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SECTORAL PRODUCTIVITY, GOVERNMENT SPENDING AND REAL EXCHANGE RATES: EMPIRICAL EVIDENCE FOR OECD COUNTRIES

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ABSTRACT

This paper investigates the long- and short-run determinants of the real exchange rate using a panel of data for fourteen OECD countries. The data are analyzed using time series and panel unit root and panel cointegration methods.

Two dynamic productivity-based models are used to motivate the empirical exercise. The candidate determinants include productivity levels in the traded and in the nontraded sectors, government spending, the terms of trade, income per capita, and the real price of oil. The empirical results indicate that it is easier to detect cointegration in panel data than in the available time series; moreover, the estimate of the rate of reversion to a cointegrating vector defined by real exchange rates and sectoral productivity differentials is estimated with greater precision as long as homogeneity of parameters is imposed upon the panel. It is more difficult to find evidence for cointegration when allowing for heterogeneity across currencies.

The most empirically successful model of the real exchange rate includes sectoral productivity measures in the long run relation and government spending in the short run dynamics.

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"[T]here is some evidence that real exchange rates tend to be higher in rich countries than in poor countries, and that relatively fast-growing countries experience real-exchange rate appreciations. But the empirical evidence in favor of a 'Balassa-Samuelson' effect is weaker than commonly believed, especially when comparing real exchange rates across industrialized countries over the post-Bretton Woods period."

Froot and Rogoff (1995: 1648)

1. Introduction

In light of Froot and Rogoff's recent assessment in the *Handbook of International Economics*, it is unsurprising that there has been a resurgence of interest in the productivity growth as a determinant of real exchange rate movements.[1] On the other hand, it *is* surprising that, despite the amount of attention lavished upon the subject, it has been difficult to find robust evidence for the Balassa-Samuelson effect in levels. This paper addresses this deficiency. It does so by investigating the long- and short-run determinants of the real exchange rate using a panel of data for the OECD countries, focusing on productivity differentials and government spending. The data are analyzed using time series and panel unit root and panel cointegration methods. However, the most fruitful results are obtained when I exploit the cross-currency information available in the panel data. To anticipate the results, I find that in this latter case, there is substantial evidence of a long run relationship between

[1] See for instance the October 1994 issue of the *Review of International Economics* dedicated to the topic, as well as Chapter 4 of Obstfeld and Rogoff's (1996) textbook.

real exchange rates, and sectoral productivity differentials; government spending also has an effect, but only in the short run.[2]

These results represent an innovation in the literature, since previous attempts using time-series techniques have had only limited success in confirming the existence, and extracting *plausible* estimates, of long-run cointegrating vectors from the short spans of data available. On the other hand, with few exceptions, other panel studies have investigated only the short-run relationship between real exchange rates and productivity differentials.[3]

The paper is organized in the following manner. Section 2 critically reviews the previous literature. Section 3 describes the model used to motivate the analysis. Section 4 describes the time series techniques implemented and results. Section 5 discusses the panel regression techniques and estimation results. Section 6 applies the Pedroni (1995) panel cointegration test to the data. Section 7 concludes.

2. A Critical Literature Review

Most earlier analyses of productivity-based models of the real exchange rate constitute straightforward interpretations of the Balassa (1964) and Samuelson (1964) approach. In those models, the relative price of nontradables is determined exclusively by supply side factors, such as productivity. A subset of this group introduces some type of rigidity, such as sectoral

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[2] It might seem that this paper's thesis is at variance with a developing consensus, built on long spans of data (Lothian and Taylor, 1996), and on large panels of data (Frankel and Rose, 1996), that purchasing power parity (PPP) holds. However, Engel (1996) has pointed out that the statistical detection of real exchange rate reversion may indicate reversion to some equilibrium value that is not necessarily a constant.

[3] See Chinn and Johnston (1996) and Canzoneri, Cumby and Diba (1996) for two exceptions. A related analysis in a non-cointegration framework is Asea and Menzdoza (1994).

re-allocation costs or imperfect capital markets, so that demand factors also determine the real exchange rate.

Suppose the price level can be expressed as a geometric average of the tradable and nontradable goods price indices, where all variables are expressed in logs:

$$p_t = \Omega p_t^N + (1 - \Omega)p_t^T \tag{1}$$

Then defining the real exchange rate as the aggregate price index deflated exchange rate yields the following expression, assuming purchasing power parity (PPP) holds for tradable goods.

$$\begin{aligned} q_t &\equiv (s_t + p_t^* - p_t) \\ &= \Omega(s_t + p_t^{N*} - p_t^N) \end{aligned} \tag{2}$$

where s is the nominal exchange rate, and an asterisk denotes the foreign country. Equation (2) states that the real exchange rate is a function of the relative price of nontradables. This point has been incorporated in various models of the nominal exchange rate where the long-run real exchange rate is allowed to vary over time.[4]

Assuming perfect international integration of goods and capital markets, the price of tradables and the interest rate are set. The former then determines the wage rate, which given intersectoral factor mobility means that relative prices are set exclusively by the level of productivity in the two sectors. Since both factors are free to move between sectors costlessly, only supply side factors matter. One then obtains:

[4] Clements and Frenkel (1983), Wolff (1987), Chinn and Meese (1995).

$$q_t = -\Omega[(\frac{\theta^N}{\theta^T})a_t^T - a_t^N] + \Omega[(\frac{\theta^{N*}}{\theta^{T*}})a_t^{T*} - a_t^{N*}] \quad (3)$$

where Θ is the labor coefficient in a Cobb-Douglas production function, a is log-total factor productivity, and the "T" and "N" superscripts denote tradable and nontradable variables and parameters, respectively. While this is the most common derivation, one can obtain similar relationships in a variety of model frameworks, such as the Ricardian or factor endowments (see Dornbusch's (1987) survey of purchasing power parity for a discussion).

Typically, the regressions are implemented in the following form:

$$\Delta q_t = \beta_0 + \beta_1 \Delta(a_t^T - a_t^N) + \beta_2 \Delta(a_t^{T*} - a_t^{N*}) + \textit{other regressors} + u_t \quad (4)$$

In one of the earliest studies of this type, Hsieh (1982) estimates the determinants of the multilateral exchange rates for Germany and Japan over the 1954-76 period, using manufacturing and service sector labor productivity levels as the explanatory variables. He finds that the coefficient β_1 (β_2) is -0.362 (.516) for Germany, and -0.538 (.538) for Japan, allowing for deviations from PPP for tradable goods. The β_1 coefficient is interpretable (approximately) as the share of nontradables in the aggregate price index. Marston (1990) adopts a similar approach, examining five bilateral exchange rates over the 1973-86 period. He obtains estimates considerably higher, ranging from -0.714 for the Franc/Deutschemark rate, to -1.244 for the Dollar/Deutschemark (β_1 is constrained to equal -β_2). Micossi and Milesi-Ferretti (1994) estimate a similar relationship for multilateral real exchange rates over the 1970-1990 period. They find that the estimates for β_1 (β_2) range from -0.10 (-0.05) to -0.76 (1.10). All the coefficients are correctly-signed except for the case of Denmark.

These studies indicate that there is a statistically significant relationship between changes in labor productivity and changes in the real exchange rate. However, the results differ by specification, by sample and data type. Moreover, it is not clear how good a proxy labor productivity is for total factor productivity (TFP), the theoretically-implied variable of interest.

DeGregorio and Wolf (1994) address this issue. In their model demand side factors affect the real exchange rate if either the assumption of perfect competition, PPP for traded goods, or perfect capital mobility are relaxed. Hence their specification nests the standard Balassa-Samuelson specification, insofar as both supply and demand shocks have an effect on the real exchange rate.

DeGregorio and Wolf estimate a number of first-differenced specifications on a panel of 14 countries; the specifications include terms of trade effects, government spending shocks, and changes in preferences regarding the consumption of nontradables, proxied by the income level. They also utilize a composite weighted average of traded and nontraded sector TFP as their productivity measure, and obtain statistically significant estimates of β_1 ranging from -0.10 to -0.26.[5]

Their results also indicate that the coefficient on a preferences variable, where preferences are proxied by income per capita, is not robust to the inclusion of terms-of-trade shocks. However, this outcome may be partly a consequence of the choice of estimating in first differences, since the low-frequency effect of changing tastes is unlikely to be manifested in year to year changes in income.

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[5] Technically the β_1 coefficient pertains to the composite variable $a^T - \mu a^N$, where μ depends on various production parameters and the rate of return to capital.

This last point leads to a more substantive problem. If the Balassa-Samuelson model is taken at face value, then these models are mis-specified since the level of the real exchange rate should move in tandem with the level of the relative productivity differential. This conclusion holds regardless of whether the series are taken to be trend stationary or difference stationary, although the implications differ. Consider the first alternative; if all series are trend stationary, then first differencing the variables induces a unit moving average error into each series. The diagnostics associated with such a regression will likely indicate the presence of serial correlation that should be addressed by some form generalized least squares. Furthermore, the estimate of the short-run dynamics will be inefficient.

Now consider the second alternative, wherein both series are integrated of order one. If the series fail to exhibit cointegration, then first differencing to avoid the problem of spurious correlation is indeed the correct procedure. However, the failure to find cointegration is indicative of model mis-specification since the theory implies a given long run relationship.

If each of the series individually contain a unit root, but together form a linear combination that is stationary, then the series are cointegrated.[6] Strauss (1996) examines the relationship in levels for six bilateral exchange rates against the Deutschemark (Belgian Franc, Canadian Dollar, Finnish Marka, French Franc, Pound, and US Dollar). Using the Johansen (1988) multivariate approach, he detects evidence of cointegration between real exchange rates and sectoral *labor* productivity. However, his estimates for β_1 (β_2) range from -1.21 (-8.72) to -10.53 (13.97)! In all cases the β_1 coefficient rejects the null hypothesis of $\beta_1 = 0$, but

[6] I omit discussion of cointegration studies using quarterly labor productivity data in manufacturing, such as Chinn (1995) and Chinn (1996), since these papers omit a proxy variable for nontradable productivity.

also rejects the null of $\beta_1 = 0.5$, which is what would be expected if about half of the CPI was accounted for by nontraded goods. Analogous results hold for his estimates of β_2, although with slightly less force (the null of $\beta_2 = -0.5$ is not rejected in two cases).

Strauss (1995) addresses both the TFP and cointegration issues. Again using the Johansen procedure he tests for a cointegrating relationship between the bilateral real exchange rate (versus Deutschemark) and relative productivity variables, where total factor productivity (TFP) instead of labor productivity is now used. While TFP is the appropriate variable, it also limits the span of the data series for five of 14 countries to 21 years. Using the conventional asymptotic critical values from Osterwald-Lenum (1992), he finds that eight cases are cointegrated at the 10% marginal significance level. However, if one adjusts for small sample effects (as in Cheung and Lai, 1993), then the number of cases of cointegration drops to a mere two: UK and possibly France.[7] Under no conditions does Japan exhibit cointegration, which is odd, given the apparent fit of the Japanese case. This oddity suggests the low power of this approach given the short span of data.[8]

Kakkar (1996) has also investigated this relationship for Japanese, Italian, German, and Canadian bilateral rates (against the US Dollar). He applies the Johansen procedure in

[7] Godbout and van Norden (1995) also demonstrate the importance of this issue of adjusting for small sample effects.

[8] Strauss does not report the parameter estimates obtained from the Johansen procedure, so it is difficult to evaluate the conformity of the results with any particular theoretical model. He does report likelihood ratio tests for restrictions on the cointegrating vector. In general the cointegrating vector linking productivity and relative prices, and the cointegrating vector linking relative prices and the real exchange rate, reject the implied restrictions. However, two points are relevant. First, the validity of such tests are conditional upon the existence of a cointegrating vector, which is in doubt. Second, these tests may also be sensitive to finite sample size effects; see Edison, Gagnon and Melick (1994) and Godbout and van Norden (1995).

conjunction with the Park (1992) Canonical Cointegrating Regression (CCR) methodology, and finds that he generally can reject the null of no cointegrating vectors using the Johansen procedure, and cannot reject the null of cointegration according to the CCR results. The results are not directly comparable to those in the other studies, since the key cointegrating vector contains the real exchange rate, US TFP in tradables, and US TFP in nontradables. This specification is selected since Kakkar finds evidence of another cointegrating relation between US tradable TFP, foreign tradable and nontradable TFP. The US tradable TFP coefficient does not then have a structural interpretation, while that on nontraded TFP has the correct sign, but is far too large to be rationalized by the Balassa-Samuelson model.[9]

In summing up the literature, one can conclude that there is some evidence for a productivity based model of the real exchange rate. However, due to statistical and data limitations, one cannot conclude that there is a robust relationship between the *level* of the productivity differential and the *level* of the real exchange rate.

3. The Theoretical Model

Froot and Rogoff (1991) present an intertemporal model incorporating nontradables. Consumption smoothing can only by mediated by exchange in tradables; since consumption of nontradables must match production, government demand shocks that fall on tradables and nontradables in different proportions than those of the private sector will have an affect on exchange rates. Assuming endogenous output and fixed sectoral capital, the relative price of nontradables in terms of tradables is then a function of productivity differentials, although the

[9] The point estimates are usually ten times too large.

intertemporal character of the model means that this price only responds to unanticipated productivity shocks. The implied long-run relationship is then:

$$p_t^N \quad p_t^T = a_t^T \quad a_t^N \tag{5}$$

assuming that the productivity differentials follow random walk processes such that the unanticipated component is a true innovation.[10]

By using the real exchange rate definition in equation (2)

$$q_t \equiv s_t + p_t^* \quad p_t$$

and appropriate substitution one obtains the following expression for the real exchange rate,

$$q_t = -\Omega[\hat{a}_t^T \quad \hat{a}_t^N] \tag{6}$$

where the hats (^) denote differences relative to the foreign country.

While the model superficially resembles the static Balassa-Samuelson formulation, the implied time series behavior of the real exchange rate is different. Anticipated or trend stationary movements in sectoral productivity have no effect on the real exchange rate, while in the Balassa-Samuelson model the exchange rate would move in tandem with productivity regardless of whether its evolution was anticipated or not.[11]

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[10] A different specification follows if the productivity processes follow I(1) processes but also exhibit higher order autoregressive behavior.

[11] Note that such a model is somewhat at variance with technology diffusion models wherein productivity levels in sector i must be cointegrated across countries. For evidence that industry productivity growth rates covary more with*in* a country than between industries of the same category in different countries, see Costello (1993)).

Rogoff (1992) extends this intertemporal model to the case of Japan, allowing for fixed factors, in order to account for the stylized fact that there is high persistence in the real exchange rate, without relying on unit root productivity shocks. The open capital account version of this model implies that unanticipated productivity shocks cause highly persistent movements in the real rate, and technically indicates a regression with near-unit root variables. If one assumes random walk processes underlying the productivity and government spending variables, then one obtains:

$$q_t = -\Omega[\hat{a}_t^T - \zeta_N \hat{a}_t^N + (\zeta_N - 1)\hat{g}_t] \tag{7}$$

where ζ_N is the ratio of nontraded goods output to private nontraded goods consumption.[12] The derivation assumes some initial fixed condition for all the variables. This equation provides us with a theoretically implied cointegrating relationship between the real exchange rate, relative productivity levels in the tradables and nontradables sectors, and government spending (expressed as a log proportion of GDP).

In the empirical portion of the paper, the robustness of the basic regression specifications is evaluated by including other candidate regressors, such as per capita income and the terms of trade, suggested by DeGregorio and Wolf (1994). Income per capita is included as a proxy for non-homotheticity of consumption preferences; that is, as income or wealth rises, consumer preferences shifts toward nontraded goods, such as services.[13] Changes

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[12] An explicit derivation of this expression is found in Chinn (1996). Note that this expression differs from Rogoff's (1992) equation (21), in that here ρ, the autoregressive coefficient on tradables productivity, is set to 1. If $\rho < 1$, then the implied time series process for all variables would be trend stationary. We view this as an empirical issue, to be addressed in Section 4.

[13] Technically, this non-homotheticity is inconsistent with the Rogoff model.

in the terms of trade have obvious implications for intratemporal prices; however, such shocks can also have intertemporal effects via wealth revaluations, and intertemporal prices. These intertemporal channels are examined, for instance, by Roldos (1990). On the supply side, the equation is augmented with the real price of oil, to account for possible shifts in the production function.[14]

4. TIME SERIES APPROACHES

4.1. Time Series Econometric Methodologies

The current standard in testing for cointegration in time series is the full-system maximum likelihood estimation technique of Johansen (1988) and Johansen and Juselius (1990). Cheung and Lai (1993), among others, have shown that finite sample critical values may be more appropriate given the relatively small samples which are generally under study.[15] Using these critical values we do not find any evidence for cointegration (in results at variance with Strauss, 1995).

I focus on estimates derived from nonlinear least squares (NLS) regression. Phillips and Loretan (1991) argue that the following NLS estimator is optimal among single-equation estimators:

[14] See also Amano and van Norden (1995) for empirical evidence on the role of oil prices in real exchange rate determination.

[15] The finite sample critical values are obtained by adjusting the asymptotic critical values for the loss of degrees of freedom due to the estimation of the parameters describing the short term dynamics. The adjustment factor is given by $(N - (p \times m))/N > 1$.

$$
\begin{aligned}
\Delta q_t = \phi_0 + \phi_1(q_{t-1} - BX_{t-1}) \\
+ \Psi\Delta X_{t+1} + \Xi\Delta X_{t-1} + u_t
\end{aligned} \tag{8}
$$

where X is a vector of explanatory variables. The leads of the differences of the right hand side variables serve to orthogonalize the error term, in the presence of endogenous RHS variables.

Since there are a very limited number of observations in each time series, I also consider a simple error correction model which excludes the leads of the differenced variables.

4.2. Data

The data are annual, covering the 1970-91 period for the United States (USA), the United Kingdom (GBR), France (FRA), Germany (DEU), Italy (ITA), Canada (CAN), Japan (JPN), Belgium (BEL), Denmark (DNK), Netherlands (NLD), Norway (NOR), Sweden (SWE), Finland (FIN) and Australia (AUS) (see the Data Appendix for details). Multilateral real exchange rates are calculated using nominal exchange rates and consumer price indices (CPIs) drawn from the IMF's *International Financial Statistics*, and weighted using trade weights used by the IMF to construct the multilateral effective exchange rates reported in *IFS*. The terms of trade and the price of oil are from the same source. The former is calculated as the log-ratio of export prices to import prices (in US dollars). The latter is the log price in US dollars, deflated by the US CPI.

The sector total factor productivity (TFP) data were constructed from the OECD's *International Sectoral Data Base* (*ISDB*) which contains industry-specific TFP data. The details of how the industry TFP were constructed can be found in Meyer zu Schlochtern and

Meyer zu Schlochtern (1994). The tradable and nontradable categorization is the same as that used by DeGregorio, Giovannini and Wolf (1994). Tradable sectors include agriculture, mining, manufacturing, and transportation, while the nontradable sectors include all other services.

The government spending variable is the log of the ratio of real government consumption to real GDP. These data come from the OECD's *National Accounts*.

The "preferences" variable is GDP per capita, where GDP is measured in Summers and Heston "International" dollars (the chain-weighted variable, RGDPCH). This variable is meant to proxy for the rising preference for services as income rises. The data were drawn from the Penn World Tables, Mark V. The rest-of-world TFPs, government spending ratios, and incomes per capita are calculated using the same weights used to calculate the real multilateral exchange rates.

4.3. Time Series Results

Preliminary analysis of the data indicated that the relevant variables could not generally reject the null hypothesis of difference stationarity at the 5% MSL, using an ADF test (1 lag, w/trend). There were three exceptions: tradable productivity differentials for Finland, nontradable productivity differential for Great Britain and the per capita income differential for Belgium. Since these three cases constitute far fewer than 5% of the total number of cases, one can be reasonably confident that the series are all I(1).

I examined the Phillips-Loretan single-equation estimates for the simple ECM specification. The basic specification is one that allows for the relative productivity variables to enter in the long run relationship, but government spending only enters in the short run dynamics, viz.

$$\begin{aligned}\Delta q_t &= \phi_0 + \phi_1(q_{t-1} - \beta_1\hat{a}^T_{t-1} - \beta_2\hat{a}^N_{t-1})) \\ &\quad + \kappa_1\Delta g_t + \kappa_2\Delta g^*_t + \gamma\Delta q_{t-1} + \nu_1\Delta\hat{a}^T_{t+1} + \nu_2\Delta\hat{a}^N_{t+1} + u_t\end{aligned} \tag{9}$$

This specification is chosen because in preliminary analysis, it was determined that government spending does not enter into the cointegrating vector with statistical significance.[16] The inclusion of *contemporaneous* home and foreign government spending variables is consistent with weak exogeneity of these variables.

Estimating the β_1 and β_2 parameters freely produces extremely poor results; hence I impose an a priori value of $\beta_1 = -\beta_2$, and collapse the lead terms into a single one.

$$\begin{aligned}\Delta q_t &= \phi_0 + \phi_1(q_{t-1} - \beta_1\hat{z}_{t-1})) \\ &\quad + \kappa_1\Delta g_t + \kappa_2\Delta g^*_t + \gamma\Delta q_{t-1} + \nu\Delta\hat{z}_{t+1} + u_t \\ \textit{where}\ \ \hat{z}_t &\equiv \hat{a}^T_t - \hat{a}^N_t\end{aligned} \tag{10}$$

This specification yields slightly more plausible estimates, reported in Table 1. The adjusted-R^2 ranges from 0.60 to a high of 0.91, indicating a substantial proportion of the variation is accounted for. Moreover, the errors appear serially uncorrelated, according to the Box-Ljung Q statistic, of order 2, with the exception of the Australian equation. Since there are effectively only 12 observations, the coefficients are very imprecisely estimated.

[16] This result is at variance with results obtained in Chinn and Johnston (1996). I conjecture that the difference occurs because the variables in that study were expressed relative to the US, which undertook a significant expansion of government spending in the mid-1980s.

In order to conserve degrees of freedom, the equations are re-estimated in a simple error-correction specification. These estimates, reported in Table 2, indicate that there is substantial, and statistically significant, evidence of reversion to conditional mean (see also Figure 1). These results obtain regardless of whether a time trend is included or not. The rate of reversion is quite rapid in certain cases (0.70 for Sweden) and quite low in others (0.15 for Finland). The Australian rate is even lower, but the associated estimate for β_1 is so implausible that I exclude them from the discussion.

More surprising is the fact that the coefficient on the composite productivity variable, the key variable in all these productivity-based model, is incorrectly signed in almost a third of the cases, and significantly so (in both economic and statistical terms) in one case. There is quite wide variation in the point estimate, as shown in Figure 2. That being said, these results are much more in line with expectations than those for bilateral rates (against the US$) reported in Chinn and Johnston (1996).[17]

The lack of robust results here mirrors those obtained by other researchers. Domestic government spending and foreign government spending coefficients (Figures 3 and 4 respectively) are correctly signed in most instances, but not statistically significant. Notice that these are short run coefficients; government spending is not included in the cointegrating relation.

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[17] These results are sensitive to the inclusion of a time trend; the Japanese coefficient for instance switches sign from roughly -2 to +0.4. The trend term is difficult to interpret, since it implies a quadratic time trend in the cointegrating vector.

If one were to end the analysis at this point, one would be forced to conclude that the evidence in favor of the levels version of the Balassa-Samuelson hypothesis was quite weak. This suggests an alternative approach.

5. THE PANEL DATA APPROACH

The motivation for appealing to cross-section information is apparent from a visual inspection of the data. For instance, Figure 5 presents a scatterplot of the real exchange rate (in units of home goods per unit of foreign) against the differential in log tradable productivity. There is a clear negative relationship that is not readily apparent in individual time series plots. On the other hand, there is no readily apparent relationship between nontradable productivity differentials and the real exchange rate in Figure 6. This outcome probably arises from the difficulties in measuring service sector productivity. Note that constraining the nontraded productivity differential to 0.4 yields the expected correlation between the exchange rate and productivity (Figure 7).

Figure 8 displays the motivation for examining the relationship between the real exchange rate and government spending in changes; in contrast an examination of the relation in levels reveals much less correlation (Figure 9). Although there is a slight negative relationship, most of it appears to be due to the Japanese observations, which are far to the left of the others.

Figure 10 displays the obvious negative correlation between the real exchange rate and the terms of trade. The reason this correlation arises may be that exogenous terms of trade shocks induce real exchange rate movements; alternatively, exogenous nominal exchange rate movements combined with incomplete pass through may be the source of the positive correlation.

Figure 11 depicts the well known "Penn-effect" (Samuelson, 1993). The higher per capita income, the higher the relative price level. Finally, Figure 12 shows the relationship between the real price of oil and the real exchange rate. No clear correlation arises in this graph.

I now turn to statistically investigating the robustness of these correlations.

5.1. Panel Regression Methodology

I consider variants of equations (6) and (7) where the data are indexed by country:

$$\begin{aligned} q_{i,t} &= -\Omega[\hat{a}_{i,t}^{T} - \hat{a}_{i,t}^{N}] \\ q_{i,t} &= -\Omega[\hat{a}_{i,t}^{T} - \zeta_N \hat{a}_{i,t}^{N} + (\zeta_N - 1)g_{i,t} - (\zeta_N - 1)g_{i,t}^{*}] \end{aligned} \tag{11}$$

The application of conventional panel regression techniques is complicated by the possible nonstationarity of the series involved. As a consequence, one must proceed with caution. I estimate the cointegrating relationships using NLS, and then test whether the residuals from these cointegrating vectors are stationary according to a unit root test.

I estimate the cointegrating relationship with individual-specific effects only in the constant, and homogeneity imposed across the individual slope coefficients. The cointegrating vectors define residuals which should be stationary. To determine whether the variables are cointegrated, the differenced regression residual (ECT) is regressed on the lagged residual, and country dummies.

$$\Delta ECT_{i,t} = \quad + \alpha ECT_{i,t-1} + \textit{currency dummies} + u_{i,t} \tag{12}$$

The t-statistic on the α coefficient is then compared against the tabulated critical values in Table 5 of Levin and Lin (1992). If the test statistic is statistically significant, then the null of no cointegration can be rejected.

Before conducting this exercise, this test was applied to a composite variable where the *theoretical* priors are imposed. These priors are: the share of nontradables in the CPI is set to 0.5 and the coefficients on the two productivity variables are of equal and opposite sign. The resulting t-statistic on the error correction term is -4.466, which exceeds the Levin-Lin 1% MSL critical value.[18] Hence, there is evidence that real exchange rates are cointegrated with productivity differentials even when using imposed coefficients.

5.2. Estimating the Cointegrating Relationships

Consider equation (8) rewritten in panel error correction form:

$$\Delta q_{i,t} = \phi_{i,0} + \phi_1(q_{i,t-1} - BX_{i,t-1}) + \Xi\Delta X_{i,t-j} + u_{i,t} \tag{13}$$

where Ψ is set to zero, and the slope coefficients ϕ_1 and B are restricted to be the same across all currencies. This equation is estimated using NLS with different variables in X. The results are reported in Table 3.

The estimated coefficient of reversion in the most basic formulation which includes *only* productivity effects (Column 1) is statistically significant, as is the coefficient on

[18] Note that the Levin and Lin procedure assumes independence of errors across individuals (here currencies). O'Connell (1996) has shown that allowing for cross-correlation increases the nominal size of such tests.

tradables productivity. The estimated rate of reversion is about 0.24, implying that the half-life of a deviation from trend is about 2.5 years. This is substantially faster than the four to five years purchasing power parity deviation half-life reported by Edison (1987), Frankel and Rose (1996) and Wei and Parsley (1995), for instance. The estimated coefficient for the tradable productivity differential if fairly plausible -- -0.485, indicating a 1% increase in relative traded sector productivity induces roughly a half percent appreciation in the real exchange rate. This result would be expected if half of the CPI was nontraded. The coefficient estimate for the nontradable productivity differential is not statistically different from zero, although the point estimate is plausible, and consistent with the Froot-Rogoff specification.

In column (2), I allow for contemporaneous, short run government spending effects. In this specification, the rate of reversion and the long run productivity coefficients are relatively unchanged. Furthermore, the short run government spending coefficients are statistically significant. They indicate that a 1% increase of government spending on goods and services relative to GDP induces an approximately one half percent appreciation in the real exchange rate, while an analogous ROW increase induces a 0.4 depreciation.

I test for the presence of long run government spending effects in column (3). In this case, the rate of reversion is unchanged, but the economic and statistical significance for the tradable sector productivity differential drops. Moreover, the estimates for the long run government spending coefficients are neither statistically significant, nor of correct sign.

If the specification in column (3) is augmented with a terms of trade variable in the long run relation, then all the long run coefficients become statistically insignificant (column 4). The long-run terms of trade coefficient itself is not statistically significant, which contrasts strongly with the results obtained by DeGregorio and Wolf (1994). This outcome suggests that terms of trade effects may have their greatest impact on exchange rates at high frequencies; there is also a problem with simultaneity in their regressions, since contemporaneous terms of trade are used (in the regression reported here, the lagged difference of the terms of trade enter with correct sign, but of an economically and statistically small magnitude).

The regression incorporating the income per capita variable is very successful in some respects. The estimated coefficient is significant at the 10% MSL. Yet, its inclusion causes the tradables and nontradables productivity coefficients to become statistically insignificant; this result suggests the presence of multicollinearity.

Finally, inclusion of the price of oil in column 6 yields in some respects unsatisfactory results. While the rate of reversion is still statistically significant, with the exception of the tradable productivity differential, the other coefficients are not statistically significant. The price of oil itself is not significant, suggesting that the cointegrating vector should not include this variable. This is an odd finding, at variance with Chinn and Johnston (1996). One surmise is that by pooling oil exporters and importers together, we conflate offsetting wealth effects. I re-estimate the equation, excluding Great Britain, Canada and Norway. In this case the results are much the same.

The fit of the preferred model is surprisingly good, considering the inadequacies of the data and the imposition of common slope coefficients. Using only the *long-run* coefficients in Column (2) of Table 3 to define the stochastic trend exchange rate, the correlation between the actual and trend is depicted in Figure 13.

One alternative interpretation of these findings is that they are due to a statistical artifact. Total factor productivity is measured as a residual of output and factor inputs. It is possible that exchange rate appreciation reduces the prices of imported goods which serve as intermediate inputs in the production process. This may in turn cause measured TFP to look larger, when in fact the calculated change is entirely due to mis-measurement. The correlation is once again negative, but for a different reason than that posited in the theoretical section of the paper.[19]

I attempt to address this concern by considering an empirical implication of this reverse causation. If exchange rate movements induce the mis-measurement of TFP, one might expect that the real exchange rate, or the cointegrating vector, should Granger cause TFP. In fact, one does not find that this is the case in this data set. The regression results are:

$$\Delta q_t = \underset{(0.017)}{-0.079}\ (q + 0.5(\hat{a}^T - \hat{a}^N))_{t-1} + v_{1t} \qquad \bar{R}^2 = 0.08 \tag{14}$$

$$\Delta \hat{a}_t^T = \underset{(0.007)}{-0.008}\ (q + 0.5(\hat{a}^T - \hat{a}^N))_{t-1} + v_{2t} \qquad \bar{R}^2 = 0.02 \tag{15}$$

$$\Delta \hat{a}_t^N = \underset{(0.017)}{-0.007}\ (q + 0.5(\hat{a}^T - \hat{a}^N))_{t-1} + v_{3t} \qquad \bar{R}^2 = 0.04 \tag{16}$$

(heteroskedasticity-consistent standard errors in parentheses); country dummies suppressed.

These results indicate that the cointegrating error is weakly exogenous for Δq; however, the two productivity variables appear to be exogenous, since they are statistically unresponsive to the disequilibrium. Further, the economic magnitudes of these parameters are negligible. Hence, the overwhelming majority of the correlation is attributable to causality running from productivity to the exchange rate.

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[19] We thank Rich Lyons for bringing this issue to our attention.

5.3. Panel Unit Root Test Results

Since the estimated rate of reversion is always statistically significant, one can be reasonably confident that the regressions include the cointegrating vector defined, by at least, the productivity coefficients. The Levin and Lin (1992) statistic for the composite variable implied by each of these sets of estimates is reported at the bottom of Table 3. They indicate that all of the series reject the unit root null. Hence, there appears to be substantial evidence for cointegration.

The specification including the preferences variable garners the greatest support on the basis of the t-statistic. However, given the difficulty in interpreting this particular cointegrating vector, and the lack of statistical significance for any of the other variables in the cointegrating vector, our preferred model is the basic specification including only real exchange rates, productivity levels in the long run relation, and short-run government spending dynamics.

6. PANEL COINTEGRATION TESTS

An alternative approach to the two-step procedure outlined above is to use the panel cointegration methodology of Pedroni (1995). Pedroni has tabulated panel cointegration test statistics for the bivariate case. I consider five cases, where the errors are assumed to be white noise.

Homogeneous Panel:

$$y_{i,t} = X_{i,t}\beta + \epsilon_{i,t}$$
$$y_{i,t} = \alpha + X_{i,t}\beta + \epsilon_{i,t}$$

Heterogeneous Panel:

$$
\begin{aligned}
y_{i,t} &= X_{i,t}\beta_i + \epsilon_{i,t} \\
y_{i,t} &= \alpha_i + X_{i,t}\beta_i + \epsilon_{i,t} \\
y_{i,t} &= \alpha_i + \delta_i t + X_{i,t}\beta_i + \epsilon_{i,t}
\end{aligned}
$$

Since critical values have been tabulated for only bivariate cases, I consider the tests where the cointegrating vector contains only the real exchange rate and the relative productivity differentials:

$$
\begin{gathered}
(q_t\ ,\ \hat{z}_t) \\
\hat{z}_t \equiv \hat{a}_t^{T} - \hat{a}_t^{N}
\end{gathered}
\tag{17}
$$

Conducting this test, one obtains the results in Table 4. The null hypothesis that q and z are cointegrated in a homogeneous panel can be rejected at the 1% MSL, regardless of whether a common constant is included or not.

Similarly if a heterogenous slopes is allowed for, once again the null hypothesis of no cointegration is again reject, but now only at the 10% MSL. If fixed effects in the constant are allowed for, or in the deterministic time trend, then one fails to reject the null hypothesis of no cointegration.

The homogeneous panel results are the most supportive of the maintained hypothesis. In these cases, the productivity long-run parameter estimate is in accord with expectations: the coefficient on z is -0.551 to -0.553, depending upon whether a constant is allowed for. This value is very similar to that obtained using NLS in Section 5, for the basic specification (2).

These results are consistent with the recent panel cointegration work undertaken by Canzoneri, Cumby and Diba (1996). They show that (i) sectoral labor productivity is cointegrated with the relative price of nontradables, and (ii) common currency prices of tradables are cointegrated, when exchange rates are expressed bilaterally against the DM. Taken together (i) and (ii) are consistent with cointegration of real exchange rates and sectoral labor productivity differentials.

7. Conclusions

The main conclusions are as follows:

1. It is difficult but possible to find a cointegrating relationship between the real exchange rate, sectoral productivity levels, 20 years of time series data. However, the implied long-run relationships seldom conform to priors in terms of either coefficient signs or magnitudes.

2. According to the Levin and Lin test, there is more evidence in favor of cointegration when analyzing panel data. The cointegrating vector definitely contains the real exchange rate and relative sectoral productivity levels. It is less likely that other variables fall into this category. However, there does seem to be a short-run, contemporaneous, effect of government spending on the real exchange rate.

3. The estimates of rates of reversion to equilibrium, as well as of the cointegrating vector, are more reliably estimated when using panel data. The half-life of a deviation from trend is on the order of 2.5 to 3 years. A one percent innovation in tradable sector

productivity induces between a 0.5 percent appreciation in the real exchange rate in the long-run.

4. Using a bivariate panel cointegration test due to Pedroni (1995), one finds that there is indeed evidence of a long-run relationship between the real exchange rate and productivity differentials of the posited type, as long as homogeneity is imposed on the slope coefficients.

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Data Appendix

Exchange rates: Multilateral nominal exchange rates, deflated using CPIs, using appropriate weights. Source: IMF, IFS lines neu (exchange rates); 64 (CPIs); and weights from IMF.

Total factor productivity: weighted sums of sectoral TFP. Source: OECD, *International Sectoral Data Base*. Traded sectors: mining, manufacturing, agriculture and transportation. Nontraded: all other services, construction. Rest-of-World variables constructed using same weights as used for exchange rates.

Government spending: log-ratio of real government consumption to real GDP. Source: OECD *National Accounts*. Rest-of-World variables constructed using same weight as used for exchange rates.

Terms of Trade: log ratio of price of exports to price of imports (in US$). Source: IMF, *IFS* lines 74 and 75 respectively.

Preferences: log GDP per capita, in 1985 International Dollars. Source: Summers and Heston *Penn World Tables* Mark V. Rest-of-World variables constructed using same weights as used for exchange rates.

Oil Prices: log crude petroleum export prices, in 1990 US$. IMF *IFS* line 76aad deflated by US CPI (line 64).

Table 1
Phillips-Loretan Estimates of the Error Correction Model

$$\Delta q_t = \phi_0 + \phi_1(q_{t-1} - \beta_1\hat{z}_{t-1}) + \kappa_1\Delta g_{t-1} + \kappa_2\Delta g^*_{t-1} + \gamma\Delta q_{t-1} + \nu\hat{z}_{t+1} + u_t$$
$$\hat{z}_{t-1} \equiv \hat{a}_{Tt-1} - \hat{a}_{Nt-1}$$

	ϕ_1	ϕ_1	κ_1	κ_2	γ	ν	$\overline{R}^2$	Q(2)
	(-)	(-)	(-)	(+)	(?)	(?)		
USA	-0.47**	1.40	-0.60	1.61	0.56*	0.22	.66	0.07 [.97]
GBR	-0.17	-0.70	0.04	-0.91	0.47	0.15	.70	1.25 [.54]
FRA	-0.45**	0.73	-1.24*	0.13	0.19	0.31	.66	0.20 [.90]
DEU	-0.34	-1.40	-0.29	1.23	0.25	0.87	.54	1.42 [.49]
ITA	-0.21	-1.30*	-0.48	-0.49	-0.25	0.01	.90	4.00 [.14]
CAN	-0.90***	-0.41***	0.07	0.55	0.74*	-.21	.76	1.19 [.55]
JPN	-0.22	-2.08	-1.52	2.86	0.13	0.23	.77	0.47 [.79]
BEL	-0.23**	0.12	-0.54	0.49	0.58***	0.11	.89	3.49 [.18]
DNK	-0.30**	0.75	0.21	0.03	0.24	-.26	.83	2.07 [.36]
NLD	-0.42***	-0.33	-1.29**	1.16**	0.24	-.47	.83	0.46 [.80]
NOR	-0.11**	2.04	-0.89**	0.18	0.22	0.09	.91	3.26 [.20]
SWE	-0.65**	-0.40	-1.25	0.38	0.31	0.40	.60	3.33 [.19]
FIN	-0.28	-3.35	0.24	0.00	0.76	-1.44	.81	3.02 [.22]
AUS	-1.17	2.48*	-4.13	5.01*	1.06*	5.10*	.88	11.32[.00]

Notes: Q(2) is the Box-Ljung Q statistic for second order serial correlation [p-values in brackets]. *(**)[***] indicates significance at the 10%(5%)[1%] marginal significance level.

Table 2
Nonlinear Least Squares Estimates of the Error Correction Model

$$\Delta q_t = \phi_0 + \phi_1(q_{t-1} - \beta_1 \hat{z}_{t-1}) + \zeta_1 \Delta g_{t-1} + \zeta_2 \Delta g^*_{t-1} + \gamma \Delta q_{t-1} + u_t$$
$$\hat{z}_{t-1} \equiv \hat{a}_{Tt-1} - \hat{a}_{Nt-1}$$

	ϕ_1	ϕ_1	κ_1	κ_2	γ	$\bar{R}^2$	Q(2)
	(-)	(-)	(-)	(+)	(?)		
USA	-0.45***	2.02	-0.40	1.41	0.45*	.70	0.37 [.98]
GBR	-0.16	-1.68	-0.12	0.78	0.47*	.76	0.91 [.63]
FRA	-0.45**	0.55*	-1.00	0.13	0.02	.64	0.63 [.73]
DEU	-0.51**	-0.60	0.42	0.32	0.36	.46	1.99 [.37]
ITA	-0.30***	-1.59***	0.12	01.12*	-0.72*	.96	0.98 [.61]
CAN	-0.52**	-0.27*	-0.50	0.70	0.40	.70	2.77 [.25]
JPN	-0.22	-2.07	-0.93	2.24	0.02	.77	1.18 [.55]
BEL	-0,26***	0.08	-0.42	0.42	0.55	.85	3.01 [.22]
DNK	-0.26**	1.04	0.09	0.26	0.20	.81	3.27 [.20]
NLD	-0.43***	-0.30	0.87*	0.91*	0.19	.79	1.13 [.57]
NOR	-0.16**	1.12	-0.87**	0.35	0.28	.91	2.72 [.26]
SWE	-0.70***	-0.53*	-1.48**	0.15	0.24	.68	2.81 [.25]
FIN	-0.15	-1.06	0.39	0.21	0.62*	.85	2.95 [.23]
AUS	-0.13	18.29	0.32	2.54	0.48	.68	0.88 [.64]

Notes: Q(2) is the Box-Ljung Q statistic for second order serial correlation [p-values in brackets]. *(**)[***] indicates significance at the 10%(5%)[1%] marginal significance level.

Table 3
Panel Estimation Results:
Determinants of the Multilateral Real Exchange Rate

$$\Delta q_{i,t} = \phi_{i,0} + \phi_1(q_{i,t-1} - BX_{i,t-1}) + \Xi \Delta X_{i,t-j} + u_{i,t}$$

Param	Predicted sign	(1)	(2)	(3)	(4)	(5)
ECT	(-)	-0.241***	-0.248***	-0.225***	-0.217***	-0.255***
		(0.031)	(0.031)	(0.033)	(0.033)	(0.036)
$\hat{a}^T$	(-)	-0.485***	-0.452**	-0.349*	-0.358	-0.244
		(0.192)	(0.184)	(0.209)	(0.225)	(0.195)
$\hat{a}^N$	(+)	0.304	0.283	0.559	0.541	0.604
		(0.340)	(0.325)	(0.380)	(0.397)	(0.338)
g	(-)			0.314	0.314	0.116
				(0.366)	(0.382)	(0.332)
g^*	(+)			-0.056	0.029	0.168
				(0.722)	(0.758)	(0.646)
tot	(-)				0.262	
					(0.163)	
$\hat{y}-\hat{n}$	(-)					-0.592**
						(0.278)
p^{oil}	(?)					
Δq	(?)	0.347***	0.306***	0.303***	0.300***	0.319***
		(0.058)	(0.059)	(0.059)	(0.067)	(0.060)
$\Delta\hat{a}^T$	(?)	0.062	0.043	0.116	0.129	0.073
		(0.119)	(0.117)	(0.120)	(0.125)	(0.123)
Δa^N	(?)	-0.382	-0.498***	-0.308	-0.318	-0.352*
		(0.174)	(0.175)	(0.187)	(0.195)	(0.210)
Δg_t	(-)		-0.411***	-0.469***	-0.473***	-0.443***
			(0.144)	(0.150)	(0.153)	(0.154)
Δg^*_t	(+)		0.537***	0.617***	0.636***	0.601***
			(0.209)	(0.222)	(0.236)	(0.223)
Δg_{t-1}				0.460**	0.474**	0.415*
				(0.182)	(0.186)	(0.219)
Δg^*_{t-1}				-0.394	-0.357	-0.351
				(0.253)	(0.259)	(0.278)
Δtot					-0.027	
					(0.061)	
$\Delta(\hat{y}-\hat{n})$						0.072
						(0.266)
Δp^{oil}						
Adj. R^2		.85	.86	0.86	.86	.86
N		271	271	271	266	271
SBC		-5.80	-5.79	-5.74	-5.70	-5.72
t		-6.080***	-6.073***	-6.145***	-5.807***	-6.125***

Notes: OLS standard errors in parentheses. ECT is the coefficient on the coint. vector; the cointegrating vectors are normalized on the real exchange rate. Regressions allow for individual country effects. SBC is the Schwartz Bayesian Information Criterion. "t" is the t-statistic on a regression of the first difference of the error correction term on the lagged error correction term and currency specific dummies; critical values from Levin and Lin (1992). *(**)[***] indicates significance at the 10%(5%)[1%] marginal significance level.

Table 3 (continued)
Panel Estimation Results:
Determinants of the Multilateral Real Exchange Rate

$$\Delta q_{i,t} = \phi_{i,0} + \phi_1(q_{i,t-1} - BX_{i,t-1}) + \Xi\Delta X_{i,t-j} + u_{i,t}$$

Param	Predicted sign	(6)	(7)
ECT	(-)	-0.242***	-0.257***
		(0.034)	(0.042)
$\hat{a}^T$	(-)	-0.354*	-0.505**
		(0.194)	(0.235)
$\hat{a}^N$	(+)	0.515	0.360
		(0.352)	(0.409)
g	(-)	0.376	0.129
		(0.345)	(0.399)
g^*	(+)	0.498	0.681
		(0.736)	(0.837)
tot	(-)		
$\hat{y}-\hat{n}$	(-)		
p^{oil}	(?)	-0.045	-0.033
		(0.025)	(0.029)
Δq	(?)	0.307***	0.298***
		(0.059)	(0.066)
$\Delta\hat{a}^T$	(?)	0.108	0.096
		(0.120)	(0.170)
$\Delta\hat{a}^N$	(?)	-0.345*	-0.331
		(0.188)	(0.231)
Δg_t	(-)	-0.448***	-0.478***
		(0.151)	(0.177)
Δg^*_t	(+)	0.768***	0.976***
		(0.258)	(0.310)
Δg_{t-1}		0.416**	0.525**
		(0.184)	(0.231)
Δg^*_{t-1}		-0.333	-0.504
		(0.255)	(0.314)
Δtot			
$\Delta(\hat{y}-\hat{n})$			
Δp^{oil}		0.001	-0.005
		(0.001)	(0.012)
Adj.R^2		.86	.86
N		271	211
SBC		-5.72	-5.63
t		-5.758***	-5.938***

Notes: OLS standard errors in parentheses. ECT is the coefficient on the cointegrating vector; the cointegrating vectors are normalized on the real exchange rate. Regressions allow for individual country effects. "t" is the t-statistic on a regression of the first difference of the error correction term on the lagged error correction term and currency specific dummies; critical values from Levin and Lin (1992). *(**)[***] indicates significance at the 10%(5%)[1%] marginal significance level.

Table 4
Pedroni Cointegration Test Results

	Homogenous Panel		Heterogeneous Panel		
	(1)	(2)	(3)	(4)	(5)
Constant	--	0.076	--	-0.112 to 0.171	-0.241 to 0.826
z	-0.551	-0.553	4.907 to -2.885	2.414 to -2.673	2.630 to -0.864
time					-0.050 to 0.008
t-stat.	6.072***	6.062***	-6.116*	-6.662	-7.462

Notes: Coefficient estimates from panel regressions (see text). Ranges are provided for heterogeneous panels. time is a linear trend. t-statistic is from the panel DF regression on the residuals from the associated regression. Critical values from Pedroni (1995), Table B.II for columns (1) and (2); from Table B.III for columns (3), (4) and (5).

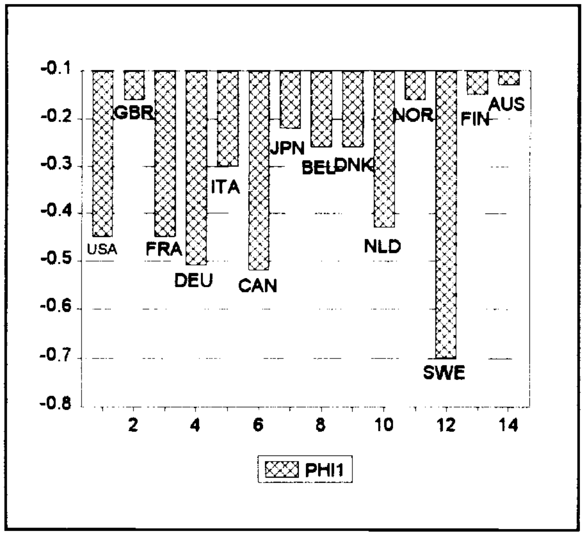

Figure 1: Reversion Parameter

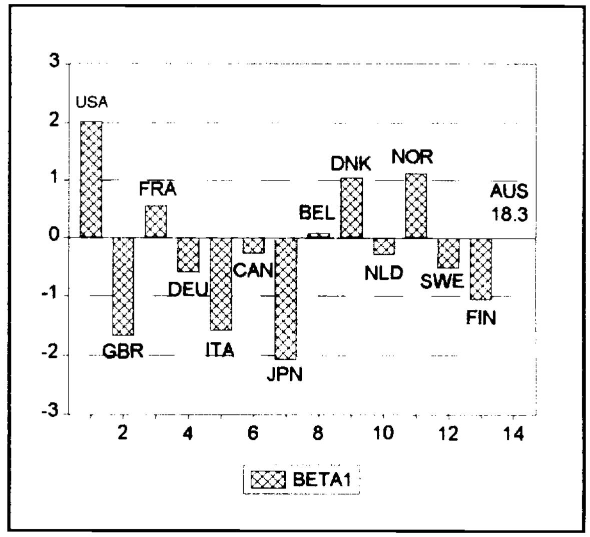

Figure 2: Long-Run Productivity Coefficient

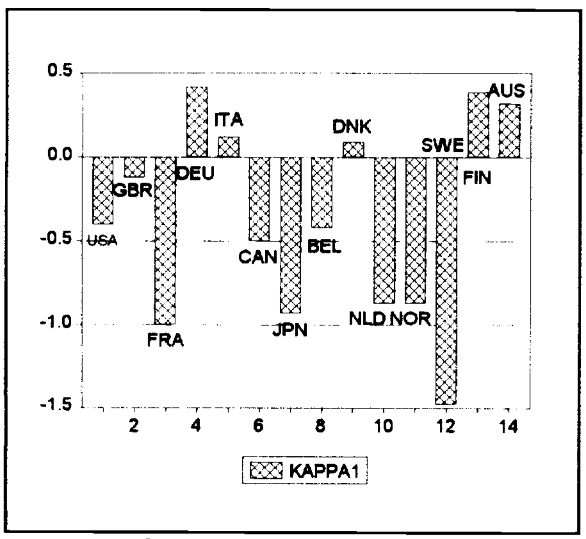

Figure 3: Short-Run Home Government Spending Coefficient

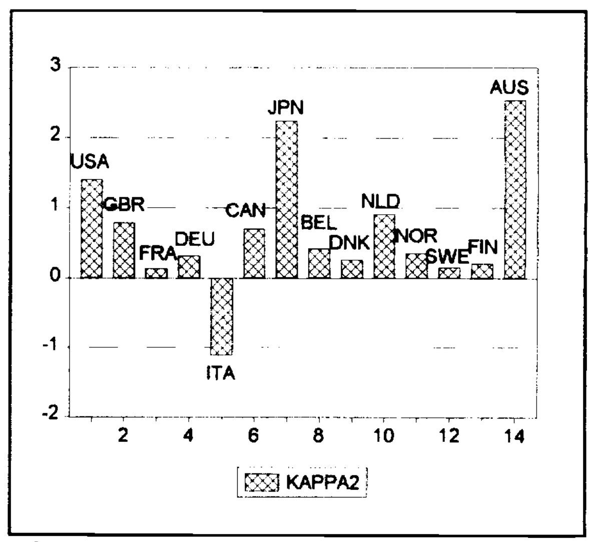

Figure 4: Short-Run RoW Government Spending Coefficient

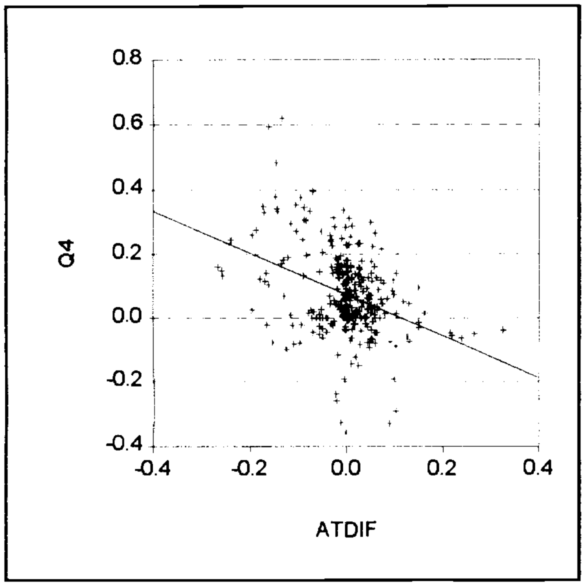

Figure 5: Real Exchange Rate and Tradable Productivity

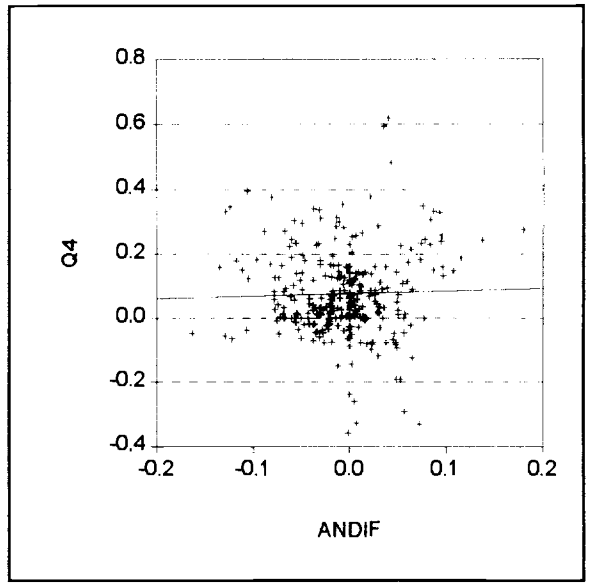

Figure 6: Real Exchange Rate and Nontradable Productivity

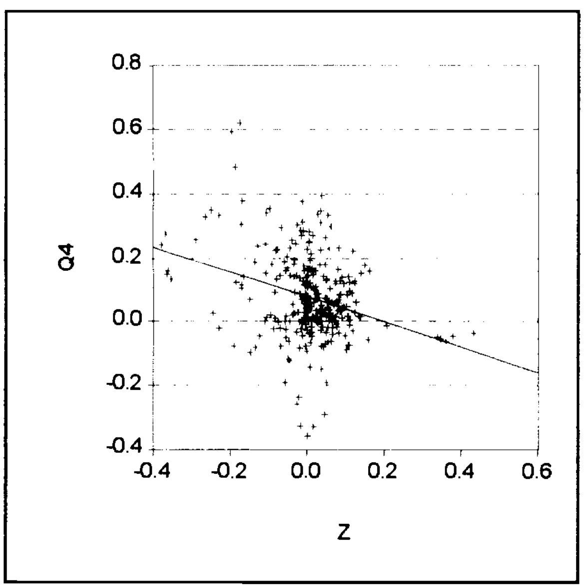

Figure 7: Real Exchange Rate and Relative Sectoral Productivity Differential

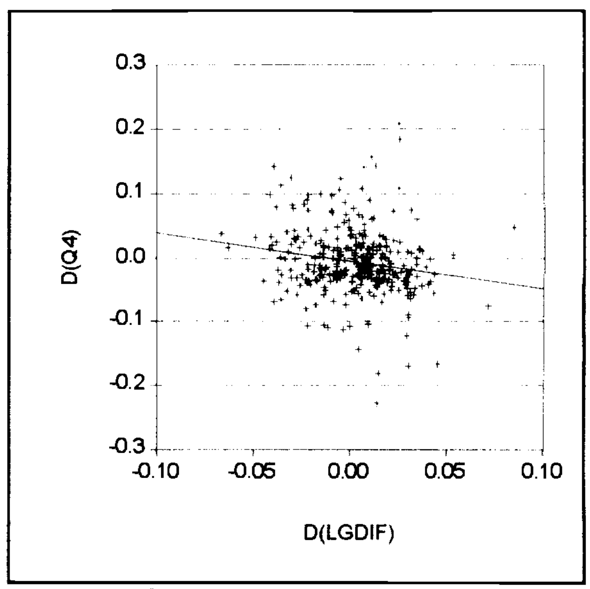

Figure 8: Change in Real Exchange Rate and Change in Government Spending

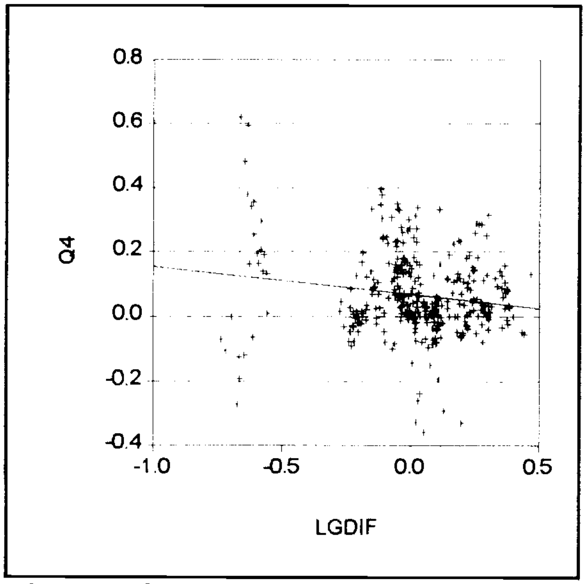

Figure 9: Real Exchange Rate and Government Spending

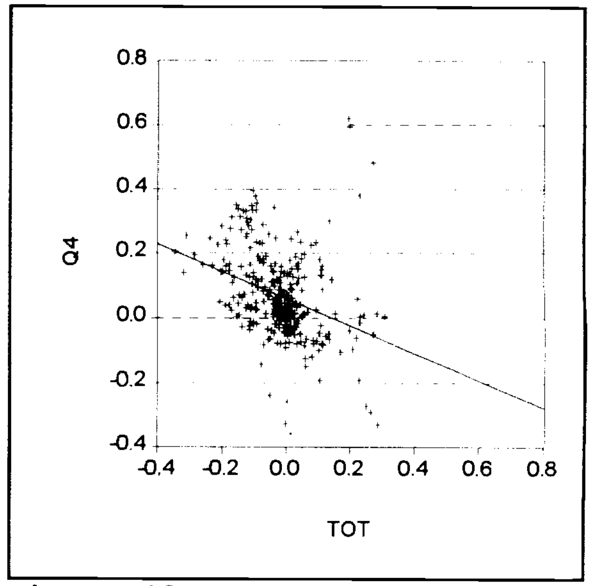

Figure 10: Real Exchange Rate and Terms of Trade

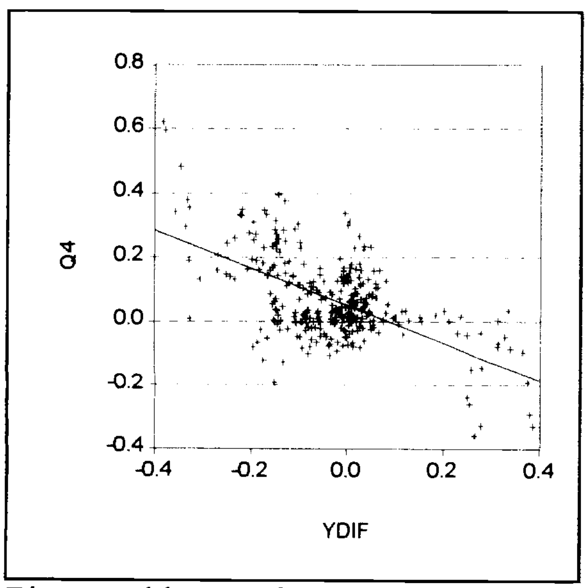

Figure 11: Real Exchange Rate and Per Capita Income (I$)

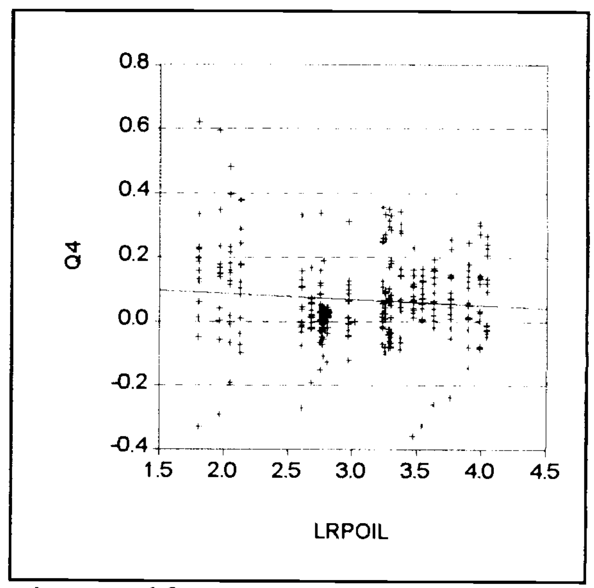

Figure 12: Real Exchange Rate and Real Price of Oil

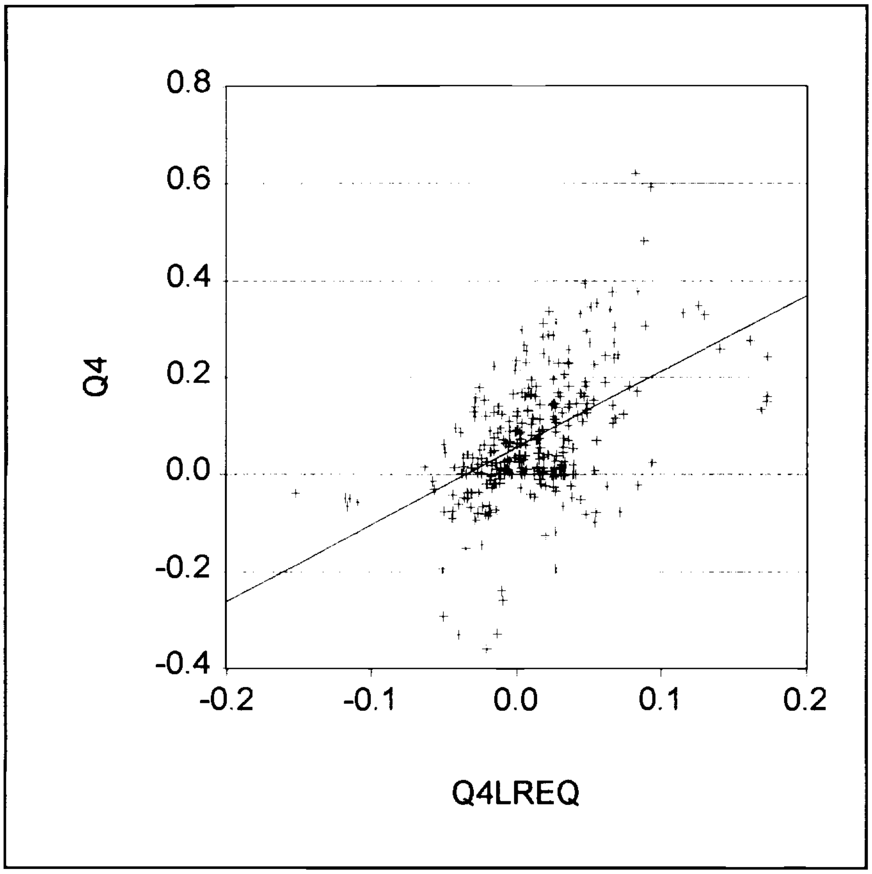

Figure 13: Real Exchange Rate and Predicted Long-Run Rate