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The Relative Price of Non-Traded Goods under Imperfect Competition*

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Abstract

We consider the role of imperfect competition in explaining the relative price of non-traded to traded goods within the Balassa-Samuelson framework. Under imperfect competition in these two sectors, relative prices depend on both productivity and mark-up differentials. We test this hypothesis using a panel of sectors for 12 OECD countries. The empirical evidence suggests that relative price movements are well explained by productivity and mark-up differentials.

JEL Classification: F31, F36, C23.

Keywords: Balassa-Samuelson hypothesis, real exchange rates, relative prices, imperfect competition.

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I. Introduction

The relative price of non-traded and traded goods is important in explaining real exchange rate movements and price convergence between countries. According to Balassa (1964) and Samuelson (1964), the relative price of non-traded goods is explained under perfect competition by differences in productivity between sectors, rather than by demand factors such as changes in fiscal policy.

The empirical literature—Bergstand (1991), Canzoneri *et al.* (1999), De Gregorio *et al.* (1994), DeLoach (2001), Froot and Rogoff (1995) and Kakkar (2003)—corroborates the fact that productivity changes in the non-tradeable and tradeable sectors are correlated with relative price changes. However, the empirical evidence also indicates that variations in aggregate demand—such as changes in public expenditure—are an important determinant of relative price variation, a fact which cannot be explained by the Balassa-Samuelson framework. Demand factors are also relevant in explaining the existence of inflation differentials in the European monetary union. Inflation in the traded sector (manufacturing) has tended to converge as a consequence of the introduction of the euro and the single market, but inflation in the non-traded sector (services) has tended to be different between countries (see European Central Bank, 1999). The Balassa-Samuelson theory suggests that these inflation variations are explained by a productivity gap between the traded and non-traded sectors (supply-side factors), with demand-side factors such as changes in fiscal policy, business cycles, etc, playing no role. However, different mark-up behaviour in services and manufacturing could be another important determinant of inflation differentials.

We examine the role played by market power in determining relative prices in the tradeable and non-tradeable sectors. Unlike in the Balassa-Samuelson framework, in an economy with imperfect competition, prices are determined by both marginal costs and mark-ups. Mark-up variations potentially amplify or dampen the price repercussions of variations in productivity. Mark-up fluctuations also provide a channel through which variations in aggregate demand could affect the relative price of non-traded goods.

We evaluate the empirical relevance of imperfect competition in explaining relative price movements using panel data for 12 OECD economies. Corroborating the previous empirical literature, we find evidence of a Balassa-Samuelson effect: an increase in the ratio between traded productivity and non-traded productivity increases the relative price of non-traded goods. Our results also show that relative prices and relative mark-ups in the non-traded and traded sectors are correlated: an increase in the non-traded mark-up relative to the traded mark-up raises the relative price of non-traded goods.

The rest of the paper is laid out as follows. In Section II, we consider imperfect competition in the Balassa-Samuelson framework and discuss the effects of variation in productivity and mark-ups on the relative prices of non-traded goods. In Sections III and IV we describe the data and our estimation procedures. In Section V we describe the empirical methodology and report our regression results for relative prices. Finally, Section VI contains our conclusions.

II. Relative prices, productivity and mark-ups

As in Balassa (1964) and Samuelson (1964), we consider an open economy that produces traded and non-traded goods, denoted by the superscripts *T* and *N*, respectively. We use lower-case letters to denote variables in logs. The log of the real exchange rate, *q*, is defined as:

$$q = s + p^* - p \,, \tag{1}$$

where s is the log of the nominal exchange rate, p is the log of the price index and * refers to foreign variables. Since the price index is a weighted average of traded and non-traded prices, $p = (1-\phi)p^T + \phi p^N$, movements in the real exchange rate can be decomposed into two components, namely, deviation from the law of one price in the traded sector and variation in the relative price of non-traded goods (assuming the same share of traded/non-traded goods):

$$q = s + p^{T*} - p^T + \phi(p^{N*} - p^{T*}) - \phi(p^N - p^T) \,. \tag{2}$$

The relevance of the relative price of non-traded goods in explaining the real exchange rate is an empirical issue. Betts and Kehoe (2006) found that movements in the relative price of non-traded goods were closely associated with movements in the real exchange rate in the USA and, furthermore, that this relationship was stronger in direct proportion to the intensity of trade between the USA and its trading partners. In contrast, Engel (1999) found that changes in the relative price of non-traded goods accounted for a small proportion of total real exchange rate changes in the USA.

As in Bergstrand (1991), De Gregorio, Giovannini and Wolf (1994), DeLoach (2001) and Kakkar (2003), we analyse the relative price of non-traded goods within the basic Balassa-Samuelson framework, but we introduce imperfect competition; thus, firms have market power to fix prices over marginal cost in both the traded and non-traded sectors. The existence of imperfect competition in the international market can be considered in different ways; for instance, in monopolistic competition—as in the New Keynesian model (see Obstfeld and Rogoff, 1996)—market power is the consequence of product differentiation. Since our results do not depend on the reasons underpinning market power, we present them in a general setup.

As in the original Balassa-Samuelson model, traded and non-traded goods are produced through a constant returns-to-scale production function:

$$Y^i = A^i L^i f^i(\kappa^i), \quad i = T, N . \tag{3}$$

The term A represents total factor productivity, κ is the capital-labour ratio, L is labour and f is the per-worker production function. Inputs can move freely across sectors, hence firms across sectors pay the same wage, w. The real interest rate, r, is determined in the international capital market, given that the economy is *small* in terms of the capital market and there is international capital mobility. However, we differ from the basic Balassa-Samuelson conditions in that firms have market power to fix their prices. Firms set their prices over marginal cost, $C(w,r)$, as:

$$P^i = \mu^i C(w,r) \quad i = T, N . \tag{4}$$

In the case of imperfect competition, $\mu^i > 1$. The mark-up μ^i is potentially affected by a range of different factors, including changes in market concentration, regulation, etc.[1] Here, the key point is that firms in the non-traded sector meet demand mainly from the domestic market, whereas firms in the traded sector meet demand from both the domestic market and abroad.[2]

For given mark-ups, we derive factor market equilibrium in the economy from equation (4). Using cost minimization, the marginal cost is represented as a function of input costs and the marginal productivity of capital and labour (the price of the traded sector is normalized to one, and hence $P = P^N$):

$$PA^N f_\kappa^N(\kappa^N) = \mu^N r \ , \tag{5}$$

$$PA^N(f^N(\kappa^N) - \kappa^N f_\kappa^N(\kappa^N)) = \mu^N w \, , \tag{6}$$

$$A^T f_\kappa^T(\kappa^T) = \mu^T r \, , \tag{7}$$

$$A^T(f^T(\kappa^T) - \kappa^T f_\kappa^T(\kappa^T)) = \mu^T w \, . \tag{8}$$

We write the marginal productivity of capital and labour in terms of the capital-labour ratio, κ^i, and the derivative of the per-worker production function, $f_\kappa^i(\kappa^i)$.

Equilibrium in the factor market is represented by the set of equations (5) to (8), where the four equilibrium variables are *P, w,* κ^T and κ^N, with the real interest rate determined by the international capital market. Given the productivity and mark-up in each sector, we use the above set of equations to determine the relative price of non-traded goods. From equation (7) we obtain the capital-labour ratio in the traded sector, κ^T, as a function of the international interest rate and the mark-up in this sector. We then compute the wage w as a function of the international interest rate and mark-up in the traded sector by substituting κ^T in equation (8) and we then solve for κ^N and P from equations (5) and (6). We differentiate the above equilibrium conditions in order to compute the effect of a

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[1] For example, in the case of the Dixit-Stiglitz model, the elasticity of demand determines the mark-up, which is constant.

[2] Thus, the non-tradeable sector is sheltered from international competition, whereas the tradeable sector is exposed to international competition. However, modelling the reasons why one sector is tradeable or non-tradeable is not relevant for our empirical analysis.

variation in relative mark-up and productivity on the relative price of the non-traded good as:[3]

$$\Delta p = \left(\frac{\alpha_N}{\alpha_T} \Delta a^T - \Delta a^N \right) - \left(\frac{\alpha_N}{\alpha_T} \Delta \mu^T - \Delta \mu^N \right), \tag{9}$$

where $\Delta p = \Delta p^N - \Delta p^T$ and Δx denotes the rate of growth of variable x, approximated for a time index t+1 as $\Delta x \approx log\, X_{t+1} - log\, X_t$—for instance $\Delta \mu_t^T = log(\mu_{t+1}^T) - log(\mu_t^T)$—and where $\alpha = \frac{F_L L}{F}$ denotes the output-labour elasticity in each sector.

Note that for the case of perfect competition ($\mu_t^N = \mu_t^T = 1$), the result in equation (9) is the original Balassa-Samuelson hypothesis, according to which variations in the relative price of non-traded goods should only be explained by variations in total factor productivity. Under imperfect competition, mark-up variations produce changes in prices provided that mark-up movements in one sector are not offset by mark-up movements in the other sector. As for productivity, the effect of an increase in the mark-up for traded goods depends on the ratio of output elasticities, because a mark-up increase in a labour-intensive sector reduces wages.

A satisfactory theory to explain the evolution of the relative price of non-traded goods cannot overlook the effect of mark-up variations on prices. It also has to distinguish between the effect of a variation in productivity compared with a variation in the mark-up on prices. Our objective was to estimate equation (9) so that we could distinguish between the effect on relative prices of productivity and mark-up variations. Changes in mark-ups only cause relative prices to vary when these follow different paths in each sector.

Different authors—for instance, Rotemberg and Woodford (1999) and Bils and Chang (2000)—argue that mark-up movements are crucial to understanding price responses to changes in marginal cost, given that countercyclical mark-up movements offset the effect on prices of pro-cyclical marginal cost movements.

[3] The derivation of this equation is similar to the perfect competition case presented by Obstfeld and Rogoff (1996), Ch. 4.

Therefore, mark-up changes will transmit the shock to the relative price of non-traded goods.

Finally, our model offers an alternative explanation for the observed positive relationship between increased public spending and non-traded sector prices (see, for instance, De Gregorio *et al.*, 1994; Strauss, 1999). Variations in mark-ups arising from fiscal expansion could affect the relative price of non-traded goods.

III. Data

We obtained sectoral data for a set of countries from the OECD STAN Database for Industrial Analysis.[4] To distinguish between traded and non-traded goods, we followed the classification proposed by De Gregorio *et al.* (1994), who grouped 18 sectors into traded and non-traded goods categories according to the ratio of exports to total production. Authors such as Canzoneri *et al.* (1999), Bettendorf and Dewachter (2007) and Méjean (2008) have also adopted this classification. The twelve traded goods sectors are agriculture, hunting, forestry and fishing; mining and quarrying; food products, beverages and tobacco; textiles, textile products, leather and footwear; wood and products of wood and cork; pulp, paper, paper products, printing and publishing; chemical, rubber, plastics and fuel products; other non-metallic mineral products; basic metals and fabricated metal products; machinery and equipment; transport equipment; and, finally, manufacturing n.e.c. and recycling. The six non-traded goods sectors are electricity, gas and water supply; construction; wholesale, retail trade, restaurants and hotels; transport, storage and communications; finance, insurance, real estate and business services; and community, social and personal services. We denote the sectors by means of an index j, with j=1,…, 12 for traded sectors and with j=1, …., 6 for non-traded sectors.

For each of the sectors we collected annual data on value added at current and constant prices, total employment and number of employees, gross capital stock at constant prices and labour cost (compensation of employees). Missing data for

[4] The database can be downloaded from http://www.oecd.org/sti/stan.

some of the variables restricted the analysis to just 12 countries in the STAN Database[5] for different time periods, resulting in an unbalanced panel. For the selected countries, Table 1 summarizes the annual periods included in the sample, the average share of non-traded goods in value added and the capital-labour ratio for the non-traded goods divided by the capital-labour ratio for traded goods. Non-traded goods represented a substantial share of total value added, ranging from 64% in Japan to 85% in France. There was a wide range of values for the relative capital-labour ratio—lowest in Japan and the UK and highest in Germany and Denmark for non-traded goods. The fact that traded and non-traded goods in different countries have different capital-labour ratios suggests that there are technological differences between countries; it therefore made sense to estimate a different output-labour elasticity for each country. Moreover, it is not always the case that non-traded goods are labour intensive; the electricity, gas and water supply sector, for example, is very capital intensive.

INSERT TABLE 1 HERE

IV. Estimation procedure and analysis of the basic variables

Below we explain how we computed the key variables in equation (9) and discuss their time-series properties.

Relative price changes

Value-added deflators measure prices for each sector, whereas changes-of-log deflators define price changes for a sector. We computed aggregate price changes in non-traded and traded goods as the weighted-average price changes for the corresponding sectors using the shares of value added, s^i_{jt}, as the weights:

$$s^i_{jt} = \frac{Y^i_{jt}}{\sum_{j=1}^{h} Y^i_{jt}}, \qquad (10)$$

[5] For Japan we took data from the OECD International Sectoral Database, since data for this country are not included in more recent versions of STAN.

where Y^{i}_{jt} denotes the nominal value added of sector j at time t and where h is the number of sectors in the non-traded or traded goods category, with h=6 and h=12, respectively. Thus, the aggregate change in prices at time t was defined as:

$$\Delta p^{N}_{t} = \sum_{j=1}^{6} s^{N}_{jt} \Delta p^{N}_{jt} \ , \qquad (11)$$

$$\Delta p^{T}_{t} = \sum_{j=1}^{12} s^{T}_{jt} \Delta p^{T}_{jt} \ . \qquad (12)$$

The change in the relative price of non-traded goods at time t for each country, denoted by the sub-index k, is defined as:

$$\Delta p_{t} = \Delta p^{N}_{t} - \Delta p^{T}_{t} \equiv p_{kt} \ . \qquad (13)$$

To simplify the notation, in the empirical model we redefine Δp_{t} for country k as p_{kt} .

Table 2 shows summary statistics for changes in non-tradeable and tradeable prices and changes in relative prices for each country in our database. The average growth in non-tradeable prices across all the countries, with the exception of Belgium, Germany and Japan, was around 5% per annum. The behaviour of tradeable prices was a little more dispersed. The average change in relative prices for non-traded goods was around 2% per annum for all the countries, with the exception of Canada and Norway (where tradeable prices grew faster than non-tradeable prices) and also Germany (where non-tradeable and tradeable prices movements were similar).

INSERT TABLE 2 HERE

Sectoral productivity changes

Productivity changes for sector j were obtained from the production function as:

$$\Delta a_{jt} = (\Delta y_{jt} - \Delta l_{jt}) - (1 - \alpha_{j})(\Delta k_{jt} - \Delta l_{jt}) \ , \qquad (14)$$

where y_{jt} , l_{jt} and k_{jt} are the logs of real output, total employment and the capital stock, respectively. Estimates of productivity changes from equation (14) required a value for output-labour elasticity (α_{j}), estimated from the production function. We chose this approach to measuring productivity rather than the Solow residual approach, because, as demonstrated by Hall (1988), the latter is not an appropriate measure of productivity under imperfect competition. To illustrate this point, consider the Solow residual, $SR = LnY - s_L Ln\left(K/L\right)$, where s_L denotes the labour share. From equation (6) we have $\alpha = \mu s_L$, since wages are not equal to the marginal productivity of labour. Hence, the Solow residual is only an accurate productivity measure when there is perfect competition, $\mu = 1$; otherwise, the Solow residual is $SR = Ln(A) + (\mu - 1) s_L Ln(K/L)$.[6]

Accordingly, we estimated output-labour elasticity by considering a Cobb-Douglas production function, in which the output for sector j in each country is given by:

$$\begin{aligned} y_{jt} &= \alpha_j l_{jt} + \beta_j k_{jt} + \xi_t + \left(a_{jt} + \eta_j + \varepsilon_{jt}\right) \\ a_{jt} &= a_j + \rho a_{jt-1} + \omega_{jt} \qquad \left|\rho\right| < 1 \end{aligned}, \tag{15}$$

where η_j is an unobserved time-invariant sector-specific effect, ξ_t is a year-specific intercept, ε_{jt} reflects serially uncorrelated measurement errors and ω_{jt} is a productivity shock. Like Blundell and Bond (2000), we maintain that employment and capital are both potentially correlated with sector-specific effects and also with productivity shocks and measurement errors. The corresponding dynamic representation of equation (15) is as follows:

$$y_{jt} = \rho y_{jt-1} + \alpha_j l_{jt} - \rho \alpha_j l_{jt-1} + \beta_j k_{jt} - \rho \beta_j k_{jt-1} + \xi_t^* + (\eta_j (1-\rho) + a_j) + (\omega_{jt} + \varepsilon_{jt} - \rho \varepsilon_{jt-1}), \tag{16}$$

where $\xi_t^* = \xi_t - \rho \xi_{t-1}$. This production function was estimated for tradeables and non-tradeables for each country and for the sectors included in the two categories. We assumed that $\alpha_j = \alpha_N$ if j was a non-tradeable sector and that $\alpha_j = \alpha_T$ if j was a tradeable sector, as the theoretical model assumes that output-labour elasticity is different for tradeables and non-tradeables but is the same for sectors in the same category. In addition, as in the original Balassa-Samuelson framework and our

[6] Under imperfect competition the Solow residual is affected by variations in output generated by changes in demand. It is therefore not suitable for testing the Balassa-Samuelson effect, as it is unable to distinguish between the effects of variations in demand or productivity on prices.

theoretical framework, a constant returns-to-scale technology was assumed; we therefore considered equation (16) under the restriction that $\alpha_j + \beta_j = 1$. Finally, we estimated output-labour elasticity for the tradeable and non-tradeable sectors using the system generalized method of moments (GMM) estimator proposed by Arellano and Bover (1995) and Blundell and Bond (1998), as it has been found to reduce the finite sample bias of the first differences GMM estimator in the estimation of a Cobb-Douglas production function (Blundell and Bond, 2000).

The output-labour elasticities from equation (16), reported in Table 3, are one-step estimates with robust standard errors, obtained using the same moment conditions as in Blundell and Bond (2000).[7] The estimated elasticities were statistically significant, indicating that the non-traded output-labour elasticity was higher than the traded output-labour elasticity for half of the countries. Using the estimated output-labour elasticity, we compute changes in productivity from equation (14) and aggregate those changes at t using the industry nominal value added share, s_{jt}. This aggregation is consistent with the Domar (1962) approach widely used in the literature. Thus:

$$\Delta a_t^N = \sum_{j=1}^{6} s_{jt}^N \Delta a_{jt}^N \ , \tag{17}$$

$$\Delta a_t^T = \sum_{j=1}^{12} s_{jt}^T \Delta a_{jt}^T \ . \tag{18}$$

Accordingly, for each country k, the relative change in productivity at time t, included in equation (9), is defined as:

$$a_{kt} = \frac{\hat{\alpha}_N}{\hat{\alpha}_T} \Delta a_t^T - \Delta a_t^N \ . \tag{19}$$

[7] The results are computed using the DPD for Ox package (Doornik, Arellano and Bond, 2006). As instrumental variables, we used lagged values for the first differences of the explanatory variables from lag t-2 to t-3 for the equation in levels, and lagged values for the explanatory variables in levels from lag t-3 to t-4 for the transformed equation. For the sake of brevity and given our interest only in the parameter of the production function that we need to obtain—equation (14)—we do not provide a detailed explanation of estimation of the production function. Further details are available on request.

As can be observed in columns 3-5 in Table 3, the average productivity growth for non-traded goods differed substantially across countries and was even negative in some countries, whereas the productivity growth for tradeables was positive in all the countries. In addition, the average changes in relative productivity confirm that productivity for tradeables grew faster than for non-tradeables, except in Japan, where non-tradeable productivity grew at a much faster rate than in any other country. Looking at the differences in productivity growth between sectors, except for Japan, the average was around 2%.

INSERT TABLE 3 HERE

Relative mark-up changes

In order to compute changes in mark-ups, we use the market equilibrium conditions—equations (6) and (8)—for a Cobb-Douglas production function which, for sector *j*, requires that:[8]

$$P_{jt}\alpha_j\left(\frac{Y_{jt}}{L_{jt}}\right)=\mu_{jt}w_{jt}\,, \tag{20}$$

where P_{jt}, μ_{jt} and w_{jt} are, respectively, the price, mark-up and wage level for sector *j* at time *t*. From this equilibrium condition, we can calculate the mark-up for sector *j* as a function of the output-labour elasticity and labour share (S_{Ljt}):

$$\alpha_j=\mu_{jt}\frac{w_{jt}L_{jt}}{P_{jt}Y_{jt}}=\mu_{jt}S_{Ljt}\,. \tag{21}$$

Therefore, mark-up changes can easily be computed as:

$$\Delta\mu_{jt}=-\Delta s_{Ljt}\,. \tag{22}$$

Note that equation (22) does not include estimates of output-labour elasticity and the value-added deflator.[9]

[8] Note that the other equilibrium equation states that the value of the marginal product of capital equals the mark-up multiplied by the cost of capital. We did not use this second condition because the estimate of the cost of capital for each sector is more inaccurate than the wage estimates given by the database.

In order to compute the labour share, we included self-employment earnings as labour income, as in Gollin (2002). We first obtained the average wage from the database as:

$$w_{jt} = \frac{\text{Compensation of employees}}{\text{Number of employees}}$$

and then multiplied this average salary by total employment, L_{jt}, which included both employees and self-employed workers. Finally, to obtain the labour share we divided the labour income in nominal terms by value added at current prices.

Once changes in mark-ups were obtained for all the sectors, we aggregated them at t using the industry value added share, s_{jt}. Thus:

$$\Delta\mu_t^N = \sum_{j=1}^{6} s_{jt}^N \Delta\mu_{jt}^N \ , \tag{23}$$

$$\Delta\mu_t^T = \sum_{j=1}^{12} s_{jt}^T \Delta\mu_{jt}^T \ . \tag{24}$$

Accordingly, for each country k, the relative change in the mark-up at time t is defined as:

$$m_{kt} = \frac{\hat{\alpha}_N}{\hat{\alpha}_T} \Delta\mu_t^T - \Delta\mu_t^N \ . \tag{25}$$

Table 4 provides summary statistics for mark-up changes in non-tradeables and tradeables and relative mark-up changes for each country in our database. The average change in tradeable mark-ups was positive (except for Japan and Spain), whereas the average change in non-tradeable mark-ups was more dispersed and even negative for 7 of the 12 countries. Additionally, the average changes in mark-ups were generally greater in the tradeable sector, except for Italy, Japan and Spain. Looking at the relative change in mark-ups, the average was quite dispersed across countries and was positive in 8 of the 12 countries.

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[9] It would also be possible to calculate the level of the mark-up from equation (21), although this could be unreliable as it might be affected by labour share measurement errors or by bias in the estimate of output-labour elasticity. Note that as long as these errors are constant over time, they are reduced by first differencing in equation (22).

INSERT TABLE 4 HERE

Non-stationarity

Using panel unit root tests, we assessed the non-stationarity properties for our three variables of interest, namely p_{kt} , a_{kt} and m_{kt}. There are several panel unit root tests available, differing in whether the null is a unit root or stationarity, whether serial correlation is removed parametrically or non-parametrically and whether the design is for cross-sectionally independent panels or for cross-sectionally correlated panels.[10] The panel unit root tests we implemented were the pooled augmented Levin, Lin and Chu (2002) and Breitung (2000) Dickey-Fuller tests (augmented Dickey-Fuller, ADF). Both test the null hypothesis of a unit root, $\beta = 0$, in the basic ADF specification:

$$\Delta y_{kt} = \beta y_{kt-1} + \sum_{j=1}^{p} \beta_{kj} \Delta y_{kt-j} + \varepsilon_{kt} , \tag{26}$$

under the assumption that β is common across a cross-sectionally independent distributed panel, with both tests taking different variable transformations. Im *et al.* (2003) proposed specifying a separate ADF regression for each cross-section and testing whether $\beta_k = 0$ for all *k*. Also, Maddala and Wu (1999) proposed a Fisher-type test that assumes heterogeneity. In considering heterogeneity and stationarity under the null, we employed the test proposed by Hadri (2000), which is a panel extension of the stationarity test described in Kwiatkowski *et al.* (1992). Finally, we considered the unit root test proposed by Pesaran (2007), which extends the Im *et al.* (2003) ADF-type regression by including cross-section averages of lagged levels and first differences for the individual series.

The results of the panel non-stationarity and stationarity tests for our three variables of interest are summarized in Table 5. Panel unit root cross-sectionally independent tests were unanimous in rejecting the presence of a unit root in relative price, productivity and mark-ups. This conclusion did not change on examining the

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[10] An exhaustive description of these tests and their properties can be found in a recent article by Breitung and Pesaran (2008).

results of the Pesaran (2007) test, which accounts for cross-section dependence. For the Hadri (2000) stationarity test, the null of stationarity was not rejected for relative changes in productivity and mark-ups. To sum up, the three series appear to be stationary according to each panel unit root and stationary test performed.

INSERT TABLE 5 HERE

V. Econometric methods and results

This section provides empirical support for the equilibrium relationship given by equation (9). Using the notation introduced in the previous section, that is, p_{kt} for the rate of growth in the relative price of non-tradeable goods, a_{kt} for changes in productivity and m_{kt} for changes in mark-ups, we estimated equation (9) by considering the following panel regression model:

$$p_{kt} = \gamma_k + \beta_k a_{kt} + \phi_k m_{kt} + \varepsilon_{kt} \tag{27}$$

for k=1,…,12 countries and a total of 304 observations for different time periods between 1970 and 2006 (see Table 1). γ_k is a country-specific factor and ε_{kt} *is i.i.d.*$(0,\sigma^2)$, capturing stochastic deviations from the equilibrium relationship given by equation (9). The coefficients β_k and ϕ_k measure the impact of relative productivity and mark-ups, respectively, on relative prices for country k at time t. Our theory in equation (9) states that those coefficients should have values of 1 and -1, respectively. We estimated equation (27) under the following parameter restrictions: (i) assuming that $\gamma_k = \gamma$, $\beta_k = \beta$, $\phi_k = \phi \quad \forall k$ and assuming that ε_{kt} *is i.i.d.*$(0,\sigma_k^2)$, in which case the estimation is called generalized ordinary least squares for the pooled data (GPOLS); and (ii) controlling for country heterogeneity by assuming that γ_k could be different for each country, $\beta_k = \beta$ and $\phi_k = \phi \quad \forall k, p$, and assuming that ε_{kt} *is i.i.d.*$(0,\sigma_k^2)$, with this estimation denominated the generalized fixed-effect estimator (GFE).[11] Detailed explanations of these models can be found in Baltagi (2008).

[11] Other specifications for the parameter restrictions and the covariance matrix are possible. For example, assuming that $E(\varepsilon_{kt}\varepsilon_{jt})=\sigma_{kj} \ \forall k, j, t$ and otherwise zero, or $E(\varepsilon_{kt}\varepsilon_{jt})=\sigma_t^2 \ \forall k, j, t$

Table 6 depicts a simple correlation analysis between these variables. Note that the correlation between p and m was strongly negative for all the countries except Japan, where the correlation was weekly negative; the correlation between p and a was generally positive, although negative for some countries. Also, relative mark-ups accounted for around 40% (country average) of the volatility in relative prices, whereas relative productivity accounted for around 27%.

INSERT TABLE 6 HERE

Table 7 reports the estimates for equation (27) under two parameter specifications. Empirical evidence supported the theoretical hypotheses that productivity had a significant positive effect and mark-up differentials had a significant negative effect on the relative price of non-traded goods. The coefficient for productivity differentials—at around 0.75—was above the values obtained by De Gregorio *et al.* (1994), who reported an average coefficient estimate of 0.23. Likewise, the effect of mark-up changes on relative price changes—at around -0.85—remained robust when we excluded the effect of productivity and the effect of the intercept on the GPOLS specification. We analysed the robustness of our results for a possible endogeneity problem caused by variable measurement errors or the omission of relevant variables. We did this by making a GMM estimation using lagged values from t-1 to t-2 for both explanatory variables as instrumental variables. It can be observed in the last two rows of Table 7 that empirical results point to the same conclusions, thus confirming that potential endogeneity does not bias our results. Standard errors reported in Table 7 indicate that the estimated coefficients are different from their theoretical values, even though they are properly signed.

INSERT TABLE 7 HERE

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and otherwise zero, or $E(\varepsilon_{kt}\varepsilon_{ks})=\sigma_{ts}\ \forall k, t, s$ and otherwise zero, we can allow for heteroskedasticity and contemporaneous correlation, as in a seemingly unrelated regression model, period heteroskedasticity and period heteroskedasticity and serial correlation, respectively. Empirical results for those specifications, not reported but available on request, lead to the same conclusions as those presented here.

Pooled mean group estimation

Given that previous empirical models impose homogeneity in the slope coefficients across countries ($\beta_k = \beta$ and $\phi_k = \phi$), we also considered the pooled mean group (PMG) estimator proposed by Pesaran *et al.* (1999), which constrains the long-run coefficients to be the same, while allowing the intercepts, short-run coefficients and error variances to differ freely across countries. The PMG procedure is attractive, as equation (9) suggests long-run homogeneity.

We assume that the long-run relative price function is given by equation (27) and consider the following autoregressive distributed lag (ARDL) (1,1,1) model:

$$p_{kt} = \theta_k + \beta_{1k} a_{kt} + \beta_{2k} a_{kt-1} + \phi_{1k} m_{kt} + \phi_{2k} m_{kt-1} + \lambda_k p_{kt-1} + u_{kt} \, . \tag{28}$$

The error correction equation is therefore:

$$\Delta p_{kt} = \delta_k (p_{kt-1} - \gamma_k - \beta_k a_{kt} - \phi_k m_{kt}) - \beta_{1k} \Delta a_{kt} - \phi_{1k} \Delta m_{kt} + u_{kt} \, , \tag{29}$$

where $\delta_k = -(1-\lambda_k)$ is the speed of adjustment coefficient, $\gamma_k = \theta_k / (1-\lambda_k)$, $\beta_k = (\beta_{1k} + \beta_{2k}) / (1-\lambda_k)$ and $\phi_k = (\phi_{1k} + \phi_{2k}) / (1-\lambda_k)$. The PMG estimate is based on equation (29), under the restriction that all long-run coefficients are equal across countries, $\beta_k = \beta$ and $\phi_k = \phi$, allowing thus for unrestricted country heterogeneity in the adjustment dynamics. The disturbances u_{kt} have zero mean and variance σ_k^2. For the purpose of the robustness check, we also provide two alternative pooled estimates: a mean group (MG) estimator and a dynamic fixed-effect (DFE) estimator. The MG estimator (Pesaran and Smith, 1995) provides an estimate of the mean long-run effect across countries, thus allowing countries to differ in terms of long-run effects. The DFE constrains all the slope coefficients and error variances to be the same. The null hypothesis of long-run homogeneity was tested using the Hausman test for equivalence between the PMG and MG estimators.

Table 8 shows estimates from the MG, PMG and DFE estimators for the ARDL(1,1,1) specification. Parameter estimates did not change very much through the different pooled methods and so were quite close to the estimates given in Table 7. In the country-specific regressions, the long-run estimates of the effects of productivity on prices were between 0.20 and 1.36, with an average estimate of 0.75, whereas the long-run estimates of the effects of mark-ups on prices ranged from -

1.49 to -0.3, with an average estimate of -0.81. Moreover, all estimates of the adjustment coefficient were negative and fell in the range -1.14 to -0.83, for an average of -0.78. Standard errors indicate that the MG parameter estimates were consistent with our theoretical parameter values in equation (9), since the null hypothesis that the coefficients for productivity and mark-ups should be 1 and -1, respectively, was not rejected. Looking at specific country estimates, the theoretical hypothesis of a unit coefficient for the relative productivity variable was rejected for just four countries, whereas for the relative mark-up variable, the hypothesis was rejected for just three countries. Imposing long-run homogeneity (that is, using PMG instead of MG estimators) resulted in more precise estimates of the long-run effect. The Hausman test statistic accepted the null of no difference between the MG and PMG estimators. As in Pesaran et al. (1999), likelihood ratio tests at conventional significance levels would reject all the restrictions, including homogeneity in the long-run coefficients. Standard errors indicate that the null hypothesis that the coefficient for productivity is 1 was rejected, whereas we cannot reject the hypothesis that the coefficient for mark-ups should be -1. On the other hand, the DFE long-run estimated parameters were lower than for the PMG estimators and had lower standard errors than the MG estimators. Pooling thus sharpened the estimates without significantly changing their values. Furthermore, we reject the hypothesis that the productivity coefficient is equal to 1 as in the case of the PMG estimators, whereas the hypothesis that the coefficient for mark-ups should be -1 is not rejected if we consider a 99% confidence interval.[12]

INSERT TABLE 8 HERE

[12] As our panel data is unbalanced, representing as it does just a handful of countries (see Table 1), the robustness of the results to variations in country coverage was checked by eliminating a single country or group of countries at a time and re-running the PMG estimation procedure. Point estimates did not differ very much from the ones reported in Table 8.

VI. Conclusions

We introduced imperfect competition in the standard Balassa-Samuelson framework and demonstrated that the relative price of non-traded goods is determined by both productivity and mark-ups.

We also estimated the effects of variation in productivity and mark-ups on relative prices for a panel of sectors belonging to 12 OECD countries, demonstrating that changes in relative productivity and mark-ups have significant and opposite effects on the relative price of non-traded goods. Faster productivity growth in the traded goods sector relative to the non-traded sector increases the relative price of non-traded goods. Our results also support the hypothesis that mark-ups in the traded and non-traded goods sectors follow different paths and generate variations in the relative price of non-traded goods. Thus, variations in mark-ups constitute a new channel through which variations in aggregate demand or other macroeconomic shocks could affect real exchange rates.

These results suggest a number of future lines of research. It could be useful to study the role of mark-ups in the propagation of business cycle fluctuations through variations in relative non-traded good prices, since mark-ups could amplify or reduce the effect of productivity shocks on prices. It would also be interesting to analyse the reasons for variations in mark-ups in each sector, in particular, different fiscal policy measures. Finally, it might also be useful to study how mark-up evolution could account for different inflation rates in the services sector, generating different national inflation rates in the Eurozone.

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TABLE 1

Time period samples for 12 OECD countries

	Coverage	Average Share of Non-Tradeables (%)	Average κ^N / κ^T
Belgium	1995-2006	79.87	1.12
Canada	1970-2001	72.58	0.66
Denmark	1970-2003	76.73	2.08
Finland	1975-2006	68.73	1.80
France	1978-2005	85.63	1.75
Germany	1991-2005	75.18	2.17
Italy	1980-2004	72.61	1.66
Japan	1970-1995	64.15	0.55
Norway	1970-2005	68.91	1.09
Spain	1985-2005	74.26	2.00
UK	1979-2003	72.80	0.59
USA	1977-2006	79.36	1.59

Notes: The average share of non-tradeables refers to the average percentage share in the value added of non-tradeables over the sampled time period. κ^N (κ^T) is the capital-labour ratio for non-tradeables (tradeables). The average κ^N / κ^T is the sample average of κ^N over κ^T .

TABLE 2

Summary statistics for price changes in 12 OECD countries

	# observations	Δp^N	Δp^T	p_k
Belgium	11	0.0201 (0.0074)	0.0035 (0.0126)	0.0167 (0.0115)
Canada	31	0.0502 (0.0354)	0.0551 (0.0602)	-0.0050 (0.0469)
Denmark	33	0.0569 (0.0376)	0.0461 (0.0451)	0.0108 (0.0344)
Finland	31	0.0531 (0.0319)	0.0240 (0.0463)	0.0290 (0.0299)
France	27	0.0418 (0.0317)	0.0242 (0.0433)	0.0176 (0.0184)
Germany	14	0.0144 (0.0177)	0.0064 (0.0151)	0.0080 (0.0171)
Italy	24	0.0671 (0.0469)	0.0460 (0.0391)	0.0212 (0.0188)
Japan	25	0.0287 (0.0220)	0.0160 (0.0320)	0.0127 (0.0155)
Norway	35	0.0522 (0.0300)	0.0667 (0.1088)	-0.0145 (0.1105)
Spain	20	0.0537 (0.0215)	0.0305 (0.0250)	0.0233 (0.0280)
UK	24	0.0556 (0.0379)	0.0349 (0.0427)	0.0207 (0.0325)
USA	29	0.0398 (0.0205)	0.0195 (0.0330)	0.0204 (0.0219)

Notes: The three columns on the right report time means (standard deviations in brackets) for price changes in the countries listed.

TABLE 3

Output-labour elasticity estimates and summary statistics for productivity changes in 12 OECD countries

	α_{N}	α_{T}	Δa^{N}	Δa^{T}	a_k
Belgium	0.72	0.63	0.0035	0.0160	0.0148
	(0.18)	(0.20)	(0.0079)	(0.0208)	(0.0241)
Canada	0.53	0.62	-0.0028	0.0127	0.0137
	(0.06)	(0.01)	(0.0100)	(0.0377)	(0.0289)
Denmark	0.80	0.85	0.0053	0.0292	0.0222
	(0.14)	(0.09)	(0.0160)	(0.0312)	(0.0352)
Finland	0.77	0.60	0.0079	0.0384	0.0414
	(0.06)	(0.26)	(0.0107)	(0.0350)	(0.0408)
France	0.45	0.53	0.0030	0.0184	0.0127
	(0.12)	(0.24)	(0.0066)	(0.0269)	(0.0226)
Germany	0.81	0.71	0.0007	0.0226	0.0250
	(0.21)	(0.16)	(0.0066)	(0.0265)	(0.0323)
Italy	0.40	0.56	-0.0070	0.0083	0.0129
	(0.08)	(0.08)	(0.0088)	(0.039)	(0.0127)
Japan	0.88	0.43	0.0154	-0.0015	-0.0191
	(17.11)	(2.21)	(0.0129)	(0.0300)	(0.0716)
Norway	0.74	0.81	0.0101	0.0279	0.0154
	(0.08)	(0.07)	(0.0109)	(0.0430)	(0.0387)
Spain	0.50	0.67	-0.0094	0.0066	0.0143
	(0.05)	(0.06)	(0.0083)	(0.0175)	(0.0162)
UK	0.51	0.48	-0.0002	0.0176	0.0189
	(0.10)	(0.05)	(0.0125)	(0.0248)	(0.0216)
USA	0.72	0.70	0.0061	0.0281	0.0228
	(0.10)	(0.09)	(0.0100)	(0.0298)	(0.0284)

Notes: The first two data columns show the output-labour elasticity estimates from the production function for non-tradeables and tradeables. Standard errors (in parentheses) were computed using standard errors robust to heteroskedasticity. The last three data columns report time means (standard deviations in brackets) for productivity changes in the countries listed.

TABLE 4

Summary statistics for mark-up changes in 12 OECD countries

	# observations	$\Delta\mu^{\mathrm{N}}$	$\Delta\mu^{\mathrm{T}}$	m_k
Belgium	11	0.0014 (0.0121)	0.0050 (0.0284)	0.0043 (0.0289)
Canada	31	-0.0030 (0.0116)	0.0130 (0.0540)	0.0141 (0.0486)
Denmark	33	-0.0017 (0.0168)	0.0066 (0.0397)	0.0079 (0.0354)
Finland	31	0.0032 (0.0134)	0.0096 (0.0444)	0.0092 (0.0501)
France	27	0.0052 (0.0113)	0.0057 (0.0288)	-0.0003 (0.0285)
Germany	14	-0.0019 (0.0105)	0.0048 (0.0200)	0.0074 (0.0206)
Italy	24	0.0060 (0.0119)	0.0023 (0.0222)	-0.0043 (0.0107)
Japan	25	-0.0008 (0.0146)	-0.0054 (0.0249)	-0.0125 (0.0566)
Norway	35	-0.0040 (0.0180)	0.0222 (0.1262)	0.0243 (0.1125)
Spain	20	0.0005 (0.0109)	-0.0048 (0.0252)	-0.0041 (0.0167)
UK	24	-0.0021 (0.0180)	0.0012 (0.0742)	0.0033 (0.0878)
USA	29	-0.0023 (0.0115)	0.0045 (0.0215)	0.0069 (0.0203)

Notes: The three columns on the right report time means (standard deviations in brackets) for mark-up changes in the countries listed.

TABLE 5

Panel unit root and stationarity test results

	p_{kt}	a_{kt}	m_{kt}
LLC	-7.25*	-7.55*	-8.11*
BRE	-3.70*	-4.23*	-6.45*
IPS	-6.72*	-6.28*	-9.25*
MW	-5.85*	-7.33*	-7.82*
HA	0.07†	0.53**	-0.75†
PE	-7.52*	-8.46*	-9.17*

Notes: Abbreviations as follows: LLC, Levin *et al.* (2002); BRE, Breitung (2000); IPS, Im *et al.* (2003); MW, Maddala and Wu (1999); HA, Hadri (2000); and PE, Pesaran (2007). * indicates rejection of the null unit root hypothesis at the 5% critical level. † indicates no rejection of the null stationarity hypothesis at the 5% level. Lags in equation (26) were selected according to the Schwarz Bayesian criterion and a trend was included in the regression.

TABLE 6

Relative price (p), relative productivity (a) and relative mark-up (m) correlations for 12 OECD countries

	Corr(p, a)	Corr(p, m)	Corr(a, m)
Belgium	-0.18	-0.61	0.84
Canada	0.02	-0.74	0.60
Denmark	0.45	-0.55	0.41
Finland	-0.22	-0.69	0.75
France	0.29	-0.30	0.71
Germany	0.32	-0.22	0.78
Italy	0.43	-0.32	0.41
Japan	0.31	-0.03	0.84
Norway	-0.14	-0.93	0.47
Spain	0.65	-0.46	0.14
UK	0.06	-0.65	0.08
USA	0.54	-0.31	0.53

TABLE 7

Estimates of the effect of changes in productivity differentials and relative mark-ups on changes in relative prices

	γ	β	ϕ	*Adj.* R^2	Sargan test
GPOLS	0.0063 (0.0010)	0.7325 (0.0034)	-0.8791 (0.0295)	0.75	
	0.0176 (0.0014)		-0.4803 (0.0377)	0.34	
GFE		0.7343 (0.0337)	-0.8591 (0.0293)	0.76	
			-0.4672 (0.0365)	0.39	
GMM-I	0.0058 (0.0010)	0.7521 (0.0345)	-0.9078 (0.0301)	0.73	0.14
GMM-II	0.0045 (0.0009)	0.7465 (0.0343)	-0.8810 (0.0297)	0.78	0.12

Notes: GPOLS, generalized pooled ordinary least squares; GFE, generalized fixed effects; GMM-I and GMM-II, generalized method of moments (GMM) estimates for the pooled data and the fixed effects model, respectively, using lagged values from t-1 to t-2 of the two explanatory variables as instrumental variables. Standard errors (in brackets) are robust to white-cross-section heteroskedasticity. The Sargan statistic tests the validity of the over-identifying restriction for GMM estimators (p values are reported).

TABLE 8

Pooled estimates for the relative price equation

	MG	PMG	Hausman test	DFE
β	0.7594*	0.7606*	0.997**	0.6527*
	(0.1301)	(0.0592)		(0.0626)
ϕ	-0.8156*	-1.0801*	0.569†	-0.9144*
	(0.1615)	(0.0357)		(0.0380)
δ	-0.7881*	-0.5805*		-1.0025*
	(0.0687)	(0.1112)		(0.0595)
Countries (n)	12	12		12
Observations (n)	291	291		291

Notes: MG, mean group estimator; PMG, pooled mean group estimator; DFE, dynamic fixed-effect estimator. Figures in brackets are standard errors. * indicates significance at the 1% level. † indicates no rejection of the null of long-run homogeneity.