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Real Exchange Rates In Developing Countries: Are Balassa-Samuelson Effects Present?

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Abstract

There is little empirical research on whether Balassa-Samuelson effects can explain the long-run behavior of real exchange rates in developing countries. This paper presents new evidence on this issue based on a panel data sample of 16 developing countries. The paper finds that the traded-nontraded productivity differential is a significant determinant of the relative price of nontraded goods, and the relative price in turn exerts a significant effect on the real exchange rate. The terms of trade also influence the real exchange rate. These results provide strong verification of Balassa-Samuelson effects for developing countries

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I. INTRODUCTION

The well-known analysis of Balassa (1964) and Samuelson (1964) provides an appealing explanation of the long-run behavior of the real exchange rate in terms of the productivity performance of traded relative to nontraded goods. Basically, the argument is that as the productivity of traded goods rises relative to that of nontraded goods, there will be a tendency for the real exchange rate to appreciate. Balassa-Samuelson effects are generally thought to be the key source of observed cross-sectional differences in real exchange rates (i.e., the same-currency prices of comparable commodity baskets) between countries at different levels of income per capita.[2] There is considerable empirical research on Balassa-Samuelson effects based on time-series data, but this research has been confined to industrial countries.[3] The time-series evidence on the working of the Balassa-Samuelson mechanism for developing countries has been largely unexplored.[4] One reason for this neglect is that sectoral price and productivity data are not readily available for developing countries. To address this problem, this paper makes use of recently available data from a number of sources to assemble a suitable data set for developing countries, which is then utilized to obtain new time-series evidence on the operation of Balassa-Samuelson effects in these countries.

Balassa-Samuelson effects can be embedded in a variety of models. These effects are typically derived within a static model, but they can be easily incorporated in the dynamic framework of the new open economy macroeconomic models.[5] Using a framework compatible with the new open economy macroeconomic approach, this paper derives two steady-state relations that capture key channels of the Balassa-Samuelson mechanism. The first relation links the real exchange rate to relative prices of nontraded goods at home and abroad. The basic version of this relation is based on the assumption that all traded goods are

[2] For a review of the evidence and a discussion of alternative explanations, see Edwards and Savastano (1999). See also Bergin, Glick, and Taylor (2004), who point out that although recent data reveal strong association between national price levels and income per capita, this association disappears in historical data going back fifty years or more.

[3] See, for example, Canzoneri, Cumby, and Diba (1999), and Lane and Milesi-Ferretti (2002).

[4] See, however, Ito, Isard, and Symansky (1997) who use time-series data to explore the Balassa-Samuelson hypothesis for APEC economies that include some developing countries.

[5] These models tend to focus on the short- to medium-term dynamics arising from nominal rigidities and have not paid much attention to long-run Balassa-Samuelson influences. Benigno and Theonissen (2002), however, do use a new open economy macroeconomic model to explore the effect of a productivity improvement in the traded good sector on the United Kingdom real exchange rate.

produced in both home and foreign countries. We also consider an alternative version, which assumes that each country specializes in the production of a different traded good. This variant of the relation then includes the terms of trade as an additional determinant of the real exchange rate. In both versions, the law of one price holds for each traded good in the long run.[6] The second relation explains the relative price of nontraded goods. Following Canzoneri, Cumby, and Diba (1999), we use restrictions on production technology to derive a simple form of the relation, which makes the labor productivity differential between traded and nontraded goods the main determinant of the relative price of nontraded goods.[7] The technology restriction used to obtain the second relation is not needed to derive the first relation.

For this study, we assembled a data set that includes time-series data for about two decades for 16 developing countries. As individual time series are not very long, we pool these series across countries to estimate our relations. Recent panel-data econometric techniques are used to identify long-run effects in these relations. The results provide strong evidence that the Balassa-Samuelson mechanism operates in developing countries. Using the United States as the reference country, we find that the U.S.-developing country differences in the relative price of nontraded goods and the terms of trade are significant determinants of the real exchange rate in the long run. The differences in the labor productivity differential, moreover, exert a significant long-run effect on the relative price differences. One puzzling result is that the estimated effect of the relative price variable is greater and that of the labor productivity variables smaller than the predicted value. We suggest explanations based on data problems to account for these discrepancies between estimated and predicted values.

The theoretical framework is outlined in Section II. The data and the empirical model are discussed in Section III. Section IV presents the results, and Section V provides some conclusions.

II. Theoretical Framework

This section outlines a framework to provide theoretical underpinnings for our empirical analysis. As we are concerned with long-term effects, we do not model short-run dynamics, but focus on steady-state relations under complete adjustment of wages and prices. We consider a multi-country framework with each country using fixed endowments of labor and

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[6] The real exchange rate for the traded-goods basket, however, need not be stationary if weights for individual traded goods differ between the home and foreign countries. Our empirical procedure accounts for this possibility.

[7] An alternative approach would relate the relative price to the total factor productivity differential. Demand-side factors would also influence the relative price in such a relation.

capital to produce traded and nontraded goods under perfect competition.[8] We use two special models of the pattern of traded-goods production. The first model follows the standard Balassa-Samuelson formulation and assumes that each country is diversified and produces all traded goods. The second model assumes that each country is specialized in the production of a country-specific traded good, as in the Armington (1969) model. We discuss below only the part of the model that is needed to derive the relations used in our empirical analysis.

Basic Setup

Households in country *i* supply a fixed amount of labor and maximize the following expected lifetime utility:

$$E_t \sum_{\tau=t}^{\infty} \delta^{\tau-t} U(C_{i\tau}),$$

where δ is the discount rate, and $C_{i\tau}$ represents a consumption index for period τ. The consumption index is defined as:

$$C_i = (C_i^T)^{\gamma_i} (C_i^N)^{1-\gamma_i} / (\gamma_i^{\gamma_i} (1-\gamma_i)^{1-\gamma_i}), \quad (1)$$

where C_i^T and C_i^N are the sub-indexes for consumption bundles of traded and nontraded goods, γ_i is the share of traded goods in aggregate consumption, and time subscripts are dropped for simplicity. The traded goods basket is assumed to be the following CES index of *m (> 1)* goods:

$$C_i^T = \left[\sum_{j=1}^{m} (v_i^j)^{1/\sigma} (C_i^{Tj})^{1-1/\sigma} \right]^{\sigma/(\sigma-1)}, \quad (2)$$

where C_i^{Tj} is the amount consumed of traded good *j*, v_i^j represents the weight placed on the good, and $\sigma\ (>1)$ is the elasticity of substitution between traded goods.

Let P_i denote the consumer price index, and P_i^T and P_i^N the price indexes for traded and nontraded goods. Using (1) and (2), we define P_i and P_i^T as the cost-minimizing prices of C_i and C_i^T, which are given by:

$$P_i = (P_i^T)^{\gamma_i} (P_i^N)^{1-\gamma_i}, \quad (3)$$

$$P_i^T = \left[\sum_{j=1}^{m} v_i^j (P_i^{Tj})^{1-\sigma} \right]^{1/(1-\sigma)}. \quad (4)$$

---

[8] Our framework can be readily extended to incorporate monopolistic competition. As such an extension would make little difference to the long-run relations derived in the paper, we assume perfect competition for simplicity.

The pattern of production for traded goods is characterized by either diversification (with each country producing all traded goods) or specialization (with each country producing a different traded good). In the case of specialization, the number of countries equals the number of traded goods. For this case, use the same index for traded goods and countries so that good i is produced by country i. Letting Y_i^N and Y_i^{Tj} denote outputs of the nontraded and j th traded good, we assume the following Cobb-Douglas production function for these goods:[9]

$$Y_i^N = A_i^N (K_i^N)^{\alpha_N} (L_i^N)^{\beta_N} , \tag{5}$$

$$Y_i^{Tj} = A_i^{Tj} (K_i^{Tj})^{\alpha_j} (L_i^{Tj})^{\beta_j} \quad \begin{cases} j = 1,\ldots,m & \textit{if diversified} \\ j = i & \textit{if specialized} \end{cases} , \tag{6}$$

where K_i^N and L_i^N represent the amounts of capital and labor used in the production of the nontraded good, while K_i^{Tj} and L_i^{Tj} are the corresponding amounts for the traded good j.

Let country 1 be the reference country, and define S_i as the exchange rate of country i (expressed as the price of country i 's currency) with respect to country 1. We distinguish between the short and the long run in the present model. The short run is characterized by nominal rigidities in the form of sticky wages and prices. The long run, on the other hand, represents steady-state equilibrium with full adjustment of wages and prices. In the short run, nominal rigidities can cause departures from the law of one price and the marginal productivity condition for labor. We assume below that there are no departures from these relations in steady state. We focus on the steady-state behavior of variables to derive Balassa-Samuelson effects.

Use a tilde over a variable to denote the steady-state value of the variable. Assuming that the law of one price holds in steady state, we can link steady-state prices of traded goods in different countries as follows:

$$\widetilde{S}_i \widetilde{P}_i^{Tj} = \widetilde{P}_1^{Tj} . \tag{7}$$

Also, assume that the marginal productivity condition is satisfied in steady state. Thus, letting W_i denote the wage rate, and using (5) and (6), we have:

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[9] The Cobb-Douglas form of the production function is used below to derive a simple relation between the relative price of nontraded goods and the labor productivity differential. Canzoneri, Cumby, and Diba (1999) discuss more general production conditions, which would also imply such a relation.

$$\widetilde{W}_i = \beta_N (\widetilde{Y}_i^N / \widetilde{L}_i^N)\widetilde{P}_i^N = \beta_j (\widetilde{Y}_i^{Tj} / \widetilde{L}_i^{Tj})\widetilde{P}_i^{Tj} \begin{cases} j = 1,...,m & if\ diversified \\ j = i & if\ specialized \end{cases} . \tag{8}$$

Key Relations

We now derive key relations in the log-linear form. Use lower case letters to denote values in logs, and define the consumption-based real exchange rate as:

$$q_i = s_i + p_i - p_1 . \tag{9}$$

The price of the nontraded good relative to the price of domestically produced traded good(s) plays an important role in the determination of the real exchange rate. Define this relative price in logs as:

$$rp_i \begin{cases} = p_i^N - p_i^T & if\ diversified \\ = p_i^N - p_i^{Ti} & if\ specialized \end{cases} . \tag{10}$$

To link q_i and rp_i, we use (4) and take a log-linear approximation of steady-state traded-goods price index around its initial value to get:

$$\widetilde{p}_i^T = \sum_{j=1}^m \overline{\theta}_i^{\,j} \widetilde{p}_i^{Tj} , \tag{11}$$

where $\overline{\theta}_i^{\,j}$ is the share of traded good j in total traded-goods consumption in the initial steady state, and the initial steady-state value of $\widetilde{p}_i^T$ is normalized to equal zero.[10] In the case of specialization, noting that only good i is exported by country i, we can express (11) as:

$$\widetilde{p}_i^T = \overline{\theta}_i^X \widetilde{p}_i^X + (1 - \overline{\theta}_i^X)\widetilde{p}_i^M , \tag{12}$$

where $\widetilde{p}_i^X = \widetilde{p}_i^{Ti}$ and $\widetilde{p}_i^M = \sum_{j \neq i} \overline{\theta}_i^{\,j} \widetilde{p}_i^{Tj} / (1 - \overline{\theta}_i^{\,i})$ are the price indexes for exports and imports, and $\overline{\theta}_i^X = \overline{\theta}_i^{\,i}$ is now the (initial-steady-state) share of the export good in total expenditure on traded goods.

[10] The share equals $v_i^j (\overline{P}_i^{Tj})^{1-\sigma} / \sum_{j=1}^m v_i^j (\overline{P}_i^{Tj})^{1-\sigma}$, where $\overline{P}_i^{Tj}$ is the price of traded good j in the initial steady state.

For the diversification case, use (3), (7), and (9)-(11) to obtain:

$$\widetilde{q}_i = \sum_{j=1}^{m} (\overline{\theta}_i^{\,j} - \overline{\theta}_1^{\,j})\widetilde{p}_1^{Tj} + (1-\gamma_i) r\widetilde{p}_i - (1-\gamma_1) r\widetilde{p}_1 \,,\; i \neq 1 . \tag{13}$$

The first term on the right hand side of (13) represents the log real exchange rate for traded goods in steady state, $\widetilde{q}_i^{\,T}$ $(= \widetilde{s}_i + \widetilde{p}_i^{\,T} - \widetilde{p}_1^{\,T})$. This term will not equal zero and may exhibit nonstationary behavior if the composition of a country's traded goods basket differs from that of the reference country. The Balassa-Samuelson analysis is often simplified by the assumption that expenditure shares are the same everywhere. In this simple case, $\widetilde{\theta}_i^{\,j} = \widetilde{\theta}_1^{\,j}$ for all *j*, $\gamma_i = \gamma_1$, and (13) can be expressed simply as $\widetilde{q}_i = (1-\gamma_1)(r\widetilde{p}_i - r\widetilde{p}_1)$.

In the case of specialization, the terms of trade also influence the real exchange rate. Letting $tt_i = p_i^X - p_i^M$ denote the log terms of trade, and using (3), (7), and (9)-(12), we obtain the relation for the specialization case as:

$$\begin{aligned}\widetilde{q}_i = &\sum_{j=1}^{m} (\overline{\theta}_i^{\,j} - \overline{\theta}_1^{\,j})\widetilde{p}_1^{Tj} + (1-\gamma_i) r\widetilde{p}_i - (1-\gamma_1) r\widetilde{p}_1 \\ &+ (1-\overline{\theta}_i^{\,X})(1-\gamma_i)\widetilde{tt}_i - (1-\overline{\theta}_1^{\,X})(1-\gamma_1)\widetilde{tt}_1\end{aligned} \,,\; i \neq 1 . \tag{14}$$

The first term on the right hand side of (14) still equals $\widetilde{q}_i^{\,T}$. Note that $r\widetilde{p}_i$ now represents the relative (steady-state) price of the nontraded good in terms of domestically-produced traded good. The terms of trade thus enter the relation because they affect the price of the traded goods basket relative to that of the traded good produced at home.

Next, the relative price of nontraded goods can be related to the productivity differential between domestically produced traded and nontraded goods. Define the log labor productivity in the two sectors as:

$$lp_i^T \begin{cases} = \sum_j \omega_i^j (y_i^{Tj} - l_i^{Tj}) & \textit{if diversified} \\ = y_i^{Ti} - l_i^{Ti} & \textit{if specialized} \end{cases} , \tag{15}$$

$$lp_i^N = y_i^N - l_i^N . \tag{16}$$

For the diversification case, ω_i^j is the weight for good *j*'s labor productivity in the aggregate labor productivity index for traded goods. Let $lp_i = lp_i^T - lp_i^N$ denote the labor productivity differential between traded and nontraded goods. In defining the diversification labor productivity index in steady state, we use the same weights as those in the traded-goods price index. Using (8), (15), and (16), and letting $\omega_i^j = \overline{\theta}_i^{\,j}$ for the diversification case, we can express the steady-state relative price as:

$$r\widetilde{p}_i = \vartheta + \widetilde{lp}_i , \tag{17}$$

where ϑ equals $\sum_{j=1}^{m} \overline{\theta}_i^{\,j} \log\beta_j - \log\beta_N$ in the case of diversification, and $\log\beta_i - \log\beta_N$ in the case of specialization.

III. Empirical Implementation

Data

We use a number of sources to put together a developing economies panel data set that includes time series from 1976 to 1994 for 16 countries.[11] This set includes 14 countries at low- and medium-income levels and 2 high-income economies (Republic of Korea and Singapore) that had lower income levels at the beginning of the sample period. Traded goods are assumed to consist of manufacturing and agriculture sectors. Nontraded goods are represented by all other sectors. The United States is chosen as the reference country. The real exchange rate is based on consumer price indexes and represents the real value of a currency in terms of U.S. dollars.

Although our classification of the traded and nontraded goods sectors is similar to the one used for industrial countries, one potential problem is that a substantial portion of the agriculture sector (and possibly of the manufacturing sector) in developing countries may consist of traditional activities producing nontraded goods. Another problem is that the quality of labor is likely to vary considerably in developing countries, and our labor productivity measure (based on employment figures unadjusted for quality changes) does not account for this variation.[12] We are unable to address these issues because of data limitations. However, we explore below certain implications of these measurement problems for the estimation of the empirical model.

Empirical Model

To undertake panel-data tests of the Balassa-Samuelson relations, we assume that long-run parameters (based on steady-state expenditure shares) are the same across our developing country set (D).[13] Thus, we set $\bar{\theta}_i^X = \bar{\theta}^X$ and $\gamma_i = \gamma$ for $i \in D$. However, to allow for possible developing-industrial country differences in expenditure shares, we do not require U.S. (country 1) parameters to be the same as the ones for our developing-country sample.

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[11] Details of the variables and data sources are provided in Appendix II.

[12] As noted in Appendix II, another limitation of the data on labor inputs is that employment measures for the manufacturing, agriculture, and other (nontraded goods) sectors come from different sources, and are not fully comparable. Also, note that labor productivity for traded goods is simply measured as the ratio of total output to total employment in the traded goods sector. For the diversification case, this index does not fully conform to the theoretical index used in (17), since the implicit weights for individual traded goods in this index could differ from the weights used in the traded-goods price index.

[13] We later allow these parameters to vary between developing countries at different income levels.

The following two equations are estimated to test for Balassa-Samuelson effects:

$$q_{it} = \mu_i + \kappa_t + \pi rpd_{it} + \tau ttd_{it} + u_{it}, \tag{18}$$

$$rpd_{it} = \psi_i + \chi_t + \lambda lpd_{it} + v_{it}, \; i \in D, \tag{19}$$

where $rpd_{it} = rp_{it} - rp_{1t}$, $ttd_{it} = tt_{it} - tt_{1t}$, and $lpd_{it} = lp_{it} - lp_{1t}$ are, respectively, the differences in the relative price of nontraded goods, the terms of trade, and the traded-nontraded productivity differential between developing country i and the United States; μ_i and ψ_i are country-specific fixed effects while κ_t and χ_t are common time effects; and u_{it} and v_{it} are error terms. Time effects represent the influence of common time-specific (short- and long-run) factors, and error terms capture the effects of short-term deviations from steady state (that are not included in time effects).

Equation (18) is derived from (13) and (14), and nests the diversification and specialization cases. Diversification implies that $\tau = 0$ while specialization implies that $\tau = (1 - \overline{\theta}^X)(1-\gamma) > 0$. In both cases, $\pi = (1-\gamma) > 0$. Equation (19) is based on (17). In this equation, $\lambda = 1$. The absence of Balassa-Samuelson effects would imply that $\pi = \tau = \lambda = 0$.[14] Note that time effects in (18) would pick up long-run movements in the real exchange rate for traded goods arising from parametric differences between developing countries and the United States.

Although the long-run parameters in (18) and (19), π, τ, and λ, are constrained to be the same across developing countries, these relations allow the short-run dynamics (reflected in the time series behavior of the error terms) to be different across countries. The explanatory variables, rpd_{it}, ttd_{it}, and lpd_{it}, can be stationary, trend-stationary or nonstationary. In the case of trend-stationary behavior, (18) and (19) can be modified to include a time trend. Coefficients of time trends in the two relations would be homogeneous across countries, and depend on the long-run parameters.[15] Note that if explanatory variables are integrated or trend-stationary, then q_{it} would also be integrated or trend-stationary. In this case, Balassa-Samuelson effects would cause permanent departures from the purchasing power parity.

[14] Tests of Balassa-Samuelson effects could also be based on alternative versions of (18) and (19), which exclude U.S. variables, rp_{1t}, tt_{1t}, and lp_{1t}, and are expressed as $q_{it} = \mu_i^* + \kappa_t^* + \pi rp_{it} + \tau tt_{it} + u_{it}^*$, and $rp_{it} = \psi_i^* + \chi_t^* + \lambda lp_{it} + v^*{}_{it}$. However, we estimate relations in the form that includes U.S. variables because this form allows us to explore whether U.S. variables exert an effect additional to their effect via rpd_{it}, ttd_{it}, and lpd_{it}.

[15] Letting $rpd_{it} = g_1 t + rpd'_{it}$, $ttd_{it} = g_2 t + ttd'_{it}$, and $lpd_{it} = g_3 t + lpd'_{it}$, we can restate (18) and (19) as follows: $q_{it} = \mu_i + \kappa_t + (g_1\pi + g_2\tau)t + \pi rpd'_{it} + \tau ttd'_{it} + u_{it}$, and $rpd_{it} = \psi_i + \chi_t + g_3\lambda t + \lambda lpd'_{it} + v_{it}$.

As discussed above, our measure for the traded goods sector (i.e., agriculture plus manufacturing) may be too broad for developing countries and could include nontraded goods. As discussed in Appendix I, the measured relative price of nontraded goods in this case would understate the true relative price, and bias the relative-price coefficient upward in (18). This measurement problem would not lead to a systematic bias in the estimation of (19), since the measured value of the traded-nontraded productivity differential would also understate its true value. A more serious problem for estimating (19) is that the labor productivity measure is not adjusted for quality variation. Appendix I shows that the estimated effect of measured labor productivity differential would be biased downward if there is a positive association between the average labor quality and the true labor productivity.

IV. RESULTS

Estimation

Before estimating (18) and (19), we examine whether the variables in these relations contain a unit root or not. Table 1 shows the results of two tests of a unit root in panel data. In the first test (LL) based on Levin and Lin (1993), the null hypothesis of a unit root is tested against the alternative of a homogeneous autoregressive coefficient. The second test (IPS) due to Im, Pesaran, and Shin (2003) tests the unit root null against a more general alternative of a heterogeneous autoregressive coefficient. Both tests indicate that q_{it} contains a unit root (with or without a time trend).[16] For the remaining variables, the tests are sensitive to whether a time trend is included or not. In the absence of a trend, the unit root hypothesis is not rejected for rpd_{it} and ttd_{it} by both the LL and IPS tests, and for lpd_{it} by the LL test. However, if a trend is present, both tests indicate that rpd_{it} and lpd_{it} are not integrated, and IPS test indicates that ttd_{it} is also not integrated.

We first consider the basic form of (18) and (19), which does not include a time trend. In this case, since there is indication of nonstationary behavior for variables in these relations, we also undertake tests for cointegration. We use two parametric tests, the panel t-test and the group t-test, suggested by Pedroni (1999). The panel t-test rejects the hypothesis that there is no cointegration for the vector (q_{it}, rpd_{it}), but not for vectors (rpd_{it}, lpd_{it})

[16] Because of the assumption of homogeneous autoregressive coefficients, the Levin and Lin (1993) test is encompassed by the IPS test. The results of the IPS test, however, are not conclusive. Although the test does not reject the unit-root hypothesis for q_{it} at the 5 percent level, it does indicate rejection at slightly higher levels (p-value = 0.069 with trend and p-value = 0.065 without trend).

Table 1. Unit Root Tests

Variable	Levin-Lin Test Statistic		Im-Pesaran-Shin Test Statistic	
	Without Trend	With Trend	Without Trend	With Trend
q_{it}	0.478	-1.008	-1.513	-1.480
rpd_{it}	0.231	-3.730 **	-0.358	-6.615 **
ttd_{it}	-0.070	-1.327	-0.388	-1.987 *
lpd_{it}	0.604	-3.297 **	-2.059 *	-6.169 **

Note: * indicates significance at the 5 percent level, and ** at the 1 percent level.

and $(q_{it}, rpd_{it}, ttd_{it})$. The group t-test rejects the no-cointegration hypothesis for all three vectors.[17] The group t-test (unlike the panel t-test) does not constrain the first-order correlation in the residuals to be homogeneous under the alternative hypothesis and is more relevant for our model, which allows the short-run dynamics to vary across countries. Thus, for the case where all variables are assumed to be nonstationary, we consider (18) and (19) to represent relations between cointegrating vectors.

We estimate these relations by Dynamic Ordinary Least Squares (DOLS), which is an appropriate framework for estimating and testing hypotheses for homogeneous cointegrating vectors.[18] The relations are estimated in the following form:

[17] For vectors, (q_{it}, rpd_{it}), (rpd_{it}, lpd_{it}), and $(q_{it}, rpd_{it}, ttd_{it})$, the panel-t test statistic is -1.730*, -1.093, and 0.278, respectively. The corresponding statistic for the group-t test is -2.074*, -1.955*, and -1.959*. An asterisk indicates significance at the 5 percent level.

[18] See Kao and Chiang (2000), and Mark and Sul (2002) for a discussion of the properties of panel DOLS.

$$q_{it} = \mu_i + \kappa_t + \pi rpd_{it} + \tau ttd_{it} + \sum_{r=-n}^{n} (\xi_{ir} \Delta rpd_{i,t+r} + \zeta_{ir} \Delta ttd_{i,t+r}) + u'_{it} \ , \qquad (20)$$

$$rpd_{it} = \psi_i + \chi_t + \lambda lpd_{it} + \sum_{r=-n}^{n} \varphi_{ir} \Delta lpd_{i,t+r} + v'_{it} \ , \qquad (21)$$

where n is the number of lags and leads used for the first-difference terms. Coefficients of these terms capture the short-run dynamics. We allow the short-run dynamics to be heterogeneous (i.e., let ξ_{ir}, ζ_{ir}, and φ_{ir} differ across i).

If a linear trend is included, unit root tests suggest that the explanatory variables in (18) and (19) are not integrated. We, thus, also consider the trend-stationary setting for estimating these relations. DOLS is a useful estimating procedure even in this case. Since first-difference terms are included in this procedure, the coefficients of level terms represent long-run effects. Therefore, we estimate (20) and (21) with trend variables to identify long-run Balassa-Samuelson influences in the trend-stationary case.

Basic Results

Tables 2 and 3 present DOLS estimates of different variants of the real exchange rate equation with one lag and one lead of the first difference terms.[19] Table 2 shows the estimates of the equation for the diversification case excluding the terms of trade variable, and Table 3 for the specialization case including this variable. For both cases, we report the results for homogeneous as well as heterogeneous short-run dynamics. Regressions 1 and 4 in these tables show estimates of the basic form of the equation without a time trend. In all of these cases, the effect of the relative price variable is positive and significant. The predicted value of this variable's coefficient equals $1-\gamma$ (which represents the share of the nontraded goods sector). The estimated value, however, is greater than unity in most cases. As discussed above, this discrepancy between the predicted and estimated values could reflect a bias arising from defining the traded-goods sector too broadly. The results also show that the terms of trade variable exerts a positive and significant effect when introduced in the real exchange rate equation (see Table 3). This finding supports the specialization version of the model, in which each country produces a different good.

[19] The short length of each time series makes it difficult to explore the possibility that the short-run dynamics involves higher lags and leads. Indeed, there are not enough degrees of freedom to estimate (20) with additional lags and leads in the case of heterogeneous dynamics. In the case of homogeneous dynamics, however, we did estimate (20) and (21) with two lags and leads, and found little difference in the results.

Table 2. The Exchange Rate Relation: The Diversification Case

Variable	Coefficient Estimates					
	(1)	(2)	(3)	(4)	(5)	(6)
	Homogeneous Short-Run Dynamics			Heterogeneous Short-Run Dynamics		
rpd_{it}	0.962 **	0.962 **	0.79 **	1.066 **	1.066 **	0.846 **
	(0.146)	(0.146)	(0.161)	(0.156)	(0.156)	(0.173)
Trend		0.057			0.071	
		(0.055)			(0.060)	
$rpd_{it}D$			0.329 *			0.401 *
			(0.129)			(0.156)
Adj. R-squared	0.997	0.997	0.997	0.997	0.997	0.997
S. E. of Reg.	0.154	0.154	0.1523	0.160	0.160	0.158

Notes: White heteroskedasticity-consistent errors are shown in parenthesis. All regressions include country-specific and time-specific dummy variables as well as three first-difference terms. Coefficients of the first-difference terms are constrained to be the same under homogeneous dynamics, and allowed to be different under heterogeneous dynamics. D is a dummy variable, which equals one for low-income developing countries and zero for others. The number of observations equals 256.

* indicates significance at the 5 percent level, and ** at the 1 percent level.

Table 3. The Exchange Rate Relation: The Specialization Case

Variable	Coefficient Estimates					
	(1)	(2)	(3)	(4)	(5)	(6)
	Homogeneous Short-Run Dynamics			Heterogeneous Short-Run Dynamics		
rpd_{it}	1.111 **	1.111 **	0.851 **	1.217 **	1.217 **	0.834 **
	(0.143)	(0.143)	(0.163)	(0.204)	(0.204)	(0.251)
ttd_{it}	0.300 **	0.300 **	0.477 **	0.332 *	0.332 *	0.565 **
	(0.091)	(0.091)	(0.103)	(0.129)	(0.129)	(0.141)
Trend		0.063			0.111	
		(0.0540)			(0.075)	
$rpd_{it}D$			0.407 **			0.601 *
			(0.143)			(0.271)
$ttd_{it}D$			-0.348 **			-0.407
			(0.123)			(0.209)
Adj. R-squared	0.998	0.998	0.998	0.997	0.997	0.997
S. E. of Reg.	0.142	0.142	0.139	0.152	0.152	0.148

See notes for Table 2. The number of observations in Table 3 regressions equals 246 because of missing terms of trade data.

Table 4 shows the results for estimating the relative price relation by DOLS. Regressions 1 and 4 in this table estimate the basic form of the relation without a time trend. The effect of the labor productivity index in both regressions is positive and significant. But the estimated values of its coefficients in the two regressions are substantially below the predicted value of unity. One possible explanation of this result, suggested above, is that measuring employment without adjustment for quality changes leads to a downward bias in the productivity coefficient.

Table 4. The Relative Price Relation

Variable	(1)	(2)	(3)	(4)	(5)	(6)
	Homogeneous Short-Run Dynamics			Heterogeneous Short-Run Dynamics		
lpd_{it}	0.287 **	0.287 **	0.345 **	0.302 **	0.302 **	0.397 **
	(0.042)	(0.042)	(0.051)	(0.048)	(0.48)	(0.062)
Trend		0.000			-0.004	
		(0.028)			(0.028)	
lpd_{it}*D			-0.152 *			-0.229 **
			(0.076)			(0.086)
Adj. R-squared	0.832	0.832	0.835	0.832	0.832	0.838
S. E. of Reg.	0.073	0.073	0.072	0.073	0.073	0.072

See notes for Table 2.

Tables 2–4 also report the results for the trend-stationary case, in which a homogeneous linear trend (with the same coefficient across countries) is included in the two relations. The tables show (see regressions 2 and 4 in each table) that the coefficient of the trend variable is insignificant in all cases, and the introduction of this variable in the regressions makes no difference to the estimates of Balassa-Samuelson parameters. We also introduced heterogeneous trends in the two relations, but this variation also made little difference to the results.

Further Analysis

Our empirical model includes time effects to allow the effect of U.S. variables to be different from that of developing countries variables because of parametric differences. Time effects are, in fact, significant in both relations. Nevertheless, we also estimated the two relations without time effects, but did not find a substantial difference in results. Our test of Balassa-Samuelson effects is based on two relations suggested by theory. Separate estimation of these relations is useful in distinguishing two key channels, through which labor productivity affects the real exchange rate. However, we also explored a variation of the test that combines the two relations, and relates the real exchange rate directly to the labor productivity measure and the terms of trade index. Estimation of this relation by DOLS also indicates that both variables are significant determinants of the real exchange rate.

We further examined whether the results are sensitive to variation in income levels. To explore this question, we divided the developing country sample into high- and low-income groups, and tested whether coefficients of Balassa-Samuelson variables differ between the two groups.[20] Regressions 3 and 6 in Tables 2–4 report the results of these tests. These regressions include interactions between explanatory variables and a dummy variable for the low-income group. Thus, coefficients of the variables show the effects for the high-income group, and interaction terms represent the additional effects for the low-income group. Interestingly, the results show that the effect of the relative price variable (in the real exchange rate regressions) is significantly higher for the low-income group while the effect of the labor productivity differential (in the relative price regressions) is significantly lower. The departures from predicted values are, thus, more pronounced for low-income countries. Since data problems are likely to be more severe for the developing countries at the lower end of the income scale, this finding is supportive of our suggested explanation that the estimates of Balassa-Samuelson effects are biased because of measurement errors. The results also indicate that the terms of trade effect is smaller for the low-income group. Therefore, the support for the specialization version seems to be weaker for the poorer developing countries.[21] Overall, these results indicate strongly that Balassa-Samuelson effects play a significant role in the determination of the real exchange rate in the long run.

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[20] The classification of countries in the two groups is based on average income per capita for the sample period. Each group includes eight countries (see Appendix II for the list of countries in the two groups).

[21] This result may seem paradoxical as production and exports of low-income countries tend to be less diversified. However, specialization could also mean production of goods (e.g., sophisticated manufactured products) that are significantly differentiated from goods produced elsewhere. Poor countries may be less specialized in this sense.

V. CONCLUSIONS

The Balassa-Samuelson hypothesis would seem to be especially relevant for developing countries where relative prices and productivities are likely to be more variable. Yet, there is little or no empirical evidence on whether Balassa-Samuelson effects can successfully explain long-run movements of the real exchange rate in developing countries. This paper presents new time-series evidence for developing countries on the presence of Balassa-Samuelson effects. To test for these effects, we estimate two long-run relations: relative prices (of nontraded goods) affect the real exchange rate in one relation, and labor productivity differentials (between traded and nontraded goods) affect relative prices in the second relation. Terms of trade also affect the real exchange rate (in the first relation) if a country is specialized in the production of traded goods. A key new finding of this paper is that the labor productivity differential exerts a significant effect on the real exchange rate via its influence on the relative price of nontraded goods.[22] The paper also finds that terms of trade are a significant determinant of the real exchange rate.

Although the effect of relative prices and labor productivity variables operates in the direction indicated by the Balassa-Samuelson hypothesis, the effect of relative prices is stronger and that of productivity differentials weaker than the predicted value. The paper also finds that the departures from predicted values are larger for developing countries with lower income levels. We suggest an explanation that attributes the above results to biases caused by measurement problems. These problems are likely to be more pronounced in countries with lower incomes and, thus, could account for differences in estimated Balassa-Samuelson effects between countries at low- and high-income levels.

Our tests of the Balassa-Samuelson explanation are based on two long-run relations, which are derived from theory under fairly general conditions, and can be implemented empirically for developing countries. Further theoretical and empirical analysis (requiring extra assumptions and more data) could extend these relations and explore the role of additional factors.[23] Such analysis is beyond the scope of the present paper. The results of this paper do suggest that the Balassa-Samuelson mechanism is an empirically useful framework for investigating the long-run behavior of the real exchange rate for developing countries.

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[22] Previous work (for example, Lane and Milesi-Ferretti, 2004), based on using GDP per capita as a proxy for the labor productivity differential, has not found a systematic effect of the productivity variable on real exchange rates in developing countries. We believe that we are able to identify this effect by using a more appropriate measure of labor productivity differential based on sectoral data.

[23] For example, Lane and Milesi-Ferretti (2004) explore the theoretical link between the real exchange rate and net foreign assets, and provide evidence that the net foreign assets position is an important determinant of the real exchange rate for developing (as well as developed) countries.

Potential Biases Due to Measurement Problems

The Traded-Goods Sector Measure Includes Nontraded Goods

Using a hat over a variable to denote the measured value, let the measured traded-goods price be $\hat{p}_{it}^{T} = \phi p_{it}^{N} + (1-\phi) p_{it}^{T}$, $1 > \phi > 0$, where ϕ is the weight for the nontraded goods that are improperly included in the traded-goods sector measure. The measured relative price of nontraded goods is then related to the true price as $r\hat{p}_{it} = p_{it}^{N} - \hat{p}_{it}^{T} = (1-\phi) rp_{it}$. Let the corresponding relation for country 1 be $r\hat{p}_{1t} = (1-\phi_1) rp_{1t}$, with $1 > \phi_1 \geq 0$. Using these relations and letting $r\hat{p}d_{it} = r\hat{p}_{it} - r\hat{p}_{1t}$, we can express (18) as:

$$q_{it} = \mu_i + \kappa_t' + \pi' r\hat{p}d_{it} + \tau ttd_{it} + u_{it},$$

where $\kappa_t' = \kappa_t + \pi[1/(1-\phi) - 1/(1-\phi_1)] r\hat{p}_{1t}$ and $\pi' = \pi/(1-\phi)$. Thus, if $r\hat{p}d_{it}$ is used instead of rpd_{it} in (18), its coefficient would be biased upward.

Note that this problem need not introduce a systematic bias in (19). For example, if we also have $l\hat{p}_{it}^{T} = \phi lp_{it}^{N} + (1-\phi) lp_{it}^{T}$, then $l\hat{p}_{it} = l\hat{p}_{it}^{T} - lp_{it}^{N} = (1-\phi) lp_{it}$. Using this relation and the corresponding one for country 1, we can show that the use of measured values in (19) would not bias the estimate of the effect of labor productivity differential.

Measured Employment Not Adjusted for Labor Quality

Express the amount of effective labor in sector $Z = T, N$, as $L_{it}^{Z} = E_{it}^{Z} \hat{L}_{it}^{Z}$, where $\hat{L}_{it}^{Z}$ is the actual (measured) quantity of labor and E_{it}^{Z} is the average quality or efficiency of labor. The measured labor productivity is related to the true productivity (in logs) as $l\hat{p}_{it}^{Z} = y_{it}^{Z} - \hat{l}_{it}^{Z} = lp_{it}^{Z} + e_{it}^{Z}$. Suppose that efficiency is positively correlated with true labor productivity. Assume that this relation takes the simple form: $e_{it}^{Z} = \rho lp_{it}^{Z}$, $\rho > 0$. Recalling that $lp_{it} = lp_{it}^{T} - lp_{it}^{N}$, it follows that $lp_{it} = l\hat{p}_{it}/(1+\rho)$. Let $lp_{1t} = l\hat{p}_{1t}/(1+\rho_1)$, $\rho_1 \geq 0$, be the corresponding relation for country 1. Then using these relations, and letting $l\hat{p}d_{it} = l\hat{p}_{it} - l\hat{p}_{1t}$, we can express (19) as:

$$rpd_{it} = \psi_i + \chi_t' + \lambda' l\hat{p}d_{it} + v_{it},$$

where $\chi_t' = \chi_t + \lambda[1/(1+\rho) - 1/(1+\rho_1)] l\hat{p}_{1t}$ and $\lambda' = \lambda/(1+\rho)$. Thus, the use of $l\hat{p}d_{it}$ instead of lpd_{it} in (19) would bias the effect of the productivity variable downward.

DATA APPENDIX

The data set consists of a number of annual time series for 16 developing countries and the United States. All series cover the 1976–94 time period. The selection of developing countries and the choice of the time period are dictated by the availability of the data.

Definitions and Data Sources

The U.S.-dollar exchange rate (*S*) and the consumer price index (*P*) are from IMF, *International Financial Statistics* (IFS). The export and import price indexes (P^X, P^M) represent the price/unit-value series from IFS, or if IFS data are not available, export and import price deflators from IMF, *World Economic Outlook* database. These indexes are used to calculate the terms of trade. The terms of trade data are not available for Singapore for the years 1976–78 and for Turkey for the years 1985–88.

Measures of the labor-productivity differential and relative price of nontraded goods are based on sectoral data on output, employment, and prices. Traded goods are represented by manufacturing and agriculture sectors, and nontraded goods by all other sectors. Value added in constant local currency units is used to measure outputs of traded and nontraded goods sectors (Y^T, Y^N). Labor inputs in the two sectors (L^T, L^N) represent the number of persons employed in each sector. Price indexes for traded and nontraded goods (P^T, P^N) are price deflators derived from value-added data in current and constant local currency units. For the United States, all of these series are from OECD, *Structural Analysis* (STAN) database. For developing countries, the series, Y^T, Y^N, P^T, and P^N, are from World Bank, *World Development Indicators* (WDI). The price deflator for services etc., which accounts for the bulk of the nontraded goods sector, is used to estimate P^N. The data on total employment in manufacturing are from World Bank, *Trade and Production* database.[24] A short gap in this data for Cameroon was filled by linear interpolation. Employment in agriculture is derived from value added per worker and total value-added series given in WDI. L^T is defined as the sum of employment in manufacturing and agriculture obtained from the above sources. L^N is measured residually as the difference between total labor force (also from WDI) and L^T. A limitation of the employment data is that employment in agriculture, manufacturing, and other (nontraded) sectors is not measured on a consistent basis. Labor productivity measures for traded and nontraded goods sectors equal Y^T / L^T and Y^N / L^N, respectively.

Income Groups

The 16 developing countries were divided in low- and high-income groups according to average GDP per capita (from WDI) for the sample period. Low- (high-) income group represents countries with per-capita income smaller (greater) than 2, 000 in 1995 U.S. dollars. The countries in each group are listed below.

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[24] See Nicita and Olarreaga (2001) for description of this database.

Low-Income Group	High-Income Group
Cameroon	Chile
Colombia	Republic of Korea
Ecuador	Malaysia
India	Mexico
Jordan	Singapore
Kenya	South Africa
Morocco	Turkey
Philippines	Venezuela

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