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The Relative Price of Non-Traded Goods under Imperfect Competition

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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 method

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 $P \quad P^N$):

$$PA^N f^{N}(\;\; ^N) \qquad ^N r \;, \tag{5}$$

$$PA^N (f^N(\;\; ^N) \qquad ^N f^{N}(\;\; ^N)) \qquad ^N w \,, \tag{6}$$

$$A^T f^{T}(\;\; ^T) \qquad ^T r \,, \tag{7}$$

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

$$p \quad \frac{N}{T} \; a^T \quad a^N \quad \frac{N}{T} \quad T \quad N \;, \tag{9}$$

where p p^N p^T and x

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.

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 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%

theoretical framework, a constant returns-to-scale technology was assumed; we therefore considered equation (16) under the restriction that ${}_j \quad {}_j \quad 1$. 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

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

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]

$$P_{jt} \quad {}_{j} \quad \frac{Y_{jt}}{L_{jt}} \qquad {}_{jt} w_{jt} \, , \tag{20}$$

where P_{jt}, ${}_{jt}$ 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}):

$${}_{j} \qquad {}_{jt} \frac{{}_{jt} \quad {}_{jt}}{} \qquad {}_{jt} \quad {}_{Ljt}$$ 0 Tw69c050nan(4 n B0r 186 4 can)5.7(caC.t4.169c[c ET Q

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:

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

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:

$${}^{N}_{t} \qquad {}^{6}_{j\;1} \; s^{N}_{jt} \quad {}^{N}_{jt} , \tag{23}$$

$${}^{T}_{t} \qquad {}^{12}_{j\;1} \; s^{T}_{jt} \quad {}^{T}_{jt} . \tag{24}$$

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

$$\frac{\hat{}_{N}}{{}_{T}} \qquad {}^{T} \qquad {}^{N}$$

INSERT TABLE 4 HERE

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, $\quad 0$, in the basic ADF specification:

$$y_{kt} \qquad y_{kt\ 1} \qquad {}^{p}_{j\ 1} \ {}_{kj} \ y_{kt\ \ j} \qquad {}_{kt} \, , \tag{26}$$

under the assumption that 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 ${}_{k} \quad 0$ 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

[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

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:

$$p_{kt} \qquad {}_{k} \qquad {}_{k} a_{kt} \qquad {}_{k} m_{kt} \qquad {}_{kt} \tag{27}$$

for k=1,…,12 countries and a total of 304 observations for different time periods between 1970 and 2006 (see Table 1). ${}_{k}$ is a country-specific factor and ${}_{kt}$ *is i.i.d.(0,* ^{2} *)*, capturing stochastic deviations from the equilibrium relationship given by equation (9). The coefficients ${}_{k}$ and ${}_{k}$ 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 ${}_{k}$, ${}_{k}$, ${}_{k}$ k and assuming that

${}_{kt}$ k *is i.i.d.*

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 $E(\ _{kt}\ _{ks})=\ _{ts}$ $\ k,t,s$ 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.

Given that previous empirical models impose homogeneity in the slope coefficients across countries (${}_{k}$ and ${}_{k}$), 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:

$$p_{kt} \quad {}_{k} \quad {}_{1k} a_{kt} \quad {}_{2k} a_{kt\ 1} \quad {}_{1k} m_{kt} \quad {}_{2k} m_{kt\ 1} \quad {}_{k} p_{kt\ 1} \quad u_{kt} . \qquad (28)$$

The error correction equation is therefore:

$$p_{kt} \quad {}_{k}(p_{kt\ 1} \quad {}_{k} \quad {}_{k} a_{kt} \quad {}_{k} m_{kt}) \quad {}_{1k} \ a_{kt} \quad {}_{1k} \ m_{kt} \quad u_{kt} , \qquad (29)$$

where ${}_{k} \quad (1 \quad {}_{k})$ is the speed of adjustment coefficient, ${}_{k} \quad {}_{k}/(1 \quad {}_{k})$, ${}_{k} \quad ({}_{1k} \quad {}_{2k})/(1 \quad {}_{k})$ and ${}_{k} \quad ({}_{1k} \quad {}_{2k})/(1 \quad {}_{k})$. The PMG estimate is based on equation (29), under the restriction that all long-run coefficients are equal across countries, ${}_{k}$ and ${}_{k}$, allowing thus for unrestricted country heterogeneity in the adjustment dynamics. The disturbances u_{kt} have zero mean and variance ${}^{2}_{k}$. 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

We introduced imperfect competition in

Arellano, M. and Bover, O. (1995). ‘Another Look at the Instrumental Variables Estimation of Error Component Models’, *Journal of Econometrics*, Vol. 68, pp. 29-52.

Balassa, B. (1964). ‘The Purchasing Power Parity Doctrine: a Reappraisal’, *Journal of Political Economy*, Vol. 72, pp. 584-596.

Baltagi, B.H. (2008). ‘Econometric Analysis of Panel Data’, 4th edition New York: Wiley.

Bergstrand, J. (1991). ‘Structural Determinants of Real Exchange Rate and National Price Levels: Some Empirical Evidence’, *American Economic Review*, Vol. 81, pp. 325-334.

De Gregorio, J., Giovannini, A. and Wolf, H. C. (1994). 'International Evidence on Tradables and Nontradables Inflation', *European Economic Review*, Vol. 38, pp. 1225-1244.

DeLoach, S. (2001). 'More Evidence in Favor of Balassa-Samuelson Hypothesis', *Review of International Economics*, Vol. 9, pp. 336-342.

Domar, E. D. (1962). 'On Total Productivity and All That', *Journal of Political Economy*, Vol. 70, No. 6, pp. 597-608.

Doornik, J.A., Arellano, M. and Bond, S. (2006). 'Panel Data Estimation using DPD for Ox', http://www.doornik.com/download/dpd.pdf.

Levin, A., Lin, C.F. and Chu, C. (2002). 'Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties', *Journal of Econometrics*, Vol. 108, pp. 1-24.

Maddala, G.S. and Wu, S. (1999). 'A Comparative Study of Unit Root Tests with Panel Data and A New Simple Test', *Oxford Bulletin of Economics and Statistics*, Vol. 61, pp. 631-652.

Méjean, I. (2008). 'Can firms' location decisions counteract the Balassa–Samuelson effect?', *Journal of International Economics*, Vol. 76, pp. 139-154.

Obstfeld, M. and Rogoff, K. (1996). 'Foundations of International Macroeconomics', MIT Press.

Pesaran, M.H. (2007). 'A Simple Panel Unit Root Test in The Presence of Cross-section Dependence', *Journal of Applied Econometrics*, Vol. 22, pp. 265-312.

Pesaran, M.H., Shin, Y. and Smith, R. (

TABLE 1

Time period samples for 12 OECD countries

	Coverage	Average Share of Non-Tradeables (%)	Average $^{N}/^{T}$
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. ^{N} (^{T}) is the capital-labour ratio for non-tradeables (tradeables). The average $^{N}/^{T}$ is the sample average of ^{N} over ^{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

	N	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	N	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	$0.07^{\dagger}$	0.53**	$-0.75^{\dagger}$
PE	-7.52*	-8.46*	-9.17*

Notes: Abbreviations as follows: LLC, Levin *et al.* (2002); BRE, Breitung (2000); IPS, Im

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

				$Adj. R^2$	Sargan test
GPOLS	0.0063 (0.0010) 0.0176 (0.0014)	0.7325 (0.0034)	-0.8791 (0.0295)		

TABLE 8

Pooled estimates for the relative price equation

	MG	PMG	Hausman test	DFE
	0.7594* (0.1301)	0.7606* (0.0592)	0.997**	0.6527* (0.0626)
	-0.8156* (0.1615)	-1.0801* (0.0357)	0.569†	-0.9144* (0.0380)
	-0.7881* (0.0687)	-0.5805* (0.1112)		-1.0025* (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