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# Testing the Balassa–Samuelson effect: Implications for growth and the 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 long run stable relationship among the variables of the model. This paper addresses these problems by using the bounds testing approach developed by Pesaran et al. [J. Appl. Economet. 16 (2001) 289]. Our empirical results do not show supportive evidence for the Balassa–Samuelson effect in the long run. This seems to suggest the holding of the PPP. However, one of the implications of the PPP is that the real exchange rate does not have any real impact on the economy. Further empirical analysis rejects this last implication. In fact, real exchange rate seems to have a long run impact on relative growth rates.


*JEL classification:* F11; F31; C22
*Keywords:* Real exchange rate; Output; Causality

## 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 such phenomenon through the differential productivity growth between the traded and non-traded goods sectors. [1]

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[1] Actually the main motivation behind the model was to explain the persistent deviations from the PPP (see Balassa, 1964; Samuelson, 1964, 1994).

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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 result in higher relative non-traded goods prices. [2] One prediction of this model is that the tradable–non-tradable price difference is lower for rich countries than for poor (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 (see Rogoff, 1996).

The Balassa–Samuelson model rests on two assumptions: (1) labor markets are competitive within each country, thus labor mobility leads to wage equalization between traded and non-traded goods sectors; and (2) the PPP holds only for tradable goods. Recent empirical work has focused in examining these assumptions. [3] Strauss (1998) rejects these 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 support that the relative prices reflect relative labor productivities in the long run. However, the evidence on PPP in traded goods is less favorable. [4]

This paper takes a different route. 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 great 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 order of integration of the time series is 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. This paper addresses these problems by using the bounds testing approach developed by Pesaran et al. (2001). Our empirical results do not show supportive evidence for the Balassa–Samuelson effect in the long run. This seems to suggest the holding of the PPP. However, one of the implications of the PPP is that the real exchange rate does not have any real impact on the eco-

[2] The other competing theories to explain the phenomenon are the relative-factor-endowments model associated to Bhagwati (1984) and the Linder-type hypothesis that stress the role of demand factors (see, Bergstrand, 1991).

[3] 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; Balvers and Bergstrand, 1997).

[4] 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.

nomy. Further empirical analysis rejects this last implication. In fact, real exchange rate seems to have a long run impact on relative growth rates.

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.

## 2. The model

There are two countries (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): [5]

$$Y_{\mathrm{T}} = f(L_{\mathrm{T}}), \quad Y_{\mathrm{N}} = g(L_{\mathrm{N}}),$$
$$Y_{\mathrm{T}}^{*} = F(L_{\mathrm{T}}^{*}), \quad Y_{\mathrm{N}}^{*} = G(L_{\mathrm{N}}^{*}).$$

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

$$P_{\mathrm{T}} f'(L_{\mathrm{T}}) = w = P_{\mathrm{N}} g'(L_{\mathrm{N}}), \tag{1}$$

$$P_{\mathrm{T}}^{*} F'(L_{\mathrm{T}}^{*}) = w^{*} = P_{\mathrm{N}}^{*} G'(L_{\mathrm{N}}^{*}), \tag{2}$$

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

$$P_{\mathrm{T}} = eP_{\mathrm{T}}^{*} \tag{3}$$

where $e$ denotes the nominal exchange rate.

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

$$P = P_{\mathrm{T}}^{1-i} P_{\mathrm{N}}^{i}, \tag{4}$$

$$P^{*} = P_{\mathrm{T}}^{*1-j} P_{\mathrm{N}}^{*j}. \tag{5}$$

To simplify matters we can make, without loss of generality, the price of tradable goods to be equal to one: $P_{\mathrm{T}} = P_{\mathrm{T}}^{*} = 1$. This implies by Eq. (3) that the nominal exchange rate is equal to one as well: $e = 1$. Eqs. (4) and (5) can be rewritten as

$$P = P_{\mathrm{N}}^{i}, \tag{4'}$$

$$P^{*} = P_{\mathrm{N}}^{*j}. \tag{5'}$$

[5] 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.

Similarly, from Eqs. (1) and (2) follows

$$P_{\mathrm{N}} = f'(L_{\mathrm{T}})/g'(L_{\mathrm{N}}), \tag{1'}$$

$$P^*_{\mathrm{N}} = F'(L^*_{\mathrm{T}})/G'(L^*_{\mathrm{N}}). \tag{2'}$$

The real exchange rate is defined as

$$\theta = P/eP^* = P/P^*. \tag{6}$$

Substituting Eqs. (1′) and (2′) into Eqs. (4′) and (5′) and them into Eq. (6) yields

$$\theta = P/P^* = \frac{[f'(L_{\mathrm{T}})/g'(L_{\mathrm{N}})]^i}{[F'(L^*_{\mathrm{T}})/G'(L^*_{\mathrm{N}})]^j}. \tag{7}$$

Eq. (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, marginal productivity of labor is proportional to the average product of labor. [6] In this case, the right-hand side of Eq. (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 argue that the relative average labor productivity between countries is proportional to relative real output per capita levels. [7] 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

$$\theta = P/P^* = \frac{[Y]^i}{[Y^*]^j}. \tag{8}$$

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

$$\theta = P/P^* = \frac{[\Delta Y]^i}{[\Delta Y^*]^j}. \tag{9}$$

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.

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

[7] We can substantiate the assumption of proportionality by using the following production functions:

$$Y_{\mathrm{T}} = f(L_{\mathrm{T}}) = L^a_{\mathrm{T}}, \qquad Y_{\mathrm{N}} = g(L_{\mathrm{N}}) = L_{\mathrm{N}},$$
$$Y^*_{\mathrm{T}} = F(L^*_{\mathrm{T}}) = L^{*\alpha}_{\mathrm{T}}, \qquad Y^*_{\mathrm{N}} = G(L^*_{\mathrm{N}}) = L_{\mathrm{N}}.$$

By using these functions one can easily obtain Eqs. (8) and (9).

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 et al. (2001) which allows for testing the existence of long run relationships between variables regardless of whether they are integrated of order 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:

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

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

where Eqs. (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. [8]

Both the price ratio and the output ratio can be either $I(1)$ or $I(0)$ variables. 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 $(1, -1)$. This would be the case if both PPP holds and there is stochastic catch-up in productivity in the sense of Bernard and Durlauf (1995). Second, in Eq. (11) by first-differentiating the output ratio we will surely end up with an $I(0)$ variable in the right-hand side. Table 1 shows augmented Dickey–Fuller tests on the different pairs of combinations of relative prices and output for Germany, Japan, United Kingdom and the US. The results show that the most likely scenario is one where the price ratio is an $I(1)$ variable and the output ratio can either be $I(1)$ or $I(0)$. The first difference of the output ratio is clearly $I(0)$. This means that in most of the specifications used we are likely to

[8] Eqs. (10) and (11), by constraining both countries' productivity to have the same coefficient, implicitly assume that the share of traded goods in the GDP of both economies is the same ($i = j$ in Eqs. (4) and (5)). This is an assumption consistent with the data, since the shares of traded sectors in the four economies considered has been similar throughout the period considered.

Table 1
ADF unit root tests

| | $\mathrm{PR}(ij)_t$ | | $\mathrm{YR}(ij)_t$ | |
|---|---|---|---|---|
| | Level | First difference | Level | First difference |
| Germany–Japan | −1.795 | −4.752* | −4.364* | −6.404* |
| UK–Japan | −0.557 | −4.906* | −5.358* | −5.970* |
| UK–Germany | −0.913 | −3.567* | −1.594 | −9.905* |
| UK–US | −1.130 | −3.662* | −2.184 | −12.929* |
| Japan–US | −0.384 | −4.576* | −3.375* | −6.007* |
| Germany–US | −1.721 | −3.927* | −2.044 | −9.634* |

*Notes*: The lag order ($p$) of the ADF test 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. The symbol *denotes rejection of the unit root null at the 5% level.

be dealing with pairs of variables with different orders of integration. However, it is a well-known fact that the usual unit root tests suffer from a very important problem of power and there remains uncertainty about the true degree of integration of the series. For these reasons, tests of the Balassa–Samuelson effect based on traditional cointegration techniques would be flawed. This is because methods such as Johansen (1991) require that both variables be $I(1)$.

In a recent paper, Pesaran et al. (2001) develop a technique to test for the existence of a long run relationship between two variables irrespective of whether they are $I(1)$ or $I(0)$. Their technique also avoids the problems of uncertainty posed by the lack of power of unit root tests commented above. For this reason, Pesaran et al.'s (2001) methodology becomes most useful in this model where variables with different orders of integration are involved. Their approach is based on the estimation of a dynamic error correction (EC) 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 (2001) 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. (2001) consider two alternatives. First, an $F$-statistic test of joint significance of the lagged levels of the variables involved. [9] 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. [10] Also, if the $F$-statistic for

[9] In case that the model contains a deterministic trend, the $F$-test also includes the null of $\alpha_1 = 0$.

[10] We refer to Pesaran et al. (2001) 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.

the joint null of zero coefficients on $y_{t-1}$ and $x_{t-1}$ shows to be insignificant, then we cannot reject the null 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 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 Eqs. (10) and (11) for a set of pairs of developed countries and analyze causality aspects of these reduced forms.

## 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 the GDP deflator measured in a common currency 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. (2001), this was done using a 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. [11] Second, we obtain the $F$- and $t$-statistics as commented above. It has to be noted that in none of the cases a deterministic trend was found to be significant and, hence, we decided to drop it from the equations estimated. [12]

The results of the $F$- and $t$-statistics are reported in Tables 2 and 3 for each one of the specifications of the model and for each pair of countries. The results in Table 2 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 both looking at the $F$- and $t$-statistics. Table 3 provides the results of the tests for the regression of the level of price ratio on the first difference of the output ratio. The results again clearly show that in none of the cases we can reject the null hypothesis of no 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 among others Lothian and Taylor (1996)). However, our exercise does not necessarily stop at this point. If one

[11] The optimal number of lags using a general-to-specific criteria gave similar results.

[12] All the results not reported in the tables are available from the authors upon request.

Table 2
$F$- and $t$-statistics for the analysis of a long run relationship in equation: $\mathrm{PR}(ij)_t = \alpha_1 + \beta_1 \mathrm{YR}(ij)_t + u_t$

| 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 |

*Notes*: 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. The $F$-statistic is compared with the critical bounds of the $F_{\mathrm{III}}$ statistic for zero restrictions on the coefficients of the lagged level variables provided in Pesaran et al. (2001, Table C1.iii). The $t$-statistic is compared with the critical bounds of the $t_{\mathrm{III}}$ statistic for zero restrictions on the lagged level of the dependent variable provided in Pesaran et al. (2001, Table C2.iii).

Table 3
$F$- and $t$-statistics for the analysis of a long run relationship in equation: $\mathrm{PR}(ij)_t = \alpha_1 + \beta_1 \Delta\mathrm{YR}(ij)_t + u_t$

| 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 |

*Notes*: 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. The $F$-statistic is compared with the critical bounds of the $F_{\mathrm{III}}$ statistic for zero restrictions on the coefficients of the lagged level variables provided in Pesaran et al. (2001, Table C1.iii). The $t$-statistic is compared with the critical bounds of the $t_{\mathrm{III}}$ statistic for zero restrictions on the lagged level of the dependent variable provided in Pesaran et al. (2001, Table C2.iii).

accepts the implications of the PPP, the real exchange rate is mean reverting 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 4 and 5 suggest a different picture. Tables 4 and 5 present the results of the tests inverting the dependent and independent variables.

In Table 4, in which the price ratio is the explanatory and the output ratio is the dependent variable, we can find mixed results. For the case of the UK and 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 5 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 per capita 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.

Table 4
$F$- and $t$-statistics for the analysis of a long run relationship in equation: $\mathrm{YR}(ij)_t = \alpha_3 + \beta_3 \mathrm{PR}(ij)_t + u_t$

| 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 |

*Notes*: 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. The symbols $^{+,*}$ and $^{**}$ denote significance at 1%, 5% and 10% levels respectively. The $F$-statistic is compared with the critical bounds of the $F_{\mathrm{III}}$ statistic for zero restrictions on the coefficients of the lagged level variables provided in Pesaran et al. (2001, Table C1.iii). The $t$-statistic is compared with the critical bounds of the $t_{\mathrm{III}}$ statistic for zero restrictions on the lagged level of the dependent variable provided in Pesaran et al. (2001, Table C2.iii).

Table 5
$F$- and $t$-statistics for the analysis of a long run relationship in equation: $\Delta\mathrm{YR}(ij)_t = \alpha_4 + \beta_4 \mathrm{PR}(ij)_t + u_t$

| 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 | $7.957^{+}$ | $-3.98^{*}$ |
| UK–US | 3 | $21.497^{+}$ | $-6.556^{+}$ |
| Japan–US | 3 | $10.003^{+}$ | $-4.414^{+}$ |
| Germany–US | 3 | $8.608^{+}$ | $-4.148^{+}$ |

*Notes*: 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. The symbols $^{+}$ and $^{*}$ denote significance at 1% and 5% levels respectively. The $F$-statistic is compared with the critical bounds of the $F_{\mathrm{III}}$ statistic for zero restrictions on the coefficients of the lagged level variables provided in Pesaran et al. (2001, Table C1.iii). The $t$-statistic is compared with the critical bounds of the $t_{\mathrm{III}}$ statistic for zero restrictions on the lagged level of the dependent variable provided in Pesaran et al. (2001, Table C2.iii).

Given the results of Table 5 we proceeded to estimate the long-run relationship between the variables involved. We used the Autoregressive Distributed Lag approach of Pesaran and Shin (1999). ARDL-based estimators of the long run coefficients are super-consistent, and valid inferences on the long run parameters can be made using standard normal asymptotic theory. This method complements well with the bounds testing procedure because it has the additional advantage of yielding consistent estimates of the long run coefficients that are asymptotically normal irrespective of whether the underlying regressors are $I(1)$ or $I(0)$. The results are reported in Table 6. [13] Except for the case of UK–Germany, the level of the relative prices has a significant impact on the first difference of the output ratio. The effect is negative except for the case of UK–Japan—although not very large—indicating that a higher relative price will have a negative impact on relative growth rates.

[13] See footnote 1 of Table 6 for the lag order selection criteria.

Table 6
Long-run estimates using the ARDL approach for equation: $\Delta\mathrm{YR}(ij)_t = \alpha_4 + \beta_4\mathrm{PR}(ij)_t + u_t$

| Countries | ARDL lag order | $\alpha_4$ | $\beta_4$ |
|---|---|---|---|
| Germany–Japan | (2, 0) | −0.004 (−2.912)* | −0.027 (−2.253)* |
| UK–Japan | (4, 1) | 0.001 (0.315) | 0.012 (2.703)* |
| UK–Germany | (3, 0) | 0.001 (0.338) | 0.003 (1.340) |
| UK–US | (1, 3) | 0.001 (0.225) | −0.002 (−2.285)* |
| Japan–US | (2, 0) | 0.012 (3.835)* | −0.009 (−2.554)* |
| Germany–US | (1, 0) | 0.001 (0.543) | −0.007 (−1.801)** |

Following Pesaran and Shin (1999), the lag order of the ARDL model was selected using the Schwarz Information Criteria and the LM tests for serial correlation. The symbols * and ** denote significance at the 5% and 10% levels respectively.

Finally, we carried out *Granger-non-causality* tests between the first difference of the per capita output and the level of the price ratio. As mentioned above, the fact that the relative price appears to be the long run forcing variable is a necessary but not sufficient condition for rejecting *Granger-non-causality* from $\mathrm{PR}(ij)$ to $\Delta\mathrm{YR}(ij)$. Hence, we constructed an ECM assuming both that the exogenous variable is $\mathrm{PR}(ij)$ and then that it is $\Delta\mathrm{YR}(ij)$. We then tested for the significance of the coefficient of the lagged EC term and the joint significance of the lagged differences of the independent variables. [14] The results are reported in Table 7. In all cases the coefficient of the EC term when $\mathrm{PR}(ij)$ is assumed exogenous is statistically highly significant. The coefficient also shows a high velocity of adjustment towards equilibrium. On the other hand, none of the coefficients on the EC term when $\Delta\mathrm{YR}(ij)$ is considered exogenous is significant and gave very small adjustment velocities. This is consistent with the evidence presented in Tables 3 and 5. Only in two cases (Japan–US and Germany–US) we find some evidence of temporal causality running also from $\Delta\mathrm{YR}(ij)$ to $\mathrm{PR}(ij)$ through the lagged first differences of $\Delta\mathrm{YR}(ij)$. Overall, the results indicate that the causal relation runs from relative prices to the first difference of relative output, confirming our previous results.

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) that use the Balassa–Samuelson effect to argue that another channel by which inflation can affect growth is through the real exchange rate. [15]

[14] This is a standard test in time series literature. *Granger-non-causality* cannot be rejected if either the coefficient on the lagged EC term or the coefficients on the lagged differences of the independent variables are not statistically significantly different from zero.

[15] 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.

Table 7
Granger-non-causality tests between $\Delta\mathrm{YR}(ij)_t$ and $\mathrm{PR}(ij)_t$

| Countries | Causal relation | $F$-test of lagged differences | $T$-test on the ECM term | Coefficient on $\mathrm{ECM}_{t-1}$ | $\overline{R}^2$ | LM-AR(4) | LM-Het | LM-Norm |
|---|---|---|---|---|---|---|---|---|
| Germany–Japan | $\mathrm{PR}(ij)_t \rightarrow \Delta\mathrm{YR}(ij)_t$ | 4.596* | −6.801* | −0.785 | 0.478 | 1.821 | 2.156 | 2.267 |
| | $\mathrm{PR}(ij)_t \leftarrow \Delta\mathrm{YR}(ij)_t$ | 1.380 | −1.424 | −0.009 | 0.259 | 1.140 | 3.289** | 2.301 |
| UK–Japan | $\mathrm{PR}(ij)_t \rightarrow \Delta\mathrm{YR}(ij)_t$ | 2.251 | −3.765* | −0.563 | 0.520 | 4.438 | 0.450 | 3.226 |
| | $\mathrm{PR}(ij)_t \leftarrow \Delta\mathrm{YR}(ij)_t$ | 1.917 | −0.071 | −0.007 | 0.256 | 3.881 | 1.116 | 3.545 |
| UK–Germany | $\mathrm{PR}(ij)_t \rightarrow \Delta\mathrm{YR}(ij)_t$ | 1.813 | −7.717* | −0.778 | 0.628 | 3.235 | 0.135 | 2.235 |
| | $\mathrm{PR}(ij)_t \leftarrow \Delta\mathrm{YR}(ij)_t$ | 1.801 | −1.650 | −0.002 | 0.455 | 6.558 | 1.005 | 3.968 |
| UK–US | $\mathrm{PR}(ij)_t \rightarrow \Delta\mathrm{YR}(ij)_t$ | 1.452 | −13.278* | −0.884 | 0.558 | 1.335 | 1.848 | 1.851 |
| | $\mathrm{PR}(ij)_t \leftarrow \Delta\mathrm{YR}(ij)_t$ | 0.755 | −0.867 | −0.055 | 0.222 | 5.879 | 2.155 | 5.571* |
| Japan–US | $\mathrm{PR}(ij)_t \rightarrow \Delta\mathrm{YR}(ij)_t$ | 1.233 | −6.754* | −0.751 | 0.419 | 4.138 | 0.487 | 4.166 |
| | $\mathrm{PR}(ij)_t \leftarrow \Delta\mathrm{YR}(ij)_t$ | 3.356* | −0.422 | −0.002 | 0.345 | 4.135 | 0.999 | 3.558 |
| Germany–US | $\mathrm{PR}(ij)_t \rightarrow \Delta\mathrm{YR}(ij)_t$ | 3.751* | −6.403* | −0.686 | 0.638 | 5.261 | 0.085 | 4.202 |
| | $\mathrm{PR}(ij)_t \leftarrow \Delta\mathrm{YR}(ij)_t$ | 7.303* | −1.428 | −0.001 | 0.503 | 2.789 | 1.035 | 6.475* |

The number of lags in the ECM is the same as the lag order selected in Tables 4 and 2. LM-AR(4), LM-Het and LM-Norm are LM tests of serial correlation of order 4, heteroscedasticity and residual normality. The symbols * and ** denote significance at the 5% and 10% levels respectively.

## 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 et al. (2001) which allows for testing the existence of long run relationships between variables regardless of whether $I(1)$ or $I(0)$.

Our results do not show supportive evidence for the Balassa–Samuelson effect in the long run. This seems to suggest the holding of the long run PPP. However, one of the implications of the 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 the PPP. In fact, real exchange rate seems to have a long run impact on relative growth rates.

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## References

Andrés, J., Hernando, I., Kruger, M., 1996. Growth, inflation and the exchange rate regime. Economics Letters 53, 61–65.

Asea, P.K., Mendoza, E.G., 1994. The Balassa–Samuelson model: A general equilibrium appraisal. Review of International Economics 2, 244–267.

Bhagwati, J.N., 1984. Why are services cheaper in the poor countries? Economic Journal 94, 279–286.

Balassa, B., 1964. The purchasing power parity doctrine: A reappraisal. Journal of Political Economy 72, 584–596.

Balvers, R.J., Bergstrand, J.H., 1997. Equilibrium real exchange rates: Closed-form theoretical solutions and some empirical evidence. Journal of International Money and Finance 16, 345–366.

Banerjee, A., Dolado, J., Mestre, R., 1998. Error-correction mechanism tests for cointegration in single-equation framework. Journal of Time Series Analysis 19, 267–283.

Barro, R.J., Sala-I-Martin, X., 1995. Economic Growth. McGraw-Hill, New York.

Bergstrand, J.H., 1991. Structural determinants of real exchange rates and national price levels: Some empirical evidence. American Economic Review 81, 325–334.

Bernard, A.B., Durlauf, S.N., 1995. Interpreting tests of convergence. Journal of Econometrics 71, 161–173.

Canzoneri, M.B., Cumby, R.E., Diba, B., 1999. Relative labor productivity and the real exchange rate in the long run: Evidence for a panel of OECD countries. Journal of International Economics 47, 245–266.

De Gregorio, J., Giovannini, A., Krueger, T.H., 1994. The behavior of nontradable-goods prices in Europe: Evidence and interpretation. Review of International Economics 2, 284–305.

Heston, A., Nuxoll, D.A., Summers, R., 1994. The differential productivity hypothesis and purchasing power parties: Some new evidence. Review of International Economics 2, 227–243.

Johansen, S., 1991. Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models. Econometrica 59, 1551–1580.

Kravis, I.B., Lipsey, R.E., 1988. National price levels and the prices of tradables and non-tradables. American Economic Review 78 (May), 474–478.

Lothian, J.R., Taylor, M.P., 1996. Real exchange rate behavior: The recent float from the perspective of the last two centuries. Journal of Political Economy 104, 488–509.

Obstfeld, M., Rogoff, K., 1996. Foundations of International Macroeconomics. MIT Press, Cambridge.

Pesaran, M.H., Shin, Y., 1999. An autoregressive distributed lag modeling approach to cointegration analysis. In: Strom, S. (Ed.), Econometrics and Economic Theory in the 20th Century: The Ragnar Frisch Centennial Symposium. Cambridge University Press, Cambridge.

Pesaran, M.H., Shin, Y., Smith, R.J., 2001. Bounds testing approaches to the analysis of long run relationships. Journal of Applied Econometrics 16, 289–326.

Rogoff, K., 1996. The purchasing power parity puzzle. Journal of Economic Literature 34, 647–668.

Sala-I-Martin, X., 1997. I just ran two million regressions. American Economic Review 87 (May), 178–183.

Samuelson, P.A., 1964. Theoretical notes on trade problems. Review of Economics and Statistics 46, 145–154.

Samuelson, P.A., 1994. Facets of Balassa–Samuelson thirty years later. Review of International Economics 2, 201–226.

Strauss, J., 1998. Relative price determination in the medium run: The influence of wages, productivity, and international prices. Southern Economic Journal 65, 223–244.