g'(L_N) \qquad (1')$$

$$P_N^* = F'(L_T^*)/G'(L_N^*) \qquad (2')$$

The real exchange rate is defined as:

$$\theta = P/eP^* = P/P^* \qquad (6)$$

Substituting equations (1') and (2') into equations (4') and (5') and them into equation (6) yields:

$$\theta = P/P^* = \frac{[f'(L_T)/g'(L_N)]^i}{[F'(L_T^*)/G'(L_N^*)]^j} \qquad (7)$$

Equation (7) is the Balassa-Samuelson effect. It asserts that if traded goods productivity relative to non-traded goods productivity is growing faster at home than abroad, then the home country should experience an appreciation of the real exchange rate.

Due to the assumption of constant returns to scale technology, the marginal productivity of labor is proportional to the average product of labor[7]. In this case, the right hand side of equation (7) can be rewritten in terms of the average productivity of labor. The impact on the real exchange rate is analogous. If traded goods' average productivity relative to non-traded goods' average productivity grows faster at home than abroad, the home country will have its real exchange rate appreciated.

Notwithstanding the theoretical flexibility of the model, the use of marginal or average labor productivity has important consequences for the empirical implementation of the Balassa-Samuelson effect. First, considering the use of average labor productivity, one can

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[7] See Canzoneri *et al* (1999) for a detailed discussion on this issue.

argue that the relative average labor productivity between countries is proportional to relative real output per capita levels. That is, the greater the ratio between the relative average productivity, the greater the output per capita ratio between countries. So, the reduced form of the model to be estimated is:

$$\theta = P / P^{*} = \frac{[Y]^{i}}{[Y^{*}]^{j}} \qquad (8)$$

Second, considering the marginal productivity of labor, one can argue that the relative marginal productivity of labor between countries is proportional to relative rate of variation of the real per capita output. So, the reduced form becomes:

$$\theta = P / P^{*} = \frac{[\Delta Y]^{i}}{[\Delta Y^{*}]^{j}} \qquad (9)$$

Instead of the variation of the real per capita output one can use the relative rate of variation of total factor productivity, as measured by the Solow residual[8].

It is easy to see that these specifications have very different econometric implications whenever time series are taken into account. Taking the same independent variable in levels or in first differences can result in different orders of integration. Moreover, in a time series context the long run relationship between the variables of the reduced forms should be addressed. Hence, problems related to spurious regression could arise from different orders of integration of the series and from the lack of a long run stable relationship among the variables of the model. This paper addresses these problems by using the bounds testing approach developed by Pesaran, Shin and Smith (1999) which allows for testing the existence of long run relationships between variables regardless of whether they are integrated of order

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[8] De Gregorio and Wolf (1994) argue that the impact of Solow residual captures the Balassa-Samuelson effect, while the variation of real GDP would capture the demand effects. However, there is a problem with this procedure since, by construction, the Solow residual is derived from the measure of output.

one or zero. In addition, the results of the long run properties of the model lead to the analysis of some causality issues. The causality results obtained can have important implications for the interpretation of the model.

3. Econometric methodology.

According to the above discussion, the two testable reduced form specifications of the model can be expressed as follows:

$$PR(ij)_t = \alpha_1 + \beta_1 YR(ij)_t + u_t \tag{10}$$

$$PR(ij)_t = \alpha_2 + \beta_2 \Delta YR(ij)_t + v_t \tag{11}$$

Equations (10) and (11) express the long run relationship between the price ratio (*PR*) of two countries *i* and *j* and their per capita output ratio (*YR*) expressed in logarithms. Both the price ratio and the output ratio can be either I(1) or I(0) variables[9]. In the case in which both variables are I(1), we could use the well-known cointegration tests for the existence of a long-run cointegration vector. However, two facts make this approach inappropriate for the purposes of our test. First, the use of price and output ratios may lead these variables to become I(0) if the log of prices (and output per capita) of countries *i* and *j* are co-integrated with a vector (0, 1). Second, in equation (11) by first-differentiating the output ratio we will surely end up with an I(0) variable in the right-hand side. For these reasons tests of the Balassa-Samuelson effect based on traditional cointegration techniques would be flawed.

In a recent paper, Pesaran *et al* (1999) develop a technique to test for the existence of a long-run relationship between two variables irrespective of whether they are I(1) or I(0). For this reason, Pesaran *et al*'s (1999) methodology becomes most useful in this model where

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[9] Based on previous unit-root tests using the Dickey-Fuller and Phillips-Perron tests we can rule out the possibility that any of the series is I(2).

variables with different orders of integration are involved. Their approach is based on the estimation of a dynamic error correction representation for the variables involved and testing whether or not the lagged levels of the variables are significant. In other words, Pesaran *et al*'s (1999) test consists of the estimation of the following conditional error correction model (ECM):

$$\Delta y_t = \alpha_0 + \alpha_1 t + \beta_1 y_{t-1} + \beta_2 x_{t-1} + \sum_{i=1}^{p-1} \varphi_i \Delta y_{t-i} + \sum_{i=1}^{p-1} \theta_i \Delta x_{t-i} + \omega \Delta x_t + u_t \qquad (12)$$

In order to test for the existence of a long run relationship Pesaran *et al* (1999) consider two alternatives. First, an F-statistic test of joint significance of the lagged levels of the variables involved[10]. Second, following Banerjee *et al* (1998), a t-ratio test for the significance of the lagged level of the dependent variable (y_{t-1}). Pesaran *et al* provide two sets of critical values assuming that both regressors are I(1) and that both are I(0). These two sets provide a band covering all possible combinations of the regressors into I(0), I(1) or mutually cointegrated[11]. Also, if the F-statistic for the joint null of zero coefficients on y_{t-1} and x_{t-1} is insignificant, then we cannot reject the hypothesis that the variable x_t is not a *long run forcing variable*. By interchanging y_t and x_t as dependent and independent variables in regression (12) we can assess whether y_t is or is not a forcing variable. Note that this is a necessary but not sufficient condition for *Granger-non-causality*.

In the next section we apply the test described above to the ECM representation of equations (10) and (11) for a set of pairs of developed countries and analyze causality aspects of these reduced forms.

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[10] When the model contains a deterministic trend, the F-test also includes the null of $\alpha_1 = 0$.

[11] We refer to Pesaran *et al* (1999) for a detailed description of the testing procedure. Note that the critical values provided contain an upper and lower bound outside which inference is conclusive. However, if the F- or t-statistics fall within these bounds, we cannot reach any conclusion unless the cointegration rank of the forcing variable x_t is known a priori.

4. Results

The empirical application was carried out for the different pairs of combinations of relative prices and output for Germany, Japan, United Kingdom and the US. Data is quarterly ranging from 1960:1 to 1996:4. This provides us with 148 observations and 37 years of sample period, giving enough degrees of freedom to apply the tests considered. Data is obtained from the OECD Statistical Compendium 1997:2. The price level is measured using the GDP deflator and output is the index of the per capita GDP at constant prices of domestic currency. All the variables considered have been transformed into logarithms.

The test for a long run relationship consists of two steps. First, the selection of the optimal number of lags of the first difference of the variables to be included in the ECM. Following Pesaran *et al* (1999), this was done using multiple criteria: the analysis of both the Schwarz Bayesian Criteria (SBC) and the Akaike Information Criteria (AIC) as well as the analysis of the LM tests for residual serial correlation of orders 1 and 4. The optimal number is chosen to maximize the SBC and AIC starting with a maximum number of lags of 8 provided that the model does not show signs of serial correlation[12]. Second, we obtain the F- and t-statistics as described above. It should be noted that in none of the cases was a deterministic trend found to be significant and, hence, we decided to drop it from the equations estimated[13].

The results of the F- and t-statistics are reported in Tables 1 and 2 for each of the specifications of the model and for each pair of countries. The results in Table 1 clearly show that the level of price ratio and the level of output ratio do not have a long run relationship in any of the cases considered looking at both the F- and the t-statistics. Table 2 provides the results of the tests for the regression of the level of price ratio on the first difference of the

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[12] The optimal number of lags using a general-to-specific criterion gave very similar results.
[13] All the results not reported in the tables are available from the authors upon request.

output ratio. The results again clearly show that in none of the cases can we accept the hypothesis of the existence of a long run stable relationship.

These results reject any of the specifications of the Balassa-Samuelson effect discussed above. This is in accordance with the great body of empirical literature that finds that the PPP hypothesis holds in the long run (see, inter alia, Lothian and Taylor, 1996). However, our exercise does not necessarily stop at this point. If one accepts the implications of PPP, the real exchange rate is constant and independent of all real variables in the economy. Therefore, one would expect no causation from the real exchange rate to any real variable. Evidence presented in Tables 3 and 4 suggest a different picture. Tables 3 and 4 present the results of the tests inverting the dependent and independent variables.

In Table 3, in which the price ratio is the explanatory variable and the output ratio is the dependent variable, we can find mixed results. For the case of the UK and the US there is clearly a long run relationship. For the cases of Germany and Japan, and Japan and the US, this relationship seems to be weak especially in the former. In the rest of the cases we cannot find any evidence of a long run relationship.

Table 4 presents the most striking results. We can accept in all cases, and with a high level of confidence, the existence of a long run relationship between the first difference of the output ratio and the level of the price ratio. These results suggest that the long run forcing variable in the estimated equations is the level of relative prices with the change in relative per capita output levels adapting to it.

These results indicate that the real exchange rate has important effects on relative growth rates. They seem to be supportive of an increasing strand of the literature on growth that stresses the empirical relevance of real exchange rate on economic growth. Sala-i-Martin (1997), for example, found that real exchange rate distortions were significant and negatively related to growth in a cross section of countries (see also, Barro and Sala-i-Martin, 1995).

This link is more specifically addressed in Andrés *et al* (1996) who use the Balassa-Samuelson effect to argue that another channel by which inflation can affect growth is through the real exchange rate[14].

5. Concluding Remarks

This paper has investigated the empirical relevance of the Balassa-Samuelson effect using a time series approach. The theoretical derivation of the model leads to two different specifications of the reduced form. These specifications have different time series behaviors. We addressed this problem together with the analysis of the existence of a long run relationship between the variables. These problems were tackled by using the bounds testing approach developed by Pesaran, Shin and Smith (1999) which allows for testing the existence of long run relationships between variables regardless of whether they are I(1) or I(0).

Our results do not show supportive evidence for the Balassa-Samuelson effect in the long run. This seems to suggest that long-run PPP holds. However, one of the implications of PPP is that the real exchange rate does not have any real impact in the economy. Further empirical analysis rejects this last implication. In other words, although the null of the Balassa-Samuelson effect is rejected, this does not mean that we can accept the possible alternative of PPP. In fact, the real exchange rate seems to have a long run impact on relative growth rates.

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[14] The real exchange rate can affect long run growth both through its effect on the sectoral allocation of resources and through its effect on the demand for exports, as stressed in the wide literature on export-led growth.

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Table 1. F- and t-statistics for the analysis of a long run relationship in equation:

$$PR(ij)_t = \alpha_1 + \beta_1 YR(ij)_t + u_t$$

Countries	Lag order (*p*)	F-Statistic	t-Statistic
Germany-Japan	2	0.955	-0.724
UK-Japan	2	2.308	-1.580
UK-Germany	3	0.615	-1.107
UK-US	3	0.773	-0.634
Japan-US	2	3.310	-0.353
Germany-US	2	0.036	-0.228

Table 2. F- and t-statistics for the analysis of a long run relationship in equation:

$$PR(ij)_t = \alpha_2 + \beta_2 \Delta YR(ij)_t + v_t$$

Countries	Lag order (*p*)	F-Statistic	t-Statistic
Germany-Japan	2	2.041	-1.473
UK-Japan	2	1.173	-0.196
UK-Germany	3	0.997	-0.807
UK-US	3	3.456	-1.401
Japan-US	2	1.331	-0.649
Germany-US	3	4.473	-0.300

Table 3. F- and t-statistics for the analysis of a long run relationship in equation:

$$YR(ij)_t = \alpha_3 + \beta_3 PR(ij)_t + u_t$$

Countries	Lag order (*p*)	F-Statistic	t-Statistic
Germany-Japan	2	7.255*	-2.844
UK-Japan	2	11.042^{+}	-3.389*
UK-Germany	3	3.040	-2.464
UK-US	3	6.425*	-3.444*
Japan-US	2	4.390	-2.963**
Germany-US	3	2.763	-2.270

Table 4. F- and t-statistics for the analysis of a long run relationship in equation:

$$\Delta YR(ij)_t = \alpha_4 + \beta_4 PR(ij)_t + v_t$$

Countries	Lag order (*p*)	F-Statistic	t-Statistic
Germany-Japan	2	14.274^{+}	-5.342^{+}
UK-Japan	2	13.452^{+}	-5.179^{+}
UK-Germany	3	7.957^{+}	-3.980*
UK-US	3	21.497^{+}	-6.556^{+}
Japan-US	3	10.003^{+}	-4.414^{+}
Germany-US	3	8.608^{+}	-4.148^{+}

<u>Notes to Tables 1-4:</u>

1. The lag order (*p*) of the underlying ECM was selected using the Schwarz Information Criteria, the Akaike Information Criteria and the LM tests for testing residual serial correlation of orders 1 and 4.
2. The symbols $^+$ and * denote significance at 1% and 5% levels respectively.
3. The F-statistic is compared with the critical bounds of the F_{III} statistic for zero restrictions on the coefficients of the lagged level variables provided in Pesaran *et al* (1999) Table C1.iii. The t-statistic is compared with the critical bounds of the t_{III} statistic for zero restrictions on the lagged level of the dependent variable provided in Pesaran *et al* (1999) Table C2.iii.