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A Cross-Country Panel Analysis Of Currency Substitution And Trade

by

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Abstract

This series evidence has been used to analyse the effect of trade on foreign holdings of domestic currency. Here we extend a model of currency substitution that incorporates a trade motive for foreigners to hold domestic currency. It uses time series and cross-sectional information for a panel of 17 industrialized countries testing two-way fixed effects models against pooled and random-effects alternatives. The cross sectional information is significant, revealing that pooling of data could result in misleading inferences, since country specific effects, regional group effects and distance are all important determinants of domestic currency holding by foreigners.

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I. INTRODUCTION

Early work on currency substitution focused on holdings of the US dollar by non-US residents, examining the limitations that this placed on monetary policy independence and the instability it introduced to the exchange rate [Miles (1978), Bordo and Choudri (1982), Cuddington (1983), Joines (1985)]. The assumption that monetary policy independence would necessarily follow in a floating exchange rate regime was challenged by Miles. Currency substitution on the demand side, resulting from transactionary, precautionary and speculative motives, could give rise to monetary policy dependence and exchange rate instability. Empirical work attempted to determine whether there was evidence for a demand side link through speculative and precautionary motives for holding assets [Mizen and Pentecost (1994)]. This paper continues the search for empirical evidence of a demand side link, but the emphasis is on the transactions motive, through the connection between currency substitution and trade.

The dollar is the most widely used currency outside domestic borders, far outstripping all the other major world currencies (see Table 1). Table 1 reports the use of currencies in official reserves and invoicing patterns by third parties. As a major reserve currency, and an invoicing or ‘vehicle’ currency – both legally, for many commodity trades, and illegally, in black market and criminal transactions - the dollar is an exceptional case. The *habitual* use of the dollar in official reserves, or for the invoicing of certain contracts between third parties, expands the use of the currency in a manner that is not influenced by US trade patterns or monetary policy. An autonomous level of currency holding in dollars exists which does not have the same impact on exchange rate instability or policy independence as genuine currency substitution involving US residents on one side of the transaction. The same can be said, in a more localised way, for other major currencies such as the DM and to a lesser extent the yen, which have similar roles in

Europe and Asia [Tavlas (1997)]. Sterling, however, is different from the dollar, DM and yen in a number of respects. It is not a major reserve currency nor a vehicle currency, but foreign residents' holdings of sterling relative to total money holdings are equivalent, in proportional terms, to the dollar, and growing. Sterling is therefore particularly suitable for an examination of the underlying relationship between currency substitution and trade. Currency substitution and trade can be analysed in the absence of reserve currency and vehicle currency effects, using the readily accessible data on holdings of sterling by non-residents and corresponding bilateral trade.

This paper aims to investigate empirically the determinants of currency substitution for the period 1976-1995. It also seeks to provide new insights in two directions. First, the effects of trade on currency substitution are fully investigated. Despite strong theoretical and empirical support [Grassman (1973); de Vries (1988)] for a transactions motive for foreigners to hold domestic currency, there are few papers that make the explicit link to trade. The notable exceptions are Ratti and Jeong (1994) and Milner, Mizen and Pentecost (1996), which offer some empirical evidence on this transactional approach. Second, in this paper we supplement the existing time series evidence by using a panel of data on domestic currency holdings by non-residents[1]. In this paper, both time series *and* cross sectional information is derived from quarterly data for the post- Bretton Woods regime (1976-95) in a panel of 17 industrial countries. This allows time and fixed (country) effects to be carefully investigated, and provides greater precision for our estimates than existing time series methods.

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[1] The use of panel data methods to supplement time series models has been popularized in recent international research by Chinn (1996) and Husted and MacDonald (1996).

The remainder of the paper is organized as follows. In the next Section we set up a model of foreigners' holdings of domestic currency. Section 3 describes the data and empirical methodology and presents the results of the panel estimation and their implications. The conclusions are reported in Section 4.

II. THE MODEL

In this section we use a model similar to that of Bergstrand and Bundt (1990). It is a two-stage model, since it separates the decision to hold financial wealth in the form of liquid or illiquid assets from the decision over the currency of denomination of the assets in each category. Taking the first decision as given, Berstrand and Bundt (1990) consider the real supply of domestic currency, m^D, that is held by domestic residents, m^{DD}, and foreign residents, m^{DF}. The total demand for domestic real money balances depends on the level of transactions by foreigners and the domestic and foreign rates and interest, r_t^D and r_t^F in the following way:

$$(1) \qquad m_t^{DF} = K^{DF}(r_t^D, r_t^F)Q_t^F$$

Q_t^F represents combined transactions effects due to foreign income and gross trade. Thus the model of foreigner's holdings of domestic currency to be estimated is as follows:

$$(2) \qquad \ln\left(m_{it}^{DF}\right) = \alpha_0 + \alpha_1 r_{it}^D + \alpha_2 r_{it}^F + \alpha_3 \ln\left(y^F{}_{it}\right) + \alpha_4 \ln\left(T_{it}\right) + \varepsilon_{it}$$

where $\ln(m_{it}^{DF})$ is the logarithm of real domestic money holdings by foreigners in country i, r_{it}^D and r_{it}^F are domestic and foreign bond yields, $\ln(y_{it}^F)$ is logarithm of foreign income, $\ln(T_{it})$ is the logarithm of gross bilateral trade and ε_{it} is a random error term. The expected signs are as follows: $\alpha_1 < 0$, $\alpha_2 > 0$, $\alpha_3 > 0$, but α_4 is ambiguous. The term r_{it}^{D} is the opportunity cost of holding domestic money and so, as it falls, foreigners will switch out of foreign money into domestic money, because although they cannot make an acquisition of bonds to reduce the total holdings of money (decided upon in the first stage), they can re-allocate their holdings between domestic and foreign monies. Conversely, if r_{it}^{F} falls (and r_{it}^{D} remains unchanged) then there is a lower opportunity cost of holding foreign currency and a tendency to switch from domestic to foreign money. The variable $\ln(y_{it}^{F})$ represents the scale of transactions, and therefore, the expected sign of this coefficient is positive. Finally, we separately include the scale of international transactions in the form of a gross trade variable. The effects of cross country variation in the scale of trade on domestic currency holdings held by foreigners is likely to be ambiguous, since trade dependence is likely to be affected by the size of a country; with larger countries having less 'dependence' on sterling balances. The volume of trade typically does not increase in proportion to country size and large countries are more likely to have currencies that are used and accepted by others. Thus there may even be a negative relationship between domestic currency holdings of foreigners and the level of trade in cross section if the use of the currency by others dominates the transactions effect.

In order to avoid the potential co-linearity problems from the inclusion of $r_{it}{}^{D}$ and $r_{it}{}^{F}$ separately, we form a differential term $(r_{it}{}^{D} - r_{it}{}^{F})$ which we expect to have a coefficient with a negative sign if currency substitution occurs. We estimate the model in dynamic equilibrium correction form, in order to allow the identification of both short run and long run relationships[2]. Thus:

$$(3)\qquad \begin{aligned} \Delta ln(m_{it}^{DF}) = {} & \beta_0 + \beta_1 \Delta\left(r_{it}^{D} - r_{it}^{F}\right) + \beta_2 \Delta ln(y_{it}^{F}) + \beta_3 \Delta ln(T_{it}) - \gamma_0\, ln(m_{it-1}^{DF}) \\ & + \gamma_1\left(r_{it-1}^{D} - r_{it-1}^{F}\right) + \gamma_2\, ln(y_{it-1}^{F}) + \gamma_3\, ln(T_{it-1}) + \omega_{it} \end{aligned}$$

where Δ is the first difference operator. In long-run equilibrium where all current and lagged values are at their equilibrium values, equation (3) reduces to:-

$$(4)\qquad \ln(m_{it}^{DF}) = \frac{\beta_0}{\gamma_0} + \frac{\gamma_1}{\gamma_0}\left(r_{it}^{D} - r_{it}^{F}\right) + \frac{\gamma_2}{\gamma_0}\ln(y_{it}^{F}) + \frac{\gamma_3}{\gamma_0}\ln(T_{it}) + u_{it}$$

provided $\gamma_0 \neq 0$. The expected signs are $\frac{\gamma_1}{\gamma_0} < 0$ and $\frac{\gamma_2}{\gamma_0} > 0$, but the last term is ambiguous from theory.

III. PANEL ESTIMATION AND EVIDENCE FOR OECD COUNTRIES

The estimation techniques which are used for the analysis of the data are standard one- and two-way fixed and random effects linear models for panel data. The one-way random effects model for example is of the form:-

$$(5)\qquad y_{it} = \alpha + \beta' x_{it} + \varepsilon_{it} + u_i$$

where ε_{it} is an idiosynchratic shock and u_i is a random effect. The model imposes the condition that $E[u_i] = 0; Var[u_i] = \delta_u^2$; and $Cov[\varepsilon_{it}, u_i] = 0$.

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[2] The variables used in equation (3) are examined individually for stationarity using the Levin and Lin (1992) procedure to test for unit roots in a panel context in the next section. None of the variables had a unit root on the basis of the tabulated critical values calculated by Levin and Lin.

The two-way fixed effects by contrast assumes that there are systematic differences across groups (countries) captured by country-specific constants, and allows for time specific effects also in the following general way:

(6) $$y_{it} = \alpha_0 + x_i + \phi_t + \beta' x_{it} + \varepsilon_{it}$$

This model has an overall constant as well as a group effect for each element in the group (x_i) and a time effect for each period (ϕ_t) [3]. All of the results were estimated using LIMDEP 7.0.

The panel includes data for 17 industrialized countries over the time period 1976 to 1995 sampled at quarterly intervals. The countries include the USA and Canada, the principal European countries (Austria, Belgium, Denmark, Finland, France, Germany, Italy, Netherlands, Norway, Republic of Ireland, Spain, Sweden, Switzerland) and Japan and Australia. The dependent variable in the regression is defined as non-bank, foreign residents' holdings of real money (sterling) balances held in UK banks as recorded by the Bank of England. These balances were translated into local currency values by deflating by the end of period spot exchange rate and then converted into constant prices by deflating by the appropriate consumer price index. The explanatory variables include local real income as a scale variable, measured as GDP at constant prices.[4] The measure of trade reported in the results is gross trade measured as the sum of export plus import values in constant local currency prices[5]. The opportunity cost of holding

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[3] The problem of multicolinearity is overcome by imposing the restriction that $\sum_u \lambda_i = \sum_t \phi_i = 0$.

[4] For Belgium, the Netherlands and the Republic of Ireland it was necessary to interpolate annual GDP figures into quarterly series by use of the quarterly industrial production index. See Bruggeman (1996) for a detailed discussion of this methodology.

[5] Alternative measures based on foreign imports (UK exports) in constant local currency prices and the Grubel-Lloyd index of intra-industry trade were also investigated.

money balances are represented by relative short-term interest rates, measured by the respective three-month Treasury Bill rates.[6]

The basic fixed effects model, allowing for intercept heterogeneity, which was used to test the model of domestic currency holding and currency substitution is set out in column 1 of Table 2.[7] In order to investigate the effects of possible co-linearity between the income variables $\ln(y_{it}^F)$ and the trade volume variables $\ln(T_{it})$, we report the sensitivity of estimated elasticities to the effects of leaving each of these variables out in turn. Although there is a change in sign on $\Delta ln(y_{it}^F)$ between columns 1 and 3, the variable is not significant in each case and the parameter estimates and significance levels of the other variables remain acceptably stable across the different specifications. In two out of three models in Table 2, however, the income variable is insignificant when the trade variable is significant, which may point to mild co-linearity between these two variables. The main results of the paper focus on the fully specified model in column 1 of Table 2.

The short-run coefficients (β_i) describe the short-run dynamics of portfolio adjustment and are not determined *a priori* by our model in terms of their magnitude or sign. The short term response coefficients on the interest rate differential and income terms reported in Table 2 are not

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[6] Stationarity tests proposed by Levin and Lin (1992) based on the regression $y_{it} = \alpha_0 + \rho y_{it-1} + \phi_t + \phi_i t + \varepsilon_{it}$ using de-meaned series, indicated that the variables were stationary. The respective t-statistics for a null that $\rho = 1$ are real money balances t = -10.632; real income t = -25.661; relative interest rates t = -12.820; gross trade t = -14.623. Since the variables are stationary in levels inferences can be drawn from classical econometric analysis without the need to difference to stationarity or to identify cointegrating relationships, Pedroni (1997a,b).

[7] Alternative panel specifications for the standard model (equation 3) were initially estimated. The results showed that the fixed effects models were preferrred to pooled and random effects alternatives under both one-way and two way specifications. The results are available from the authors on request.

significantly different from zero. However, the highly significant positive coefficient on the trade variable ($\Delta ln(T_{it})$) merits attention. This result is consistent with a transactions motive: an increase in trade induces a short-term, temporary increase in holdings of domestic currency by foreigners. This interpretation is not sensitive to the measure of trade expansion nor the currency that importers and exporters invoice in. If the growth in trade involves growing imports from the UK that need to be paid for in domestic currency, then the coefficient on the variable is consistent with a short-run, positive impact on the demand for active liquidity services. Alternatively, if there are growing exports to the UK which are invoiced and paid for in domestic currency, then temporary accumulation of domestic currency may reflect sluggishness in adjusting the asset composition of portfolios.[8] Estimation of the model using trade measured by the foreign currency value of imports and the Grubel-Lloyd index of two-way trade show exactly the same pattern of positive short-term changes in foreigners' domestic currency balances in response to increased trade.[9]

The long-run relationships implied by the estimated model in Table 2 column 1 between domestic currency holdings and foreign income, $ln(y_{it}^F)$, and the interest rate differential (r_{it}^D - r_{it}^F) confirm the hypothesized relationships derived from theory. The coefficients γ_0 and γ_2 are negative and significant at the 1% level, while γ_1 is positive and significant at the 10% level. The long-run elasticity of foreigners' domestic currency demand with respect to income (as summarized in Table 3) is positively income-elastic in line with a domestic transactions/wealth effect explanation. The long-run interest rate effect is less robust but a rise in foreign interest rates

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[8] Expectations of exchange rate adjustments may also affect the speed of adjustment of domestic currency holdings by foreign residents.

[9] The share of matched or two-way trade in gross trade.

relative to domestic rates induces foreigners to substitute domestic cash balances for their own currency, implying that currency substitution occurs. This effect is much more robustly identified when, in Table 4, the interest rate differential is adjusted for exchange rate changes (r_{it}^{D} - r_{it}^{F}- Δe_{it}). In this case β_1(ie the short run effect) is also negative and significant at the 1% level. Reassuringly, the coefficients on the other variables in both the dynamic and long-run form of the estimated model remain fairly stable. Indeed the specific values of the long-run coefficients in both the 'core' model (column 2, Table 3) and the 'alternative' specification (column 2, Table 4) give very similar income and trade elasticites for domestic currency holdings by foreigners.

The long-run relationship between foreigners' domestic currency holdings and gross trade differs from the short run response (see Table 3). While foreigners' accumulate domestic currency in the short-run $\left(\beta_3 > 0\right)$, in the long-run they decumulate domestic currency as trade increases $\left(\frac{\gamma_3}{\gamma_0} < 0\right)$.[10] In the long term there can still be a motive for holding domestic currency by foreign residents, but the size of those holdings may decrease with the increased expectations of matched trade in equilibrium. We may need to consider also the effects of country size both on trade volume and on the need to hold other currencies. Although the UK's trade will tend to be greater in absolute terms with larger countries, large countries tend to be relatively less open than small ones and their own currencies are more likely to be used as "vehicles" by other countries. They therefore have a reduced incentive to hold "domestic currency" than small countries for transactions and precautionary purposes in the longer term. In the short term, however, both large and small countries may expect bilateral trade imbalances and their residents may be sluggish in adjusting portfolios in the face of imbalances.

[10] This result is robust for measures of trade based on the foreign currency value of imports and the Grubel-Lloyd

It is interesting to note that this distinction between the short-run and long-run effect of trade would have been obscured by pooled data. Consider Figure 1 which gives a scatter plot of the domestic currency balances held by foreign residents and the trade volumes for each country. There is a strong positive relationship evident between the volume of trade and domestic currency holdings of foreigners in the pooled data. Moreover, this positive relationship is preserved when domestic currency balances are plotted against lagged trade, to allow for a typical 60-day settlement period for invoices, which could lead *current* transactions to be recorded as currency flows in the *subsequent* quarter. When the data plots are examined by region and country in Figure 2, however, the positive association holds for only a few countries. The positive pattern in the data as a whole is the result of the *combination* of data points from different countries that do not reveal a systematic positive relationship when assessed individually.[11] Clearly a model which pooled the present data would conclude, incorrectly as it turns out, that there is a positive long-run relationship between trade and currency holdings. Using a pooled regression method masks the country-specific information in the data and should be rejected in favor of a model which can capture individual country (group) effects.[12]

We report country fixed effects for our 'core' model in Table 3. The short-run country effects are expressed as deviations (λ_i) from the average group value of the constant term (β_0), while the implied long-run values are recorded as country constant terms $\frac{\beta_0 + \lambda_i}{\gamma_0}$. If we focus first of all

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index shows that the result is robust.

[11] Time series evidence would also give mixed results, depending on the particular country considered.

[12] This reveals why both pooled and random effects models were comprehensively rejected in favor of the model which allows two-way group and time effects.

on short-run effects, we can see that large positive values for the λ_i apply to Italy, Japan, Spain and to a lesser extent USA. By contrast negative values are associated with Ireland, Finland, Norway, Netherlands, Switzerland, Australia, Denmark, Canada, Sweden and Austria. Since the model is predicting logarithmic values of domestic currency balances held by foreigners, all of the long-run values of the country constants (which are negative) become more negative as the absolute value of any *positive* fixed domestic currency holdings approaches zero. In other words larger negative values for the long-run country constant terms imply smaller but positive autonomous domestic currency holdings. Thus Italy, Japan, Spain and the USA are the countries with smaller (unexplained) fixed holdings of domestic currency, while Ireland, the British Commonwealth and smaller north European countries have larger fixed holdings of domestic currency. This latter set of countries includes ones that we would expect to have larger holdings than that simply accounted for by income and trade levels and the degree of financial integration, given their geographic or cultural proximity.

To account for the possibility that country type may be important we re-estimate the model including country-type effects in place of individual-country effects. These measure the effects of collaboration between countries in trade areas such as the European Union, the European Free Trade Association and the British Commonwealth on domestic currency balances. The fixed effects are reported in Table 5 in columns 1 and 2. The results reveal that participation in EFTA reduced domestic currency balances significantly, but membership of the EU and the Commonwealth had no significant effect on foreigners' holdings of domestic currency. The negative effect of EFTA membership within this sample probably reflects the relatively high income and trade levels of these Scandinavian countries with the UK.

The effect of distance on trade through a gravity-type relationship is included in this model by letting foreigners' domestic currency balances vary non-linearly with distance.[13] The results in column 3 of Table 5 show that there is a significant negative relationship in distance, as expected, and a positive relationship in distance squared which is on the borders of significance at the 10% level; increased distance from the UK reducing domestic currency holdings, but at a decreasing rate. This may capture both a direct effect associated with reduced financial integration and cultural proximity as distance increases. It may also capture an indirect effect of reduced gravitational pull on trade flows and a resulting reduction in transactions demand for the domestic currency. Interestingly, once we control for distance the positive coefficient on the Commonwealth membership dummy is also (weakly) significant.

IV. CONCLUSIONS

Substitution of currencies on the demand side can result in greater monetary policy interdependence and exchange rate instability even under floating exchange rates: it is this possibility that has spurred on the search for evidence of currency substitution in practice. Most empirical work has focused on the demand for currency as an alternative asset, invoking speculative and precautionary motives. Comparatively little work has been done previously on the relationship between currency substitution and trade. This paper considers this issue.

By forming a panel of data from 17 OECD countries over the period 1976-95 the precision of the estimates from the available span of data are increased. The findings reported in the results of the panel estimations are encouraging and informative. The effect of income and relative yields have

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[13] This relationship has been explored in recent papers by Bergstrand (1985, 1991) and Harrigan (1993).

comparable effects on currency substitution as those recorded by time series results, but this paper provides new evidence of a transactions effect in the short-term through the impact of trade on currency substitution. Holdings of domestic currency by foreigners are clearly influenced by changing patterns of trade in the short run. In the long term, geographical proximity becomes a more important influence on the use of currency by non-residents, as we might expect. This information is brought to light in cross-sectional analysis of panel data but is easily obscured by pooling the data. The country-specific features of the impact of trade on currency substitution would have been less evident in time series regressions on a country-by-country basis. We conclude that cross-sectional panel analysis has an important contribution to make when assessing the relationship between currency substitution and trade.

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Table 1: International Use of the Dollar, DM, Sterling and the Yen

Non-Bank Private Sector Holdings of Foreign Currency ($bn)

Currency	End 1985	End 1989	End 1993	End 1996
US dollar	201.1	264.0	318.9	432.2
DM	16.8	84.5	199.7	201.1
Yen	4.5	14.8	17.8	48.8
Sterling	4.9	17.1	31.5	39.7
All Currencies	250.4	457.7	735.9	976.8

Non-Bank Private Sector Holdings of Foreign Currency
as a Proportion of Total Money Stock (%)

Currency	End 1985	End 1989	End 1993	End 1996
US dollar	5.25	5.39	6.17	7.13
DM	5.15	12.3	16.2	13.0
Yen	0.3	0.4	0.4	0.9
Sterling	2.82	6.58	8.71	9.10

Official Monetary Institutions Holdings of Foreign Currency ($bn)

Currency	End 1985	End 1989	End 1993	End 1996
US dollar	60.4	71.5	92.3	126.1
DM	20.0	39.3	21.4	44.6
Yen	5.5	14.6	12.6	9.9
Sterling	0.5	1.3	5.4	3.7

Currency Denomination of Exports, Selected Industrial Countries 1992-96 (%)

	US dollar	DM	Yen	Sterling
United States	98.0	0.4	0.4	0.3
Germany	9.8	76.4	0.6	2.4
Japan	52.7	0.0	35.7	0.0
United Kingdom	22.0	5.0	0.7	62.0

Sources: BIS International Banking and Finance Market Developments, various issues
IMF, International Finance Statistics, various issues
Tavlas (1997) Table 9.

Table 2: Alternative Fixed Effect Models

	1*	2*	3*
Constant	0.934	-0.062	0.854
	(2.890)	(-1.679)	(2.523)
$D(r_{it}^{D} - r_{it}^{F})$	0.001	(0.002) (0.864)	0.001
	(0.562)		(0.537)
$Dln(y_{it}^{F})$	-0.054		0.023
	(-0.491)		(0.200)
$Dln(T_{it})$	0.567	0.572	
	(11.736)	(11.818)	
$ln(m_{it-1})$	-0.187	-0.171	-0.120
	(11.499)	(-11.065)	(-9.737)
$(r_{it-1}^{D} - r_{it-1}^{F})$	0.002	0.002	0.001
	(1.320)	(1.806)	(0.689)
$ln(y^{F}_{it-1})$	-0.366	-	-0.242
	(-3.099)		(-1.994)
$ln(T_{it-1})$	0.150	0.134	-
	(7.326)	(6.748)	
Group effects	fixed	fixed	fixed
Time effects	yes	yes	yes
R^2	0.269	0.262	0.165
(adjusted) R^2	(0.208)	(0.201)	(0.097)
F statistic	4.410	4.348	2.428
	(98.12)	(96.18)	(96.18)
log-likelihood	1433.27	1427.11	1348.86

* *t* statistics in brackets

Table 3: Short and Long run Coefficients and Country Fixed Effects*

Explanatory Variable	Short-run	Long-run
$\Delta(r_{it}^{D} - r_{it}^{F})$	$\beta_1 = 0.001$	
$\Delta ln(y_{it}^{F})$	$\beta_2 = -0.054$	
$\Delta ln(T_{it})$	$\beta_3 = 0.567$	
$ln(m_{it-1})$	$\gamma_0 = -0.187$	
$(r_{it-1}^{D} - r_{it-1}^{F})$	$\gamma_1 = 0.002$	$\gamma_1/\gamma_0 = -0.011$
$ln(y^{F}_{it-1})$	$\gamma_2 = -0.366$	$\gamma_2/\gamma_0 = 1.957$
$ln(T_{it-1})$	$\gamma_3 = 0.150$	$\gamma_3/\gamma_0 = -0.802$
constant	$\beta_0 = 0.934$	
+ country 1 (Australia)	$\beta_0 + \lambda_i = -0.265$	$\frac{(\beta_0 + \lambda)}{\gamma_0} = -3.578$
+ country 2 (Austria)	= -0.103	= -4.444
+ country 3 (Belgium)	= 0.104	= -5.551
+ country 4 (Canada)	= -0.187	= -3.995
+ country 5 (Denmark)	= -0.251	= -3.652
+ country 6 (Finland)	= -0.394	= -2.888
+ country 7 (France)	= 0.008	= -5.037
+ country 8 (Germany)	= 0.007	= -5.032
+ country 9 (Italy)	= 0.987	= -10.273
+ country 10 (Japan)	= 0.762	= -9.070
+ country 11 (Netherlands)	= -0.265	= -3.578
+ country 12 (Norway)	= -0.275	= 3.524
+ country 13 (Rep. of Ireland)	= -0.676	= -1.380
+ country 14 (Spain)	= 0.584	= -8.118
+ country 15 (Sweden)	= -0.162	= -4.128
+ country 16 (Switzerland)	= -0.267	= -3.567
+ country 17 (USA)	= 0.393	= -7.096

* Based on equation 1 in Table 2, with all country effects significant at 1% level

Table 4: Alternative Dynamic and Long run Estimated Model

	Dynamic Model *	**Long run Model***
Constant	0.809	-4.731
	(2.684)	
$\Delta(r_{it}^{D} - r_{it}^{F} - \Delta e_{it})$	-0.004	
	(-5.897)	
$\Delta ln(y_{it}^{F})$	-0.057	
	(-0.556)	
$\Delta ln(T_{it})$	0.225	
	(4.537)	
$ln(m_{it-1})$	-0.171	
	(-11.390)	
$(r_{it-1}^{D} - r_{it-1}^{F} - \Delta e_{it})$	0.007	-0.041
	(7.816)	
$ln(y^{F}_{it-1})$	-0.326	1.906
	(-2.958)	
$ln(T_{it-1})$	0.145	-0.848
	(7.650)	
Group effects	fixed	fixed
Time effects	yes	yes
R^2	0.392	
F statistic	7.715	
	(97.12)	
log-likelihood	1530.64	

* *t* statistics in brackets

Table 5: Country-Type and Gravity Effects

	1*	**2***	**3***
$D(r_{it}^{D} - r_{it}^{F})$	0.0002 (0.170)	0.0002 (0.183)	0.0004 (0.290)
$Dln(y_{it}^{F})$	-0.086 (-0.943)	-0.086 (-0.948)	-0.078 (-0.865)
$Dln(T_{it})$	0.447 (10.791)	0.446 (10.761)	0.445 (10.730)
$ln(m_{it-1})$	-0.033 (-5.313)	-0.031 (-5.185)	-0.035 (-5.593)
$(r_{it-1}^{D} - r_{it-1}^{F})$	0.002 (2.770)	0.002 (2.677)	0.002 (3.087)
$ln(y^{F}_{it-1})$	-0.004 (-0.881)	-0.002 (-0.487)	0.006 (1.169)
$ln(T_{it-1})$	0.030 (3.810)	0.026 (3.709)	0.025 (3.665)
EU1	-0.007 (-0.894)	-	-
EU2	-	-0.0008 (-0.128)	-0.0008 (1.081)
COMMONWEALTH	0.005 (0.676)	0.006 (0.721)	0.020 (1.740)
EFTA	-0.015 (-2.191)	-0.011 (-2.031)	-0.013 (-2.222)
DISTANCE	-	-	-0.105 (-2.033)
DISTANCE SQUARED	-	-	0.073 (1.777)
R^2	0.111	0.110	0.113
F statistic	17.49 (9, 1265)	17.40 (9, 1264)	14.67 (11, 1263)
log-likelihood	1309.12	1307.62	1309.87

* *t* statistics in brackets