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**The Real Exchange Rate in Small Open Developed Economies:  
Evidence from Cointegration Analysis**

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## **Abstract**

We examine the effects of the terms of trade and the expected real interest rate differential on the real exchange rate in a sample of small open developed economies. We employ cointegration analysis to search for possible long-term linkages. We find that while both the terms of trade and the expected real interest rate differentials affect the real exchange rate in the long run, the role of the terms of trade generally proves more consistent across countries. The speed of adjustment for the expected real interest rate differential in the error-correction model, however, is quantitatively larger than it is for the terms of trade.

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## **I. Introduction**

Few issues in international finance attract more attention than the determination of the real exchange rate. Yet this issue remains open; little consensus exists on the appropriate set of fundamental factors that explain the real exchange rate (e.g., Baxter 1994, Chinn 1991, Coughlin and Koedijk 1990, MacDonald 1998, and Mussa 1990). Different studies consider different fundamental factors. Moreover, sometimes the theoretical work does not connect with empirical practice. For example, while theoretical work investigates extensively the relationship between terms-of-trade shocks and the real exchange rate, empirical work on large developed countries generally overlooks the role of the terms of trade in determining the real exchange rate.

The existing empirical literature generally considers either large developed or small developing countries. That literature explores the effect of the expected real interest rate differentials (both short-term and long-term) on real exchange rates, usually for large countries, and generally finds little evidence of a long-run relationship. In large economies, however, other fundamental factors (e.g., domestic and foreign productivity, capital accumulation, wealth, saving, and so on) may determine the real exchange rate. On the other hand, the existing research on real exchange rates for small economies generally considers developing countries, where fundamental differences in capital mobility and trade barriers exist.

We examine whether the real interest rate differential and the terms of trade possess a long-run relationship with the real exchange rates (against the US dollar) in nine small, developed economies (Australia, Austria, Canada, Italy, Finland, New Zealand, Norway, Portugal, and Spain).<sup>1</sup> Each country incorporates a relatively high degree of openness in both goods and financial markets. Although researchers propose numerous possible determinants, we confine our analysis to the long-term expected real interest rate differential and the terms of trade as the major proximate exogenous long-run

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<sup>1</sup> We do not choose other small European countries because the movements of their nominal exchange rates were restricted within the 2.25 band of the European Monetary System (EMS) after 1979. We consider Italy, however, because its currency moved within a wider band and it realigned its exchange rate more frequently than other members of the EMS.

determinants of the real exchange rate.<sup>2</sup> In those theoretical models that maintain a monetary spirit but incorporate rigidities, a currency's appreciation inversely relates to the real interest rate differential (between foreign and home real interest rates) (Meese and Rogoff 1988, Frankel 1979, Mussa 1982, and Obstfeld and Rogoff 1996). The failure to find strong empirical support for this relationship prompts researchers to seek other variables that along with the interest rate differential cointegrate with the real exchange rate.<sup>3</sup> In this paper, we find cointegration between the real exchange rate, the real interest rate differential and the terms of trade. The expected real interest rate differential captures financial market developments, especially capital flows, and the terms of trade captures goods market developments.

Although some researchers (Gruen and Wilkinson 1994, and Amano and van Norden 1995) investigate the relationship between the real exchange rate, the terms of trade, and the expected real interest rate differential in the context of small developed economies (i.e., Australia and Canada, respectively), our research differs qualitatively. By examining a group of small developed economies, we can determine if any of the findings of Gruen and Wilkinson (1994) or Amano and van Norden (1995) represent general rather than idiosyncratic results.

Since we find that the real exchange rate cointegrates with the long-term interest rate differential and the terms of trade, the findings of Gruen and Wilkinson (1994) and Amano and van Norden (1995) generalize to our nine-country sample. Moreover, the terms of trade more consistently and more strongly affect the real exchange rate than the expected real interest rate differential. That result contrasts with Gruen and Wilkinson's (1994) finding that the long-term interest rate differentials prove quantitatively more important than the terms of trade. Finally, although the real exchange rate cointegrates with the real

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<sup>2</sup> That we find cointegration between those three variables validates our narrow focus and allows the estimation of the error-correction model that encompasses only those three variables. Further research can expand this list of potential long-run determinants. Our current investigation provides consistent evaluation of the potential role of the terms of trade and the expected real interest rate differential across a relatively homogeneous group of small open developed economies.

<sup>3</sup> For example, Blundell-Wignall and Brown (1991) and Edison and Pauls (1993) consider the cumulated current account balance in their analysis. While Blundell-Wignall and Brown succeed in their attempt, Edison and Pauls do not.

interest rate differential and the terms of trade, the magnitude and signs of such effects differ across countries.

## **II. Literature Review**

Empirical research on exchange rates in the 1970s and 1980s focuses largely on the short-run movements of exchange rates reflecting the increased exchange rate volatility after the abandonment of the Bretton-Woods system and the failure of asset models to provide an adequate explanation of exchange rate changes. More recently, however, some focus shifts to real variables and the long-run adjustment of real exchange rates.

A number of studies consider whether real interest rate differentials explain real exchange rate movements. Meese and Rogoff (1988), Edison and Pauls (1993), and Coughlin and Coedijk (1990) do not find a cointegrating relationship between real exchange rates and expected real interest rate differentials. These three papers use the Engle-Granger cointegration method.<sup>4</sup> Huizinga (1987), after decomposing real exchange rate movements into permanent and transitory components, shows that the transitory component of the real exchange rate accounts for only a small portion of actual real exchange rate variation. Moreover, Campbell and Clarida (1987) conclude that movements in the ex ante real interest rate differential do not offer large or persistent enough change to account for the variability in the real exchange rates.

The failure of expected real interest rate differentials alone to explain real exchange rate movements prompted researchers to expand the set of explanatory variables. The empirical literature now considers domestic and foreign productivity proxies (Edwards 1989, Meese 1990, Huizinga 1987, Coughlin and Koedijk 1990, Faruquee 1995, Stein 1994, Strauss 1996, Williamson 1994, and Zhou 1995), capital accumulation (Edwards 1989, Stein 1994, and Williamson 1994), cumulated current account balances (Edison and Pauls 1993, and Coughlin and Koedijk 1990), the level and composition of

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<sup>4</sup> Chinn (1991) argues that while the fundamentals are appropriate, the functional form is not.

government spending (Edwards 1989, Meltzer 1993, and Zhou 1995), saving (Stein 1994), and the terms of trade (Faruquee 1995, Gruen and Wilkinson 1994, Edwards 1989, and Amano and van Norden 1995).<sup>5</sup>

Empirical analysis of real exchange rates usually perform better at longer horizons, making techniques, such as Johansen's (1988) cointegration analysis, relevant for examining long-run relationships between real exchange rates, expected real interest rate differentials, and other explanatory variables. Amano and van Norden (1995) find that the Canadian-U.S. real exchange rate depends on movements in the terms of trade, and that the influence of monetary factors, as reflected in expected real interest rate differentials, is only secondary. Gruen and Wilkinson (1994), also using cointegration methods, consider Australia's trade-weighted real exchange rate and find that both the terms of trade and the real interest rate differential explain the real exchange rate during the period of floating exchange rates. Unlike Amano and van Norden (1995), Gruen and Wilkinson (1994) find that the real interest rate differential is qualitatively more important than the terms of trade in explaining the real exchange rate. Hansen and Hutchison (1996) consider the terms of trade along with the supply of nontraded goods as possible "real" determinants of the nominal exchange rate. They use an error-correction model to examine New Zealand data for the period 1979-1993 and find that a long-run equilibrium relationship exists among these variables.<sup>6</sup> MacDonald (1998) examines the long-term determinants of real effective exchange rates for Japan, Germany, and the U.S., reversing much of the findings in the previous literature and finding evidence of cointegration between the real exchange rate and real interest rate differentials as well as other determinants – the terms of trade, productivity differentials, relative fiscal balances, and net foreign assets.<sup>7</sup>

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<sup>5</sup> The terms of trade and the expected real interest rate differentials appear together in a few studies (Gruen and Wilkinson 1994, and Amano and Van Norden 1995).

<sup>6</sup> Faruquee (1995) and Kawai and Ohara (1997) both include the terms of trade in the real exchange rate equation along with a broad set of other variables (productivity, relative price of traded to nontraded goods, and the stock of net foreign assets as a share of GNP). They find that the real exchange rate may or may not cointegrate with the terms of trade; the results are idiosyncratic.

<sup>7</sup> MacDonald and Nagayasu (1999) discover a long-run relationship between the real exchange rate and the real interest rate differential for a panel of 14 OECD countries (including the countries in our sample, except for Finland,

Several studies consider these issues in samples of developing countries. Odedokun (1997) uses the terms of trade in a real exchange rate equation for developing (African) countries. Krumm (1993) considers the terms of trade as one factor among a set of structural determinants of the (equilibrium) real exchange rate in the medium-term for Tanzania and the Philippines. Edwards (1989) examines pooled data for 12 developing countries concluding that the external terms of trade, a real fundamental, affects the equilibrium real exchange rate.<sup>8</sup>

Trade theorists see the effects of changes in the terms of trade mainly as an exogenous shock to the real exchange rate in the context of small open economies.<sup>9</sup> The existing theoretical research, however, provides conflicting views on whether a terms-of-trade deterioration appreciates or depreciates the real exchange rate (see Neary, 1988).

Thus, the empirical literature on real-exchange-rate determinants considers primarily large developed economies (e.g., Meese and Rogoff 1988, Edison and Pauls 1993, and so on) and/or developing and middle-income countries (e.g., Edwards 1989). Less attention focuses, however, on small open developed economies. We examine whether expected real interest rate differentials and the terms of trade explain the real exchange rate in such countries. The small country assumption implies that the country takes prices in goods markets and the terms of trade as exogenous. Real interest differentials capture financial market developments, while the terms of trade affects the relative demand for and supply of domestic and foreign goods. All countries in our sample possess a relatively high degree of openness in both goods and assets markets.

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Portugal, and Spain). Interestingly, for our purposes, they find little evidence of a long-run relationship on a country-by-country basis.

<sup>8</sup> Specifically, this empirical work supports the view that an improvement in the terms of trade results in an appreciation of the real exchange rate. Edwards (1989) also includes other variables -- government consumption of nontraded goods, a proxy for exchange and trade controls, a measure of technological progress, the ratio of investment to GDP, and the lagged ratio of net capital flows to GDP. He does not consider, however, real interest rate differentials.

<sup>9</sup> The real exchange rate equals the price of traded to nontraded goods.

### III. The Model

Two basic methods exist for specifying real exchange rate determination – structural and time-series models.<sup>10</sup> The structural approach uses time-tested theoretical concepts such as purchasing power and uncovered interest rate parity to specify the relationship between the real exchange rate and its determinants. Meese and Rogoff (1988) provide a typical derivation of the short-run relationship, for example, between the real exchange rate, on the one hand, and the expected real interest rate differential and the long-run real exchange rate, on the other.<sup>11</sup>

The typical result of such an exercise<sup>12</sup> gives the following expression:

$$q_t = \gamma ({}_k r_t - {}_k r_t^*) + \bar{q}_t, \quad (1)$$

where  $q_t$  and  $\bar{q}_t$  are the natural logarithms of the short- and long-run real exchange rates,  $({}_k r_t - {}_k r_t^*)$  is the real interest rate differential, and  $\gamma = -1$ . Variations in the real exchange rate partly reflect variations in the flexible-price equilibrium real exchange rate and in the expected real interest rate differential. Although this model incorporates sticky-price considerations (as reflected in the expected real interest rate differential), it does not permit real shocks that affect the long-run equilibrium real exchange rate.

Researchers then hypothesize determinants of the long-run real exchange rate and derive the final estimating relationship.<sup>13</sup> A number of real variables possibly determine  $\bar{q}_t$ . In Hooper and Morton (1982), for example,  $\bar{q}_t$  varies over time as a function of home and foreign cumulated trade balances. Edison and Pauls (1993) introduce other variables such as cumulated current accounts, but generally with negative results. As noted above, MacDonald (1998) finds cointegration between the real exchange rate

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<sup>10</sup> Zellner and Palm (1974) show the direct link between dynamic structural and time-series models. That is, dynamic structural models reduce to a series of univariate ARIMA models for the endogenous variables with restrictions on the coefficients of the ARIMA models dictated by the structural specification. Ahking and Miller (1987) provide an example for models of exchange rate determination.

<sup>11</sup> Other researchers (e.g., Baxter 1994, Coughlin and Koedijk 1990, and Gruen and Wilkinson 1994) also adopt this approach.

<sup>12</sup> Appendix A provides details.

<sup>13</sup> Note that the resulting relationships are basically error-correction models with coefficient restrictions that limits the scope of short-run adjustment based on the short-run theoretical structural model initially proposed.



and the real interest rate differential alone, and in combination with the terms of trade, differences in productivity, relative fiscal balances, and net foreign assets.

The non-structural time-series approach begins by defining the determinants of the long-run real exchange rate, which is identified as a cointegrating relationship. Then the dynamic short-run movements in the real exchange rate flow out of an error-correction model that incorporates the real exchange rate and its long-run determinants.

We postulate that the long-run real exchange rate depends on the real interest rate differential and the terms of trade.<sup>14</sup> We also adopt the non-structural, time-series method for modeling the real exchange rate in the long and short run. Thus, we specify the long-run equation as follows:

$$q_t = b_1 + b_2 id_t + b_3 \tau_t \quad (2)$$

where  $id_t$  is the real interest rate differential, and  $\tau_t$  is the natural logarithm of the terms of trade. Then the short-run adjustment follows an error-correction specification as follows:

$$\begin{aligned} \Delta q_t = & \alpha_{10} + \alpha_1 (\beta_1 q_{t-1} - \kappa - \beta_2 id_{t-1} - \beta_3 \tau_{t-1}) \\ & + \sum_{i=1}^s \alpha_{11,i} \Delta q_{t-i} + \sum_{i=1}^s \alpha_{12,i} \Delta id_{t-i} + \sum_{i=1}^s \alpha_{13,i} \Delta \tau_{t-i} + u_{1,t} \end{aligned} \quad (3)$$

$$\begin{aligned} \Delta id_t = & \alpha_{20} + \alpha_2 (\beta_1 q_{t-1} - \kappa - \beta_2 id_{t-1} - \beta_3 \tau_{t-1}) \\ & + \sum_{i=1}^s \alpha_{21,i} \Delta q_{t-i} + \sum_{i=1}^s \alpha_{22,i} \Delta id_{t-i} + \sum_{i=1}^s \alpha_{23,i} \Delta \tau_{t-i} + u_{2,t} \end{aligned}, \text{ and} \quad (4)$$

$$\begin{aligned} \Delta \tau_t = & \alpha_{30} + \alpha_3 (\beta_1 q_{t-1} - \kappa - \beta_2 id_{t-1} - \beta_3 \tau_{t-1}) \\ & + \sum_{i=1}^s \alpha_{31,i} \Delta q_{t-i} + \sum_{i=1}^s \alpha_{32,i} \Delta id_{t-i} + \sum_{i=1}^s \alpha_{33,i} \Delta \tau_{t-i} + u_{3,t} \end{aligned} \quad (5)$$

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<sup>14</sup> The natural logarithm of the long-run real exchange rate in equation (1) equals the natural logarithm of the long-run nominal exchange rate and the difference in the natural logarithms of the long-run price levels in the two countries. Engel (1993, 1999) and MacDonald (1998) define the domestic and foreign price levels as geometric weighted averages of traded and nontraded goods price indexes. Substitution into equation (1) then implies that the real exchange rate depends on the terms of trade and the traded to nontraded goods price indexes in each country. Appendix A provides details.

where  $(\beta_1 q_{t-1} - \kappa - \beta_2 id_{t-1} - \beta_3 \tau \tau_{t-1})$  represents the lagged residuals of the cointegrating relationships and  $\alpha_j$  ( $j=1,2,3$ ) equal the speed of adjustment parameters.<sup>15</sup> Note that  $E(u_{1,t}) = E(u_{2,t}) = E(u_{3,t}) = 0$ ,  $Var(u_{1,t}) = \sigma_{u_{1,t}}^2$ ,  $Var(u_{2,t}) = \sigma_{u_{2,t}}^2$ , and  $Var(u_{3,t}) = \sigma_{u_{3,t}}^2$ .

#### IV. Data and Method

##### *Data*

The sample periods for six out of nine countries run from 1973 to 1994 or 1995. The sample periods for Finland, Portugal, and Spain start in 1978, 1983, and 1978, respectively. The sample period selection for each country was constrained by data availability. We employ quarterly time-series data from the International Financial Statistics on CD-ROM.

The series for the *bilateral* real exchange rate ( $q$ ) is constructed as  $q = eCPI^{US}/CPI^H$ , where  $CPI^{US}$  and  $CPI^H$  are the consumer price indices in the U.S. (foreign country) and the country under consideration (home country), respectively, and  $e$  is the average quarterly nominal exchange rate.<sup>16</sup> For the terms of trade, we use the ratio of export unit value to import unit value when available and the ratio of export prices to import prices otherwise. The expected real interest rate equals the difference between the interest rate on long-term government bonds and the expected inflation rate, where the expected inflation rate equals a two-year centered moving average, incorporating both backward and forward-looking elements. All variables are in logs, with the exception of the real interest rates.

##### *Unit-Root Tests*

We perform the augmented Dickey-Fuller (ADF) test to determine whether the real exchange rates, the real interest differentials, and the terms of trade are stationary. The Akaike information criterion (AIC) and Schwartz Bayesian criterion (SBC) select the optimal lag length ( $k$ ). Usually these two criteria identify the same lag length. When the lag lengths selected differ but the results of the two models agree regarding the existence of the unit root, we report the statistics with the lag length suggested by SBC.

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<sup>15</sup> Equation (2) normalizes the coefficient on the real exchange rate to equal one. Thus,  $b_1$ ,  $b_2$  and  $b_3$  correspond to  $(\kappa/\beta_1)$ ,  $(\beta_2/\beta_1)$ , and  $(\beta_3/\beta_1)_t$  respectively.

Some series, however, seem to have a unit root when we use the lag length suggested by one criterion and seem not to have a unit root when we use the lag length suggested by the other criterion. For these series, we report both statistics (i.e., the real exchange rate for Spain and the expected real interest rate differential for Portugal).

The ADF tests for the real exchange rate ( $q_t$ ), the terms of trade ( $\tau\tau_t$ ), and the real interest differential ( $id_t$ ) rely on the following specification:

$$\Delta x_t = \delta_0 + \delta_1 tr + \delta_2 x_{t-1} + \sum_{j=1}^k \psi_j \Delta x_{t-j} + \varepsilon_{it}, \quad (6)$$

where  $x_t$  equals  $q_t$ ,  $\tau\tau_t$ , and  $id_t$  in turn,  $tr$  is a time trend, and  $\varepsilon$  is a well-behaved random error.

#### *Cointegration and Error-Correction Analyses*

We adopt Johansen's (1988) maximum likelihood technique to conduct our cointegration analysis. If at least one cointegrating vector exists, then according to Granger (1983) and Engle and Granger (1987) a valid error-correction representation of the data must exist, as seen in equations (3), (4), and (5).

In the error-correction model, the disturbance term related to the real interest rate could potentially exhibit autocorrelation as a result of the overleaping in data from using quarterly data when expectations are calculated on a 2-year horizon. Such concerns, however, are alleviated by the lag selection procedure of the lag length of the VAR. This specification assumes one cointegrating vector, which, as we show below, is consistent with our findings. If we had two cointegrating vectors, then we would need two error-correction terms in the error-correction model.

#### *Lag-Order Determination*

Before proceeding further, the optimal lag length ( $s$ ) for the vector error-correction model ensures Gaussian residuals. We first consider the AIC and SBC criteria, applied to the unrestricted VAR model in levels given by equations (3) to (5). In some cases, however, these two criteria suggest a relatively large number of lags, consuming degrees of freedom. We use a likelihood ratio test statistic to exclude

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<sup>16</sup> The nominal exchange rate equals the price of foreign currency (US dollars) in terms of the home currency.

unnecessary lags (Sims 1980, Holden and Perman 1994, and Enders 1995). This method follows the general-to-specific model selection procedure.

While some researchers (e.g., Hatanaka 1996, and Holden and Perman 1994) test sequentially for the exclusion of individual lags, others test sequentially for the exclusion of groups of lags (e.g., Enders 1995). We test the sequential exclusion of lags in both forms.

#### *Deterministic Components in the Data and the Cointegration Space*

Inappropriate assumptions about the existence of deterministic components in the time series can produce misleading inferences. For example, the exclusion of a relevant trend causes bias, making the rejection of the null hypothesis of non-stationarity unlikely.

Different assumptions about the presence of deterministic components in the time series and about the presence of intercepts and/or trends in the cointegrating relationships relate to different asymptotic distributions (Johansen 1992). No generally accepted criterion exists that allows a researcher to specify the best model a priori. Deciding, for example, whether a trend exists (and if it is linear or not) by visual inspection of time-series plots of the data reflects standard practice (e.g., Strauss 1996). Such an ad hoc approach, however, can easily prove misleading.<sup>17</sup> Although Johansen (1992) discusses the issue of joint determination of the cointegration rank and of the existence of a linear trend, he does not provide a model selection method. Instead, he suggests that researchers consider all possible sub-hypotheses about the form of the model and construct a test statistic for each one of the limiting distributions. Then one can reject the null hypothesis only after rejecting all the sub-hypotheses. This suggestion generalizes Pantula's (1989) work on unit-root testing in univariate time series. Harris (1995) interprets this method (i.e., moving from the most restrictive sub-hypothesis to less restrictive ones and stopping the first time that a model fails to reject the null) as more narrow than Johansen's suggestion. For some data sets, the more restrictive sub-hypotheses make the rejection of the null of no-cointegration easier.

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<sup>17</sup> Another method tests more formally for the presence of a trend in a series  $y_t$  by running a regression,  $y_t = \gamma_0 + \gamma_1 tr + \gamma_2 (tr)^2 + e_t$ . Since we know, however, that the series are non-stationary, the  $R^2$  and the  $t$ - and  $F$ -statistics are not reliable.

We follow Harris's (1995) approach and consider different models that provide different combinations of assumptions about the deterministic components in the series and the presence of trend or intercept in the cointegrating equations following the work of Osterwald-Lenum (1992). In particular, we consider a model with a constant in the cointegrating vectors, a model with trends in the time series and a constant in the cointegrating vectors, and a model with both a constant and a linear trend in the cointegrating vectors. All three models allow for a constant in the cointegrating space because we use index numbers in our equations.

## **V. Results**

Table 1 reports the results of the unit-root tests. Statistics on constants and trends appear only when they are significant and are included in the model. Of the nine countries examined, five clearly have unit-roots in all three series (Australia, Canada, Finland, Italy, and New Zealand). The remaining countries also have unit-roots in most series, but with weaker evidence of non-stationarity. The real exchange rate and the terms of trade for Austria and Norway are non-stationary, but we cannot reject the null of non-stationarity for the expected real interest rate differential at the 5-percent level. Some evidence of stationarity of the expected real interest rate differential also exists for Portugal and Spain. The terms of trade test as non-stationary for all countries except Portugal and Spain. The only country with some evidence of stationarity in the real exchange rate is Portugal. If the 1-percent significance level identifies stationary series, then only the terms of trade in Portugal and Spain exhibit stationary behavior.

Table 2 provides the results of the cointegration analysis. The model with a constant in the cointegration space emerges as the selection for all nine countries. Whenever cointegration occurs only one cointegrating vector exists. Both the  $\lambda$ -trace and  $\lambda$ -max statistics indicate a unique cointegrating vector for Australia, Austria, Canada, Norway, New Zealand, and Portugal at the 90 percent level (Table 2). The  $\lambda$ -max statistic also indicates a unique cointegrating vector for Finland and Italy at the 90 percent level; the  $\lambda$ -trace statistic indicates no cointegration. Finally, the  $\lambda$ -max statistic indicates two

cointegrating vectors at the 90-percent level for Spain; the  $\lambda$ -trace statistic indicates only one cointegrating vector. In sum, we adopt the specification of one cointegrating vector for further analysis.

Table 3 reports the estimated cointegrating relationships normalized by the real exchange rate. Our results confirm the theoretical ambiguity regarding the effects of terms of trade on the real exchange rate. Typically, a terms-of-trade improvement leads to a real exchange rate appreciation (negative coefficient). That prediction occurs in four of the nine countries Australia, Canada, New Zealand, and Spain considered. Chen and Rogoff (2002) argue that Australia, Canada, and New Zealand “are near perfect examples of....well-developed, small open economies ... where internal and external markets operate with little intervention, and where floating exchange rate regimes have been implemented for a sufficiently long period of time” (p.7). In the other countries, a terms of trade improvement leads to a real depreciation.

The positive sign on the terms of trade in some countries does not necessarily contradict theory, however, because the predictions of the theoretical literature for the effects of terms-of-trade changes on the real exchange rate are more complicated. First, the real appreciation of the domestic currency when the terms of trade improve may reflect a relatively high weight of exportables in the home price level so that the rise in the price of the exportables raises the home price level. Similarly, the depreciation of home currency rate when the terms of trade improve may result from a relatively high weight of importables in the home price level. Second, a terms-of-trade improvement should unambiguously lead to a real exchange rate appreciation only when the traded and nontraded goods are substitutes. In that case, both the direct effect from a terms-of-trade improvement on income and the indirect effect through the relative price of traded goods result in an increase in the demand for nontraded goods and therefore to a real exchange rate appreciation. If, however, the traded and nontraded goods are complements, then the indirect effect of the terms-of-trade improvement through the relative price of traded goods leads to a decline in the demand for nontraded goods and to a real exchange rate depreciation. Therefore, the overall effect of the a terms-of-trade improvement on the real exchange rate cannot be determined without

knowing the relative strength of the various effects.<sup>18</sup> Third, the countries with a positive sign between the real exchange rate and the terms of trade are open and trade mainly in manufacturing products. But at the same time they have a large component of services (nontraded goods) sector, which may display a low degree of substitutability. Fourth, rigidities like “pricing-to-market” and “local-currency pricing” may distort the workings of the indirect effect identified above and thus impede the substitutability of traded and nontraded goods. The expectation that terms of trade improvements associate with real exchange rate appreciation emerge because of the unitary pass-through assumption, which underlies many traditional models. The new open economy macroeconomics literature, however, emphasizes “pricing-to-market” and “local-currency pricing” behavior. Obstfeld and Rogoff (2000), for example, develop such a model where imports are invoiced in the importing country’s currency. That model shows that unexpected currency depreciations associate with improvements in the terms of trade.

The expected real interest rate differential possesses the predicted negative relationship with the real exchange rate in five countries (Australia, Canada, Finland, Italy, and Portugal). Failure to verify empirically that high real domestic interest rates lead to a real currency appreciation is not uncommon, usually attributed to the use of monetary policy to defend pegged exchange rates (e.g., International Monetary Fund 1996). Since our sample considers exchange rates that were more or less flexible, we believe that this result may exist because of the presence of different restrictions on capital flows.<sup>19</sup>

We test whether we can exclude the real interest rate differential from the estimated cointegrating relationship, using the likelihood ratio test suggested by Johansen and Juselius (1990). Testing for exclusion restrictions, we cannot exclude the real interest rate differential in the five countries where it displays the predicted negative sign (Table 4). We can exclude it, however, in two countries where it displays the non-predicted, positive sign (Austria and Norway). That leaves the task of resolving the puzzle (of the positive sign) in two countries (New Zealand and Spain). New Zealand preserved capital

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<sup>18</sup> For a discussion of the terms of trade effects on the real exchange rate in explicit terms of income and substitution effects, see Neary (1988), who provides a microeconomic model based on utility theory.

controls until 1983, and also experienced both fixed and flexible exchange rate regimes during that period. Although we break the sample to include only fixed or flexible exchange rates below, we still cannot obtain definitive results (possibly because the sub-samples are too short). Spain, on the other hand, kept capital controls throughout the period considered, including both a flexible exchange rate period and their ERM participation with wide margins.

We also test whether we can exclude the terms of trade from the estimated cointegrating relationships (Table 4). We can strongly reject the null hypothesis of excluding the terms of trade from the cointegrating vectors for all countries. In general, the terms of trade prove more important at the margin than the expected real interest rate differentials. Those results prove consistent with the earlier findings of Gruen and Wilkinson (1994) and Amano and van Norden (1995) for Australia and Canada, respectively, showing that in addition to monetary factors, the terms of trade provide an important explanatory variable of the real exchange rate in the long run. In contrast to Gruen and Wilkinson (1994), however, our results suggest that the terms of trade prove quantitatively more important in explaining the long-run real exchange rate than the real interest rate differential.<sup>20</sup> Our results for Canada support Amano and van Norden's (1995) findings.

Table 5 provides the estimates of the speed of adjustment coefficients ( $\alpha_i$ ) in the error-correction model, where the error-correction term is normalized on the real exchange rate (i.e., the error correction term equals  $[q_{t-1} - (\kappa / \beta_1) - (\beta_2 / \beta_1)id_{t-1} - (\beta_3 / \beta_1)\tau\tau_{t-1}]$ ). Suppose that the error-correction term exceeds zero. Then stability of the adjustment process requires that as the real exchange rate ( $q$ ) falls, the expected real interest rate differential ( $id$ ) rises (falls) if its coefficient ( $\beta_2/\beta_1$ ) is positive (negative), and the terms of trade ( $\tau\tau$ ) rises (falls) if its coefficient ( $\beta_3/\beta_1$ ) is positive (negative), or some combination thereof. Now, compare Tables 3 and 5. If ( $\beta_2/\beta_1$ ) or ( $\beta_3/\beta_1$ ) are positive (negative), then stabilizing

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<sup>19</sup> Although the economies in our sample are relatively open, the condition of perfect capital mobility is not met.

<sup>20</sup> This result holds for all countries in our sample, including Australia. While Gruen and Wilkinson's (1994) sample period extends to 1990, our sample extends to the third quarter of 1995.



adjustment requires that the corresponding  $\alpha_2$  or  $\alpha_3$  is positive (negative). That outcome occurs in every case except for Australia's expected real interest rate differential, and Spain's terms of trade. For those two anomalies, neither speed of adjustment coefficient is significant. Stabilizing adjustment also requires that the speed of adjustment coefficient in the real exchange rate equation ( $\alpha_l$ ) is negative, which occurs in every case.

Limiting our discussion to significant and stabilizing speed of adjustment coefficients, several observations emerge. First, movements in the expected interest rate differential provide stabilizing adjustment toward long-run equilibrium in six countries (Austria, Canada, Finland, Italy, Norway, and Spain). Second, the terms of trade produce stabilizing adjustment toward long-run equilibrium in five countries (Australia, Austria, Italy, New Zealand, and Portugal). Finally, the real exchange rate generates stabilizing adjustment toward long-run equilibrium in four countries (Canada, Norway, Portugal, and Spain). Thus, stabilizing adjustment comes from two sources in six countries (Austria, Canada, Italy, Norway, Portugal, and Spain) and from only one source in three countries (Australia, Finland, and New Zealand).

The countries in our sample have not operated under a pure float throughout the sample periods. The relationship between the real exchange rate and its determinants may display different properties under different exchange rate regimes. All countries in our sample, however, display some degree of exchange rate flexibility. The degree of flexibility varies across countries, but each exchange rate provides sufficient flexibility. For example, we only consider those European countries within the Exchange Rate Mechanism that possessed wide bands and frequently realigned. More importantly, all exchange rates are bilateral rates with the US dollar. Thus, since the US did not peg its exchange rate during that period, all countries that pegged the currency to another country's currency or to a basket of currencies experienced some degree of exchange rate flexibility.<sup>21</sup>

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<sup>21</sup> Appendix B provides results for splitting samples into fixed and flexible exchange rates, where such splitting is feasible.

In sum, the expected real interest rate differential plays an important role in the adjustment toward long-run equilibrium. It represents an equilibrating source in six countries and its speed of adjustment coefficient exceeds that of the terms of trade and the real exchange rate in every instance. That finding accords with intuition; capital flows adjust quicker than trade flows.

## **VI. Conclusion**

The recent literature on real exchange rate determination considers the role of financial markets and capital flows, using the expected real interest rate differential as an explanatory variable. Most of this research examines large developed economies. On the other hand, empirical research on real exchange rate determinants in developing countries pays more attention to the role of exogenous real shocks.

We consider the terms of trade and the long-term expected real interest rate differential as exogenous fundamental determinants of the long-run real exchange rate in small open developed economies. The terms of trade incorporate goods market developments, through the relative prices of internationally traded goods. The long-run real interest rate differential incorporates financial market developments, through the relative interest rates of internationally traded assets. In this way, we attempt to capture long-run movements in the real exchange rate that reflect developments in both financial and goods markets.

Our results suggest that the terms of trade and the expected real interest rate differential prove important in explaining the long-run real exchange rate of small open developed economies in the post Bretton-Woods era. For the small open developed economies in our sample, a long-run equilibrium relationship exists among the real exchange rate, the terms of trade, and the long-term expected real interest rate differential. The terms of trade prove an important and non-excludable element in each cointegrating relationship, while the expected real interest rate differentials prove excludable in two countries. The expected real interest rate differential, however, always possesses a quantitatively larger speed of adjustment in moving the real exchange rate toward its long-run equilibrium, than does the terms of trade or the real exchange rate. In some countries, a terms of trade improvement causes a real appreciation while in other countries, a real depreciation. If traded and nontraded goods are complements

and the substitution effect dominates the income effect and/or if importables are a large component of the domestic price level, then a terms-of-trade improvement leads to a real exchange rate depreciation, and vice versa. Alternatively, a negative association between the terms of trade and the real exchange rate may also reflect incomplete pass-through and/or pricing-to-market behavior or local-currency pricing behavior. On the other hand, if traded and nontraded goods are substitutes and/or exportables are a large component of the domestic price level, then a terms-of-trade improvement leads to a real exchange rate appreciation. Our analysis suggests that the evidence provided by single-country studies generalizes to a larger group of small and open developed economies. Whether the real interest rate differential and the terms of trade are the primary factors that affect the real exchange rate in the long-run, rendering other variables redundant, constitutes the subject of further research.

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**Table 1: ADF Tests for Unit Roots in the Real Exchange Rate, Terms of Trade, and Real Interest Rate Differential**

$$\Delta x_t = \delta_0 + \delta_1 tr + \delta_2 x_{t-1} + \sum_{j=1}^k \psi_j \Delta x_{t-j} + \varepsilon_t$$

Country	Sample	Real Exchange Rate ( $q$ )		k ( $q$ )	Terms of Trade ( $\tau\tau$ )		k ( $\tau\tau$ )	Real Interest Differential ( $id$ )		k ( $id$ )
		$\delta_2$	$t(\delta_2)$		$\delta_2$	$t(\delta_2)$		$\delta_2$	$t(\delta_2)$	
				<b>AIC / SBC</b>			<b>AIC / SBC</b>			<b>AIC / SBC</b>
Australia	73:1-95:3	-0.0696	-2.11	3 / 3	-0.0881	-2.46	5 / 5	-0.0523	-1.74	NA / 4
Austria	73:1-94:1	-0.0718	-2.10	3 / 3	-0.1311	-1.72	2 / 2	-0.0705	-2.49**	3 / 3
Canada	73:1-95:2	-0.0451	-2.00	NA / 4	-0.0952	-2.14	2 / 2	-0.0342	-1.31	1 / 1
Finland	73:1-94:1	0.0003	0.08	NA / 1	-0.0201	-0.98	NA / 1	-0.0338	-1.38	1 / 1
Italy	73:1-95:2	-0.0694	-2.09	1 / 1	-0.0245	-1.22	5 / 5	-0.0429	-1.49	1 / 1
New Zealand	73:1-95:2	-0.0720	-1.95	1 / 1	-0.1419	-3.27	5 / 5	-0.0453	-1.59	1 / 1
Norway	73:1-95:2	-0.0793	-2.08	1 / 1	-0.0599	-1.62	1 / 1	-0.0685	-2.22**	1 / 1
Portugal	83:1-92:4	-0.3975	-3.76**	NA / 1	-0.6048	-4.50*	11 / 11	-0.0663	-2.18**	NA / 3
Spain	78:2-95:2	-0.3014	-2.40	2 / NA	-0.1397	-4.36*	4 / 4	-0.0545	-1.29	5 / NA
		-0.0002	-0.16	1 / 1				-0.0883	-2.28**	NA / 2

NOTE: The variable  $x$  in the equation equals the real exchange rate ( $q$ ), the terms of trade ( $\tau\tau$ ), and the real interest rate differential ( $id$ ) for the alternative tests. The test statistics [ $t(\delta_2)$  for the coefficient  $\delta_2$ ] are as follows:  $\tau_\tau$  with a constant and a time trend for the terms of trade in Australia, Canada, New Zealand, and Spain and for the real exchange rate in Portugal;  $\tau_\mu$  with a constant but without a time trend for the real exchange rate in Australia, Austria, Canada, Italy, Norway, and New Zealand;  $\tau$  without a constant and a time trend for all other tests. The critical values for the various  $\tau$ -statistics are reported in MacKinnon (1991). The Akaike information and Schwartz Bayesian criteria are reported as *AIC* and *SBC*, respectively. Usually, the AIC and SBC suggest the same lag length. If there is disagreement about the number of lags and the stationarity tests yields the same qualitative outcome, we report the results for the SBC. In two cases (Portugal and Spain) the AIC and SBC lead to different lag lengths and the test statistics differ as to whether the series in question are stationary. In those cases, we report both results. Finally, NA means not applicable.

\* means significantly different from zero at the 1-percent level

\*\* means significantly different from zero at the 5-percent level

**Table 2: Model with Intercept in the Cointegrating Vector and No Trend in the Data**

				<b><math>\lambda</math>–max Tests</b>								
$H_0$	$H_a$	n-r	90%	Australia	Austria	Canada	Finland	Italy	New Zealand	Norway	Portugal	Spain
$r=0$	$r=1$	3	14.09	26.65†	20.51†	26.43†	18.95†	18.21†	61.51†	22.29†	37.94†	31.94†
$r=1$	$r=2$	2	10.29	7.72	9.56	5.06	5.10	8.49	7.45	7.61	5.92	12.46†
$r=2$	$r=3$	1	7.50	3.82	2.68	3.62	2.31	1.48	3.92	2.89	3.68	1.93
				<b><math>\lambda</math>–trace Tests</b>								
$H_0$	$H_a$	n-r	90%	Australia	Austria	Canada	Finland	Italy	New Zealand	New Zealand	Portugal	Spain
$r=0$	$r>0$	3	31.88	35.18†	32.75†	35.11†	26.36	28.18	72.88†	32.79†	47.54†	46.32†
$r\leq 1$	$r>1$	2	17.79	9.53	12.24	8.68	7.41	9.97	11.37	10.50	9.60	14.39
$r\leq 2$	$r>2$	1	7.50	3.82	2.68	3.62	2.31	1.48	3.92	2.89	3.68	1.93
Number of Lags				6	4	4	2	2	4	2	3	2

Note:  $H_0$  is the null hypothesis and  $H_a$  is the alternative hypothesis. The number of variables and cointegrating vectors are n and r, respectively. The critical values for the  $\lambda$ –max and  $\lambda$ –trace tests are from Johansen and Nielson (1993), as displayed by *CATS*.

† Means significant at the 10-percent level.



**Table 3: Estimated Cointegrating Vectors Normalized by Coefficient of the Real Exchange Rate**

$$q = (\kappa/\beta_1) + (\beta_2/\beta_1)id + (\beta_3/\beta_1)\tau\tau$$

Country	Coefficient of Variable			
	$q$	$id$	$\tau\tau$	$constant$
Australia	1	-0.041	-1.385	0.235
Austria	1	0.031	4.967	2.764
Canada	1	-0.039	-0.696	0.305
Finland	1	-0.025	3.287	1.828
Italy	1	-0.356	10.065	7.891
New Zealand	1	0.028	-3.184	0.411
Norway	1	0.003	0.847	1.935
Portugal	1	-0.077	2.704	5.053
Spain	1	0.158	-4.033	4.290

Note: See Table 1. The normalization also affects the size, but not the sign, of the speed of adjustment parameters reported in Table 5.

**TABLE 4: Tests for Excluding the Expected Real Interest Rate Differential and Terms of Trade from the Cointegrating Vector**

$$\beta_1 q - \beta_2 id - \beta_3 \tau\tau - \kappa \sim I(0)$$

Country	LR( $\eta$ ) statistic		$\eta$
	$H_0: \beta_2 = 0$	$H_0: \beta_3 = 0$	
Australia	6.63**	11.44**	1
Austria	0.20	2.77**	1
Canada	11.20**	13.15**	1
Finland	10.44**	12.33**	1
Italy	9.20**	6.21**	1
New Zealand	13.57**	53.79**	1
Norway	0.09	12.95**	1
Portugal	13.39**	25.33**	1
Spain	17.41**	1.55**	1

Note: The LR( $\eta$ ) is the likelihood ratio statistic with  $\eta$  degrees of freedom.  $H_0$  is the null hypothesis.

\* means significant at the 1-percent level.

\*\* means significant at the 5-percent level

**Table 5: Estimated Speed of Adjustment Coefficients ( $\alpha$ )**

$$\alpha_i [q - (\kappa/\beta_1) - (\beta_2/\beta_1)id - (\beta_3/\beta_1)\tau\tau]$$

<b>Speed of Adjustment Coefficient (<math>\alpha_i</math>) for Variable's Equation</b>						
Country	<i>q</i>	<i>t-value</i>	<i>Id</i>	<i>t-value</i>	<i><math>\tau\tau</math></i>	<i>t-value</i>
Australia	-0.023	-0.75	0.296	0.58	-0.094	-4.96**
Austria	-0.017	-1.26	0.472	2.41**	0.032	4.09**
Canada	-0.042	-2.18**	-2.142	-4.59**	-0.013	-0.52
Finland	-0.066	-1.03	-4.061	-3.70**	0.019	0.58
Italy	-0.005	-0.74	-0.193	-1.53†	0.015	4.13**
New Zealand	-0.001	-0.05	0.188	0.31	-0.111	-9.30**
Norway	-0.204	-4.45**	2.130	2.57**	0.031	0.55
Portugal	-0.235	-3.20**	-1.171	-0.62	0.179	5.00**
Spain	-0.037	-2.41**	1.673	6.36**	0.010	1.11

Note: Results based on the normalization of the cointegrating vector by the coefficient of the real exchange rate.

\*\* means significant at the 5-percent level.

† means significant at the 10-percent level

## Appendix A:

This appendix provides a derivation of the relationship between the real exchange rate, and the expected real interest rate differential and the terms of trade. While the part that links the real exchange rate and the real interest rate differential corresponds to typical derivations (e.g., Meese and Rogoff, 1988), most discussions usually fail to provide a derivation that incorporates the terms of trade as well.

Define  $q_t$  as the log of the real exchange rate,  $e_t$  the log of the nominal exchange rate,  $p_t$  the log of the home price level, and  $p_t^*$  the log of the foreign price level. Thus, the real exchange rate is defined as follows:

$$q_t \equiv e_t + p_t^* - p_t. \quad (A1)$$

Consider uncovered interest rate parity, where  ${}_k i_t$  and  ${}_k i_t^*$  denote the home and foreign nominal interest rates at period  $t$  for  $k$  periods ahead. That is,

$$E_t e_{t+k} - e_t = {}_k i_t - {}_k i_t^*. \quad (A2)$$

This leads to the real interest rate parity condition<sup>22</sup>

$$E_t (q_{t+k} - q_t) = {}_k r_t - {}_k r_t^*, \quad (A3)$$

where  ${}_k r_t$  and  ${}_k r_t^*$  denote expected home and foreign real interest rates.

Under fully flexible prices, the long-run equilibrium real exchange rate for period  $t$  is  $\bar{q}_t$ , and for period  $t+k$  is  $\bar{q}_{t+k}$ . If no real shocks exist (or if all real shocks do not affect the expected long-run equilibrium value of the real exchange rate), then

$$E_t \bar{q}_{t+k} = \bar{q}_t. \quad (A4)$$

Now consider a real exchange rate adjustment mechanism that restores the real exchange rate to its long run equilibrium value  $\bar{q}_{t+k}$  (for  $k=0,1,2,\dots,n$ ) as follows:

$$E_t (q_{t+k} - \bar{q}_{t+k}) = \phi^k (q_t - \bar{q}_t), \quad 0 < \phi < 1, \quad (A5)$$

where  $\phi$  is a speed of adjustment, showing how fast the real exchange rate returns to its long run equilibrium.<sup>23</sup> This stochastic process allows for price-stickiness as in the Dornbush (1976) and Frankel (1979) models.

Combining equations (A4) and (A5) produces the following expression for the real exchange rate

$$E_t (q_{t+k}) - \bar{q}_t = \phi^k (q_t - \bar{q}_t). \quad (A6)$$

Solving equation (A3) for  $E_t (q_{t+k})$  and substituting into equation (A6) gives

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<sup>22</sup> This result emerges when one subtracts the expected inflation differential from both sides of equation (A2).

<sup>23</sup> The higher  $\phi$  is, the slower the adjustment process.

$$q_t = \gamma(r_t - r_t^*) + \bar{q}_t, \quad (\text{A7})$$

where  $\gamma = 1 / (\phi^k - 1) < -1$ .

Variations in the real exchange rate partly reflect variations in the flexible-price equilibrium real exchange rate and in the expected real interest rate differential. Although this model incorporates sticky-price considerations (as reflected in the expected real interest rate differential), it does not permit real shocks that affect the long-run equilibrium real exchange rate. A number of real variables possibly determine  $\bar{q}_t$ .

Assuming that the nominal exchange rate ( $e_t$ ) follows a random walk in the face of real shocks (or assuming that the foreign exchange market is in equilibrium), we write  $\bar{q}_t$  as follows (all variables are in logs):

$$\bar{q}_t = \bar{e}_t + p_t^* - p_t. \quad (\text{A8})$$

where  $\bar{e}_t$  is the long-run equilibrium nominal exchange rate.

Now, following Engel (1993, 1999) and MacDonald (1998), we write the domestic and foreign price levels as geometric weighted averages of traded and nontraded goods prices ( $p_t^T, p_t^{T*}$  and  $p_t^N, p_t^{N*}$ , respectively). Thus, we have

$$p_t = ap_t^N + (1-a)p_t^T \quad \text{and} \quad p_t^* = a^* p_t^{N*} + (1-a^*)p_t^{T*}. \quad (\text{A9})$$

Equations (A8) and (A9) lead to

$$\bar{q}_t = \bar{e}_t + (p_t^{T*} - p_t^T) - [a(p_t^N - p_t^T) - a^*(p_t^{N*} - p_t^{T*})]. \quad (\text{A10})$$

The term in brackets represents the weighted differences between the log of the relative prices of traded and nontraded goods in the home and foreign countries. Finally, substituting equation (A10) into equation (A7) produces

$$q_t = \gamma(r_t - r_t^*) + \bar{e}_t + (p_t^{T*} - p_t^T) - [a(p_t^N - p_t^T) - a^*(p_t^{N*} - p_t^{T*})]. \quad (\text{A11})$$

We abstract from the traded-nontraded-goods relation, and focus on the difference of the (logarithms of the) price indices of traded goods between home and foreign countries by assuming that both home and foreign countries specialize in the production of their exportables.<sup>24</sup> Thus, the domestic price level for traded goods consists of the price level of only home exportables and the foreign price level for traded

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<sup>24</sup> This is for theoretical reasons, since we focus on the traded goods component of the real exchange rate, and for practical reasons, since the data required for such an exercise are hard to find. As Engel (1999) notes “determining precise price indexes for nontraded goods and traded goods is an impossible task given the quality of data available”

goods consists of the price level of foreign exportables, the price level of home importables. So  $(p^T)$  equals the logarithm of the price of home exportables and  $(p^{T*})$  equals the logarithm of prices of home importables. Since  $(\bar{e}_t)$  equals the nominal exchange rate (and assuming that the foreign exchange market is in equilibrium), then the term  $[\bar{e}_t + (p_t^{T*} - p_t^T)]$  measures the logarithm of the inverse of the terms of trade of the home country. Alternatively, it measures the real exchange rate for traded goods (MacDonald 1998).

## **Appendix B:**

We redo the cointegration analysis by splitting the samples where appropriate. For the nine countries, this was only feasible and meaningful for Australia and New Zealand.<sup>25</sup> In particular, Australia followed a trade-weighted peg until 1983 and moved to flexible exchange rates afterwards. New Zealand pegged either to the US dollar or to a trade weighted index until 1978, then the exchange rate floated between 1979-1981 before it returned again to a trade-weighted peg between 1982-1984 and finally the exchange rate became flexible.<sup>26</sup> In every instance except for New Zealand under flexible exchange rates, we find one cointegrating vector. For the exception, we find two cointegrating vectors.

We provide the results of the split-sample analysis in Tables B3 and B5. Splitting the sample does not affect the results for the cointegrating vectors in Australia. As in the full sample results, the coefficients on both the real interest rate differential and the terms of trade are negative for both the fixed and flexible exchange rate periods. In New Zealand, the signs on the interest rate differential and the terms of trade during the fixed exchange rate period are the same as when we consider the full sample.

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(p. 508). Moreover, he finds that the relative prices of nontraded to traded goods accounts for little of the movements of the U.S. real exchange rate.

<sup>25</sup> Splitting the sample did not apply to other countries for various reasons. For example, the exchange rate regime in Austria and Canada did not change throughout the sample. Finland and Norway, on the other hand, switched from one exchange regime to another too often to make the splitting feasible (or meaningful). Moreover, such frequent changes of the peg and switches from pegged to flexible exchange rates practically imply a high degree of flexibility. Italy, Spain, and Portugal had flexible exchange rates before joining the ERM and once they joined all three possessed bands of fluctuations ( $\pm 6\%$ ) wider than the standard. In addition, Italy realigned its central parity extremely often. Thus, the exchange rate enjoyed a significant degree of flexibility in practice. Our sample covers only one year of ERM participation for Portugal. Finally, splitting the sample is not feasible for Spain, since that leaves only six years of observations during the ERM period, too little data for cointegration analysis.

<sup>26</sup> For a detailed chronology and typology of exchange rate regimes we follow Cottarelli and Giannini (1997).

When we consider the flexible exchange rate period the coefficients emerging from the first cointegrating vector are consistent with those of the full sample and the fixed exchange rates period while the coefficients from the second cointegrating vector are reversed. We suspect that this result may reflect the short data span in the split-samples. On balance, however, the signs correspond for the fixed and flexible exchange rates periods.

The speed-of-adjustment coefficients in the error-correction models prove qualitatively consistent when considering the full sample and the fixed exchange rate period. With flexible exchange rates, the sign on the real interest rate differential coefficient becomes negative. In New Zealand, the coefficient on the real exchange rate is always negative regardless of the sample. It is statistically significant, however, only for one of the two cointegrating vectors of the flexible exchange rate period. The coefficient on the interest rate differential is positive in all cases except for one of the two cointegrating vectors of the flexible exchange rates period. This is, however, the only case that this coefficient appears statistically significant. The terms-of-trade adjustment coefficient has a negative sign and is statistically significant for the full sample period and for the fixed exchange rates period. When we consider the data from the flexible exchange rates period this coefficient turns positive but its statistical significance is much lower.

The above results, however, should be treated with caution since in most cases we rely on around a decade's quarterly data and this may not be the ideal data span for cointegration analysis.

**Table B3: Estimated Cointegrating Vectors Normalized by Coefficient of the Real Exchange Rate (for Fixed and Flexible Exchange Rates)**

Country	Coefficient of Variable			
	$q$	$Id$	$\tau\tau$	$Constant$
Australia 73:1-83:3 (fixed)	1	-0.089	-3.190	0.460
Australia 83:4-95:3 (flexible)	1	-0.053	-0.831	0.327
New Zealand 73:1-85:2 (fixed)	1	0.027	-6.349	0.059
New Zealand 85:2-95:2 (flexible), 1 <sup>st</sup> coint. Vector	1	0.080	-2.275	0.338
New Zealand 85:2-95:2 (flexible) 2 <sup>nd</sup> coint. Vector	1	-0.060	0.633	0.725

Note: See Table 1. The normalization also affects the size, but not the sign, of the speed of adjustment parameters reported in Table 5.

**Table B5: Estimated Speed of Adjustment Coefficients ( $\alpha$ ) (for Fixed and Flexible Exchange Rates)**

$$\alpha_i [q - (\kappa/\beta_1) - (\beta_2/\beta_1)id - (\beta_3/\beta_1)\tau\tau]$$

Country	Speed of Adjustment Coefficient ( $\alpha_i$ ) for Variable's Equation					
	$q$	$t\text{-value}$	$id$	$t\text{-value}$	$\tau\tau$	$t\text{-value}$
Australia 73:1-83:3 (fixed)	-0.017	-1.496*	0.145	0.518	-0.042	-3.983**
Australia 83:4-95:3 (flexible)	-0.142	-1.401*	-1.533	-1.219	-0.125	-2.592**
New Zealand 73:1-85:2 (fixed)	-0.007	-0.827	0.111	0.432	-0.029	-4.736**
New Zealand 85:2-95:2 (flexible), 1 <sup>st</sup> coint. Vector	-0.189	-5.619**	0.260	0.331	0.040	1.580*
New Zealand 85:2-95:2 (flexible) 2 <sup>nd</sup> coint. Vector	-0.047	-1.277	-3.137	-3.678**	0.007	0.266

Note: Results based on the normalization of the cointegrating vector by the coefficient of the real exchange rate.

\*\* means significant at the 5-percent level.

† means significant at the 10-percent level