AUSTRALIA'S REAL EXCHANGE RATE – IS IT EXPLAINED BY THE TERMS OF TRADE OR BY REAL INTEREST DIFFERENTIALS?

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ABSTRACT

We use time series techniques to examine the behaviour of Australia's real exchange rate from 1969 to 1990. The real exchange rate exhibits non-stationary behaviour over this period, in contrast to simple purchasing power parity theory. We find weak evidence that the real exchange rate exhibits a stable long run relationship with the terms of trade. There is no stable long run relationship between the real exchange rate and either short or long real interest differentials between Australia and its major trading partners.

Since the float of the Australian dollar and the world-wide deregulation of financial markets, we find some evidence that the real exchange rate exhibits a stable relationship with the terms of trade alone, and with **long** real interest differentials alone. The evidence for a stable relationship is clearest with long real interest differentials.

After the float, we also find evidence that the terms of trade and long real interest differentials **together** help to explain the Australian real exchange rate. We estimate the number of independent long run relationships between the real exchange rate, the terms of trade and long real interest differentials and, for some specifications, find evidence of two independent relationships.

Since the float, our best estimates are that a 1 per cent improvement in the terms of trade leads to an appreciation of the Australian real exchange rate of about 0.3 to 0.5 per cent, while an increase of 1 percentage point in the differential between Australian and world long real interest rates is associated with an appreciation of the Australian real exchange rate of about 2 to 3 1/2 per cent.

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1. INTRODUCTION

As part of the deregulation of Australia's financial markets, the Australian dollar was floated in December 1983. The real exchange rate[1], after having been roughly steady since the mid-1970s, depreciated by about 35 per cent between the beginning of 1984 and mid-1986. Over the rest of 1980s, about two-thirds of this very large depreciation was unwound.

Many have argued that movements in the Australian real exchange rate are substantially influenced by shifts in the terms of trade (see, for example, McKenzie (1986), Blundell-Wignall and Thomas (1987), Simes (1988), Blundell-Wignall and Gregory (1990), Freebairn (1989) and Murphy and Smith (1991)). As these studies recognise, in a small open commodity exporting economy such as Australia's, the real exchange rate should shift in response to movements in real fundamentals such as the terms of trade.

Another strand of literature (see, for example, Sachs (1985), Dornbusch and Frankel (1987), Isard (1988), Meese and Rogoff (1988) and Blundell-Wignall and Browne (1991)) has attempted to explain movements in the real exchange rates of large OECD countries by real interest differentials. In a world with deregulated financial flows, it is consistent both with sticky-price monetary models such as the Dornbusch (1976) overshooting model and with portfolio-balance models such as Branson (1979) that real exchange rates should be correlated with real interest differentials.

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[1] We use the term "real exchange rate" to mean the nominal exchange rate adjusted for differences in the price level (which we measure by the Consumer Price Index) between between Australia and its trading partners. It is thus an "external" measure of the real exchange rate rather than the relative price of non-traded to traded goods, as implied by the Swan-Salter definition of the real exchange rate. Dwyer (1987) examines the relationship between these two measures of the real exchange rate.

In this paper we assess empirically whether, over the time periods we consider, Australia's real exchange rate is explained by the terms of trade, by real interest differentials, or by a combination of the two. We use time series techniques to address these issues and focus on whether stable long run (so-called cointegrating) relationships exist between all or a subset of these variables.

2. THE STYLISED FACTS

Graph 1 shows Australia's trade-weighted real exchange rate and the terms of trade for goods and services from the December quarter 1969 to the March quarter 1991.[2] There appears to be a clear correlation between movements in the terms of trade and the real exchange rate. Three periods stand out for special mention. Firstly, between 1978 and 1984, the real exchange rate appreciated and was considerably more volatile than the slowly falling terms of trade. Secondly, although in proportionate terms the real exchange rate fell more than the terms of trade between early 1984 and mid-1986, it appreciated less than proportionately in the subsequent four years. Finally, the terms of trade fell 10% from its average level in 1989 to its level in the March quarter 1991. By contrast, the real exchange rate fell 3.5% from its average level in 1989 to its average level in the March quarter 1991, and, by the end of March 1991, was back to its average 1989 level.

Graphs 2 and 3 show Australia's real exchange rate along with the short and long real interest differentials respectively. The real interest differentials are calculated as the difference between Australian real rates and an arithmetic average of real rates in the US, UK, Japan and Germany, using actual inflation over the past 12 months as a proxy for expected inflation. Over the 1970s, these series do not appear to move with the real

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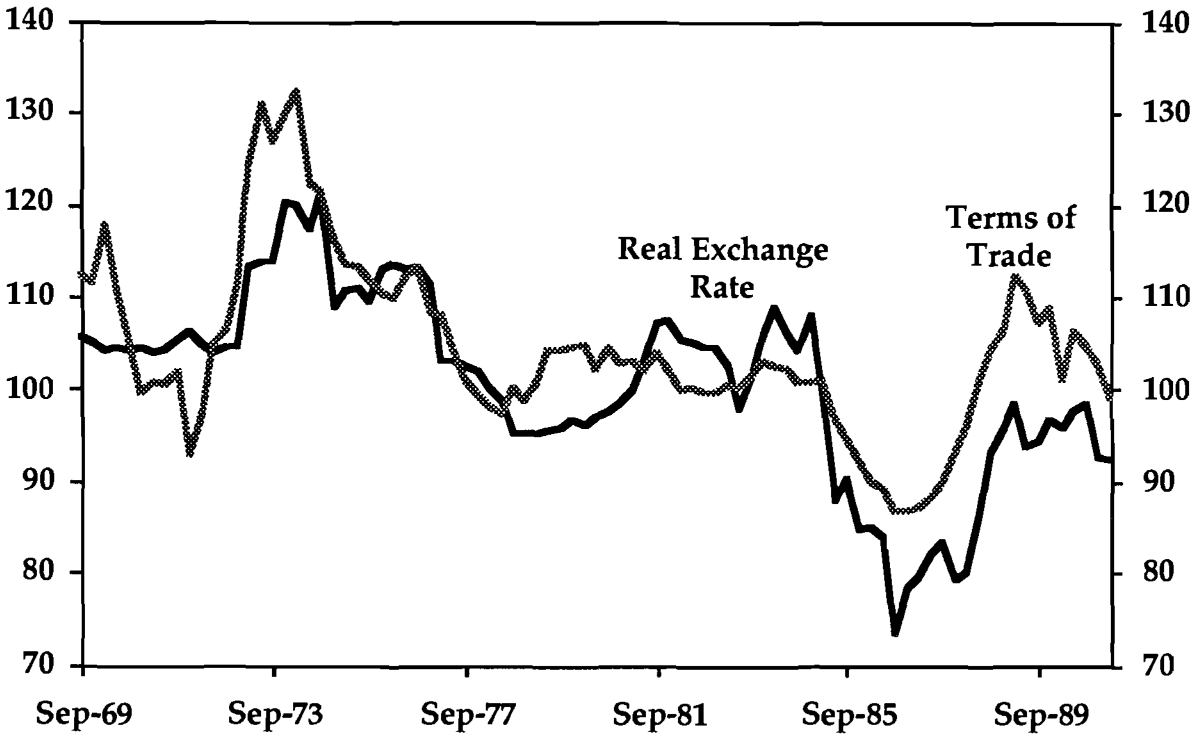
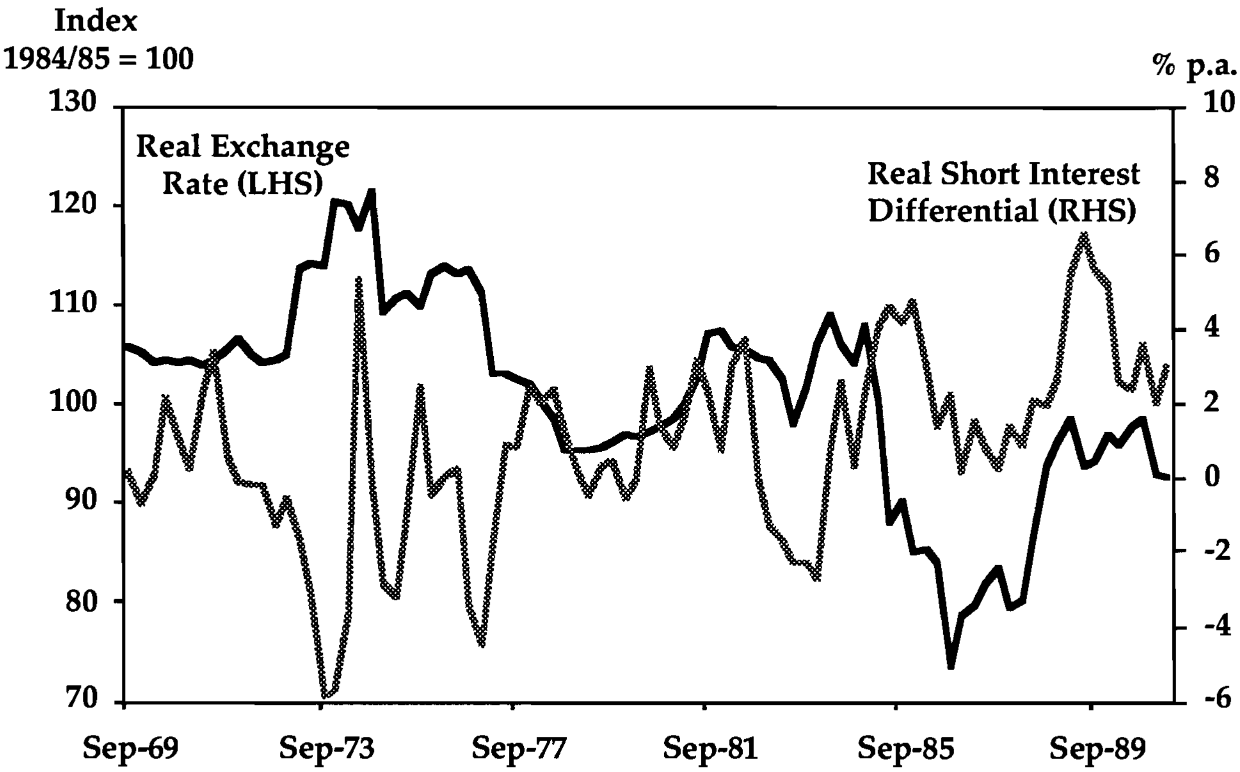
[2] All the graphs in the paper show variables in **levels**. By contrast, in all the econometric estimation, the real exchange rate and measures of the terms of trade are used in log form. We have taken considerable care to construct an accurate measure of Australia's real exchange rate. Our measure is a trade-weighted exchange rate adjusted ratio of the Australian "Medicare-adjusted" CPI to the CPIs of its 22 major trading partners. Trade weights are calculated as an average of annual trade flows from 1980 to 1989. All exchange rates are quarterly averages. See the Data Appendix for further detail.

exchange rate. However, since the early 1980s the real exchange rate appears to be more closely correlated with real interest differentials (especially with the long real differential). These graphs provide some evidence that in the most recent period, relatively high Australian real interest rates have tended to keep the real exchange rate high.

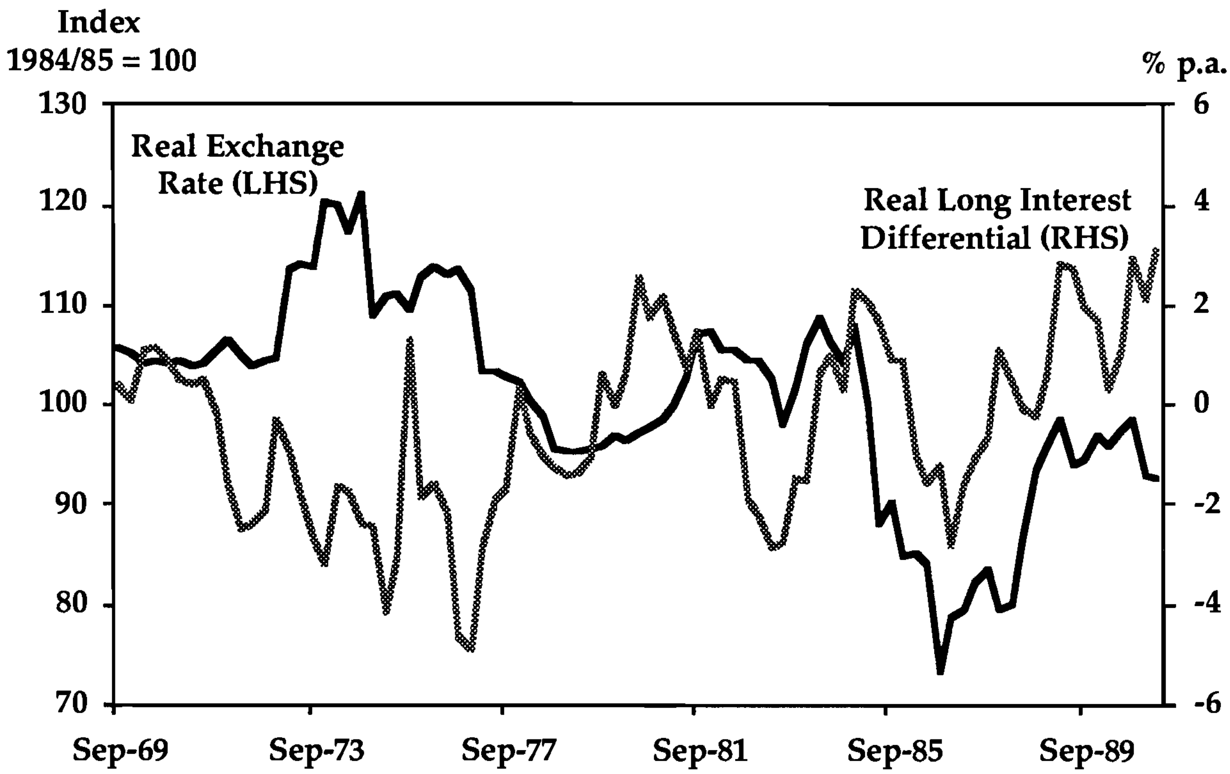
Graph 1: Real Exchange Rate and Terms of Trade
Index 1984/85 = 100



Graph 2: Real Exchange Rate and Real Short Interest Differential



Graph 3: Real Exchange Rate and Real Long Interest Differential



(a) Correlation Analysis

Table 1 shows the correlations between the variables used to explain the real exchange rate for two different sample periods. As expected, the Table shows that the real exchange rate is positively correlated with the terms of trade both over the whole sample period and since the float. It also shows that while the real exchange rate is negatively correlated with long and short real interest differentials over the whole sample period, it is positively correlated with both since the float. Note that the correlation is stronger with long differentials than with short ones after the float. Post-float, the terms of trade is also correlated with real interest differentials – and again the correlation is stronger with long real differentials than with short ones.

Table 1: Correlation Coefficients for Quarterly Data

	RER	Terms of Trade	Short Interest Differential	Long Interest Differential
1969:4 to 1990:4				
RER	1.00	0.78	-0.42	-0.27
Terms of Trade		1.00	-0.40	-0.17
Short Interest Differential			1.00	0.75
1984:1 to 1990:4				
RER	1.00	0.77	0.26	0.63
Terms of Trade		1.00	0.54	0.76
Short Interest Differential			1.00	0.66

3. METHODOLOGY

(a) Theoretical Motivation

In this paper we follow and extend the methodology of Meese and Rogoff (1988). As in their paper, we use real versions of the models proposed by Dornbusch (1976), Frankel (1979) and Hooper and Morton (1982).

There are three main assumptions behind these models. The first is that

$$E_t(q_{t+k} - \overline{q}_{t+k}) = \theta^k (q_t - \overline{q_t}), \qquad 0<\theta<1,$$

where $q_t \equiv s_t + p_t - p_t^*$.

s_t is the nominal exchange rate (foreign currency price of domestic currency), p_t and p_t^* are the domestic and foreign price levels, and q_t is the real exchange rate, all measured in logs. E_t is the time-t expectations operator, θ is the speed of adjustment parameter, and $\overline{q_t}$ is the long run real exchange rate – that is, the real exchange rate which would prevail at time t if all prices were fully flexible.

Thus after a temporary shock, the real exchange rate is assumed to move back to its long run equilibrium at a constant rate.

Secondly it is assumed that:

$$E_t(\overline{q}_{t+k}) = \overline{q_t},$$

i.e. the long run real exchange rate is assumed to follow a martingale process or a random walk. In Hooper and Morton (1982) $\overline{q_t}$ is posited to be a function of the cumulated current account deficits of both countries (which are themselves assumed to be random walks).

The third major assumption is that uncovered interest parity in its real form holds,

i.e. $$E_t(q_{t+k} - q_t) = {}_kR_t^* - {}_kR_t,$$

where ${}_kR_t$ is the domestic real interest rate of maturity k at period t, and ${}_kR_t^*$ is the corresponding foreign rate.

The combination of these three assumptions implies that the real exchange rate, q_t, is determined in the following manner:

$$q_t = \alpha({}_kR_t - {}_kR_t^*) + \overline{q_t}, \qquad \alpha = \frac{1}{1-\theta^k} > 1. \tag{1}$$

The assumptions imply that both q_t and $\overline{q_t}$ are non-stationary variables, while $({}_kR_t - {}_kR_t^*)$ is a stationary variable. Changes in the real interest differential are therefore expected to have only a temporary impact on the real exchange rate. While securities of alternative maturity affect the magnitude of α, they should not affect the underlying relationship between q_t and $\overline{q_t}$.

Equation (1) applies to the bilateral exchange rate between any two countries. As our focus is on the Australian trade-weighted real exchange rate, we study a trade-weighted version of equation (1). Rather than constructing a trade-weighted real interest differential, data limitations led us to use the differential between Australian real interest rates and an arithmetic average of real interest rates in the US, UK, Japan, and Germany.[3]

To extend this model we note that Blundell-Wignall and Gregory (1990) have shown that for a small open economy subject to terms of trade shocks, internal balance requires that the long run real exchange rate should be a function of the terms of trade. Hence in this paper, we posit that Australia's long run real exchange rate is a function of the terms of trade, that is,

$$\overline{q_t} = \gamma + \beta TOT_t, \quad \beta > 0.$$

Hence, our model for the real exchange rate is

$$q_t = \alpha({}_kR_t - {}_kR_t^*) + \overline{q_t} + u_t, \text{ where } \overline{q_t} = \gamma + \beta TOT_t. \tag{2}$$

[3] An alternative would have been to trade-weight these four foreign real rates. The resulting series are almost identical to the series we use, and the estimation results are almost unchanged.

We test four hypotheses:

1) That the real exchange rate, q_t, is non-stationary. The alternative hypothesis is that q_t is stationary and hence that the real exchange rate exhibits only temporary deviations from purchasing power parity.

2) That over the sample, there is a cointegrating relationship between the terms of trade and the real exchange rate.

3) That over the sample, there is a cointegrating relationship between the real interest differential and the real exchange rate.[4]

4) That over the sample, there is a cointegrating relationship between the real exchange rate and both the terms of trade and the real interest differential. In this case, we test if $\alpha > 1$ and $\beta > 0$ as is required by the model underlying equation (2).

(b) Econometric Methodology

Over recent years, a number of techniques have been developed to establish whether stable long run (or cointegrating) relationships exist between non-stationary variables and to estimate these relationships. Probably the most widespread method used is the Engle-Granger (1987) procedure. This has the advantage of being straightforward to apply, relying as it does on single equation OLS estimation. However, it has two main limitations.

The first is that while coefficient estimates from the Engle-Granger procedure are "super consistent" (see, for example, Pagan and Wickens (1989) for a definition), inference cannot be made on these estimates because the t-statistics do not possess a t-distribution. This limitation can be overcome by making a non-parametric adjustment to the OLS coefficient estimates and standard errors (Phillips and Hansen (1990)). This procedure yields "fully modified" coefficient estimates with their associated t-statistics.

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[4] This hypothesis requires that the real interest differential is non-stationary - which is contrary to the theory. We discuss this issue in the Discussion section.

The second limitation of both the Engle-Granger (E-G) and Phillips-Hansen (P-H) procedures arises when there are more than two variables in a system. In this case, there may be more than one cointegrating relationship between them and the E-G and P-H approaches do not provide a method of examining this issue.

By contrast, the Johansen (1988) procedure[5] addresses the problem directly. It involves applying maximum likelihood techniques to estimate a full vector autoregressive system of equations which includes both levels and first differences. Note, however, that when more than one cointegrating relationship is identified, the estimated relationships are not unique, as any linear combination of the estimated relationships is also cointegrated.[6] In this case, interpretation is not so clear.

In this paper we make use of all these techniques. Before estimating the cointegrating relationships, we examine each series to see whether it is non-stationary in a unit-root sense. Three tests are used to assess a series' stationarity. The first (the Augmented Dickey Fuller [ADF] test, Said and Dickey (1984)) and second (the Z(t) test, Phillips (1987)) assume the null hypothesis that the series is non-stationary. The third test (the G(p,q) test, Park and Choi (1988)) assumes the null hypothesis that the series is stationary. All three tests suffer from low power – that is, it is common to accept the null hypothesis even when it is false.

Having established to our satisfaction that our series are non-stationary, we use the Engle-Granger and Phillips-Hansen procedures to estimate cointegrating relationships. In all cases, we allow for a constant in the cointegrating relationship, but no time trend. We apply two tests (the ADF

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[5] See Clements (1989) for a clear description of the Johansen procedure. For applied examples, see also Johansen and Juselius (1990).

[6] Assume that x_t, y_t and z_t are three I(1) series with two independent cointegrating relationships between them: $z_t = \alpha_0 + \alpha_1 x_t + \alpha_2 y_t + \varepsilon_{1t}$ and $z_t = \beta_0 + \beta_1 x_t + \beta_2 y_t + \varepsilon_{2t}$ with errors, ε_{1t} and ε_{2t}, which are I(0). A linear combination of these equations:
$z_t = \gamma \alpha_0 + (1 - \gamma) \beta_0 + [\gamma \alpha_1 + (1 - \gamma) \beta_1] x_t + [\gamma \alpha_2 + (1 - \gamma) \beta_2] y_t + \gamma \varepsilon_{1t} + (1 - \gamma) \varepsilon_{2t}$
is also a cointegrating relationship since the error term in this new equation is a linear combination of the original I(0) errors and hence is also I(0). Thus γ can be chosen so that the coefficient on either x_t or y_t is zero. For a fully determined system, both cointegrating relationships must be specified.

and Phillips' Z(t) tests) to the residuals from the Engle-Granger regression to establish whether the series are cointegrated. Using the Phillips-Hansen estimation, we also apply Park's H(p,q) test for cointegration (Park (1988)).[7]

Finally, we apply the Johansen procedure. Where the real exchange rate is found to be cointegrated with both the terms of trade and real interest differentials, this gives us an indication of how many cointegrating relationships can be identified between the series.

(c) Series Used in Estimation[8]

Quarterly Series:

RER	log of Australia's real exchange rate with its 22 major trading partners, using quarterly average bilateral exchange rates and consumer price indices.
TOT	log of the terms of trade of goods and services.
TOT(C)	log of the ratio of the RBA Commodity Price Index (quarterly averages in $A) to the implicit price deflator for imports of goods and services. (Only available post-float.)
SHORT(F3)	expected short real interest differential using CPI inflation over the next quarter to proxy for expected inflation.
SHORT(B3)	expected short real interest differential using CPI inflation over the past quarter to proxy for expected inflation.
SHORT(B12)	expected short real interest differential using CPI inflation over the past year to proxy for expected inflation.
LONG	expected long real interest differential using CPI inflation over the past year to proxy for expected inflation.

[7] This test is similar to the G(p,q) test discussed earlier. We do not use the Durbin-Watson statistic to test for a cointegrating relationship because of its undesirable asymptotic properties (Phillips (1987)).

[8] The Data Appendix provides definitions and sources for the series.

Monthly Series[9]:

RERM, SHORTM(B12) and LONGM are the monthly series equivalent to RER, SHORT(B12) and LONG respectively. Note however, that RERM is calculated using **end month** exchange rates.

TOTM(I) log of an interpolation of the quarterly terms of trade of goods and services.

TOTM(X) log of the monthly Export Price Index deflated by an interpolation of the quarterly implicit price deflator for imports.

TOTM(C) log of the RBA Commodity Price Index (in $A) deflated by an interpolation of the quarterly implicit price deflator for imports.

Many alternative series could have been chosen for this exercise. In particular, the short and long real interest differentials can be calculated using different assumptions for inflationary expectations. A range of proxies for inflationary expectations have been proposed, from entirely backward-looking models to forward-looking models, to a mixture of the two. Campbell and Clarida (1987) compare survey data with a range of proxies for long-term inflationary expectations to illustrate how sensitive any calculation of the long-term real interest differential is to the proxy chosen. Mishkin (1987) agrees with their conclusions and states even more strongly that "research on the linkage between real interest rates and the exchange rate based on the examination of long-term real-interest differentials cannot be taken seriously."(p. 143) Nevertheless, others continue to use such proxies. In this paper, we do not add to this debate, but we do follow others (Meese and Rogoff (1988)) in our use of such proxies.[10]

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[9] When it is necessary to derive a monthly series as an interpolation of a quarterly series, part 2 of the Appendix gives details of the interpolation.

[10] The strongest justification for our use of past 12 months inflation in calculating the long real interest differential is that inflationary expectations are adaptive. Some (see, for example, Blinder (1988) and Ball (1991)) regard this as the most reasonable assumption. Of course, this assumption may affect our estimation results, which we recognise as a lack of robustness of our analysis.

A number of different terms of trade measures are also used in this paper. The differences between these series is a combination of coverage, timing and periodicity.[11] While Simes (1988) argues that the real exchange rate should be determined by expected market prices, and hence expected commodity prices, rather than the prices exporters actually receive, this should only introduce a lag between the terms of trade and the real exchange rate due to the existence of contracts. Since the focus of this paper is on long run relationships, these short-term lags are presumably of less relevance.

4. RESULTS

(a) Unit Root Tests

All series used in the estimation are tested for deterministic and stochastic non-stationarity. The results of these tests are presented in Appendix 1. For many of the series, the results are ambiguous. For some of the series, there is evidence of 2, 1 or 0 unit roots. This lack of clarity is due in part to the small sample period being used.

Our first conclusion is that the real exchange rate has one unit root. This conclusion is based on the fact that, for a majority of the three time periods studied, two of the three tests accept this hypothesis at a 1 per cent level of significance. This result confirms the results of other studies (e.g. Blundell-Wignall and Gregory (1990) and Corbae and Ouliaris (1991)) and accepts the hypothesis that the Australian real exchange rate follows a non-stationary process rather than exhibiting only temporary deviations from purchasing power parity.

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[11] The RBA Commodity Price Index includes 19 commodities and covers about sixty per cent of Australia's commodity exports, which is around forty per cent of Australia's total exports of goods and services. This series is very highly correlated with the implicit price deflator for exports of goods and services. Using quarterly data from September 1982 the correlation coefficient between the IPD for exports of goods and services and the Export Price Index is 0.998 and the correlation coefficient between the IPD for exports of goods and services and the RBA Commodity Price Index (all items in $A) is 0.976.

By the same criterion, we also find that the terms of trade and long real interest differential series have one unit root.[12]

By contrast, the short real interest differential series with three month expectations (SHORT(B3) and SHORT(F3)) appear to be stationary, that is, they have no unit roots, again looking at all the three different time periods.[13] However, the short real interest differential series with 12 month backward looking expectations (SHORT(B12)) shows much weaker evidence of being stationary than the other two series. Over the post-float period, we tentatively conclude that it has 1 unit root.

Examination of the Tables in Appendix 1 shows the low power of all three tests. Note, in particular, that the G(p,q) test only rejects the null of stationarity in one case.

In this paper we are attempting to explain long run movements in the level of Australia's real exchange rate especially over the period since the float. Since we conclude that this series is non-stationary, its long run behaviour must be explained by other non-stationary series. Hence at this point we exclude the stationary short real interest differential series (SHORT(B3) and SHORT(F3)) from any further analysis. We apply the technique of cointegration to the real exchange rate, terms of trade, long real interest differential and non-stationary short real interest differential (SHORT(B12)) series.

(b) The Real Exchange Rate and the Terms of Trade

Using quarterly data 1970 – 1988, Blundell-Wignall and Gregory (1990) find evidence for a stable long run relationship between the Australian real exchange rate and the terms of trade. Table 2 displays our results. There is

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[12] In a world with deregulated financial flows, it is hard to understand why either short or long real interest differentials could be non-stationary. This issue is addressed in the Discussion section.

[13] This statistical evidence – that real short-term interest differentials seem to be stationary – is consistent with the findings of Meese and Rogoff (1988). It is also consistent with the work of Campbell and Clarida (1987) and Tarditi and Menzies (1991) who examine the relationship between the level of real short-term interest differentials and the log level of the real exchange rate. Both of these studies find this relationship to be insignificant.

mixed evidence for a long run stable relationship between the terms of trade (TOT) and the real exchange rate. For both the long sample period (1969:4 to 1990:4) and the post-float period (1984:1 to 1990:4), the three statistical tests for cointegration give conflicting results. For both samples, the H(p,q) test accepts the null hypothesis of a cointegrating relationship, while the ADF and the Z(t) tests give mixed signals. Thus, the ADF test accepts the null of no cointegration for the long sample, but rejects it (at 10% level of significance) for the post-float period. The Z(t) test rejects the null of no cointegration (at 15% level of significance) for the long sample, but accepts it for the post-float period.

Table 2: Dependent Variable – Real Exchange Rate (quarterly)

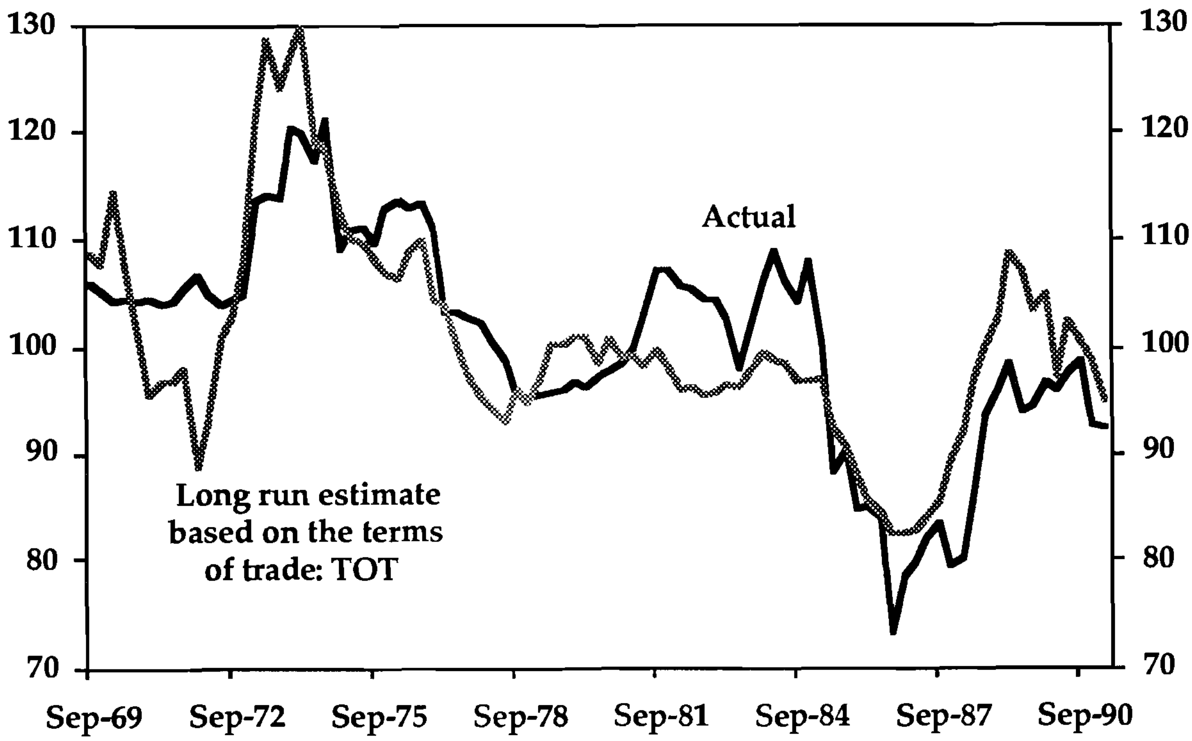
Explanator	Sample period	Coefficient Estimate			Test Statistics[a]		
		E-G	P-H[b]	t-statistic	ADF	Z(t)	H(p,q)
TOT	69:4 - 90:4	0.91	1.08	6.11	-2.13	-2.89*	6.30
	84:1 - 90:4	1.06	1.08	4.42	-3.07**	-2.61	7.84
TOT(C)	84:1 - 90:4	0.82	0.92	3.02	-2.26	-2.57	8.03
LONG	69:4 - 90:4	-0.016	-0.029	-1.84	-2.00	-2.15	5.02
	84:1 - 90:4	0.045	0.056	3.74	-3.39***	-3.52***	7.29
SHORT (B12)	69:4 - 90:4	-0.017	-0.029	4.63	-2.16	-2.45	4.35
	84:1 - 90:4	0.015	0.020	1.07	-2.09	-2.51	6.67

a. All of these statistics are based on regressions containing a constant but no time trend. The ADF and Z(t) statistics are based on the E-G regressions, while the H(p,q) statistics are based on the P-H regressions. For the ADF and the Z(t) tests, the null hypothesis is **no cointegration**. If the test statistic is **more negative** than the critical value then the null is rejected, i.e. cointegration is accepted. 5%, 10% and 15% critical values for these test statistics are: -3.37, -3.07, and -2.86 (Phillips and Ouliaris (1990)). For the H(p,q) statistic the null hypothesis is **cointegration**. Under this null, this statistic is asymptotically distributed as a chi-squared with q-p degrees of freedom. We use p = 0, q = 5. 5%, 10% and 15% critical values for this test statistic are: 11.07, 9.24 and 8.12.

b. 10 lags in the Bartlett window are used when deriving the long run variances used in the Phillips-Hansen estimates for the full sample period. 5 lags are used for the shorter sample period.

***, **, * indicates the null is rejected at a 5%, 10%, 15% level of significance.

Graph 4: The Real Exchange Rate



Interestingly, in the post-float period, two of the three tests (the ADF and Z(t) tests) find that the real exchange rate is not cointegrated with TOT(C) – the terms of trade measured by the ratio of commodity prices to import prices. This finding is surprising as most of the movement in export prices comes from changes in commodity prices, and thus there is a general expectation that the real exchange rate is primarily determined by commodity prices.

Blundell-Wignall and Gregory estimate that, in response to a 1% change in the terms of trade, the real exchange rate changes by about 0.63% pre-float, by about 1.4% post-float, and by about 1.05% for their full sample. By contrast, using a somewhat longer run of data, our estimated relationships are almost the same in the post-float period as in the whole sample. For a 1% change in the terms of trade, we estimate a change in the real exchange rate of between 0.82% and 1.08%. Graph 4 shows the **level** of the real exchange rate over the whole sample period and the long run estimate of the real exchange rate derived from the Phillips-Hansen estimated cointegrating relationship with the terms of trade (TOT).[14]

(c) The Real Exchange Rate and Real Interest Differentials

Meese and Rogoff (1988) examine several bilateral real exchange rates with the US to determine whether there is a long run stable relationship with the respective real interest differentials. They find that their short real interest differentials are stationary, and although their long real interest differentials and real exchange rate series are both non-stationary, there is no stable long run relationship between them.

Table 2 (above) contains coefficient estimates and cointegration tests for the relationship between the Australian real exchange rate and real interest differentials. As the Table shows, over the full sample period there is no strong evidence that the real exchange rate is cointegrated with either real interest rate differential, and the coefficient estimates on the interest differentials are of the wrong sign. After the float, there continues to be no strong evidence of cointegration with the short real interest differential. By contrast, there is very strong evidence of cointegration between the real exchange rate and the long real interest rate differential – with all three tests accepting cointegration at a 5% level of significance. Both the Engle–Granger and Phillips-Hansen methods give coefficient estimates on the long real interest differential of about 0.05. Because the real interest differentials are entered in levels and the real exchange rate in logs, these results imply that a 1 **percentage point** increase in the long real interest differential is associated with an average appreciation of the real exchange rate of about 5 **per cent**.

(d) Do the Terms of Trade and Real Interest Differentials Jointly Explain the Real Exchange Rate?

We have established that there is some evidence after the float of a stable relationship between the real exchange rate and the terms of trade and

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[14] The real exchange rate presumably responds more to terms of trade changes which are perceived to be permanent than to those perceived to be transitory. Our approach does not take this distinction into account.

strong evidence of a stable relationship between the real exchange rate and long real interest differentials. In this sub-section, we extend our analysis in three ways. Firstly, we examine both monthly[15] and quarterly data. Secondly, we extend our analysis to examine cointegrating relationships between all three variables. Finally, we use the Johansen (1988) procedure to examine the number of cointegrating relationships between the three series.

Tables 3 and 4 present the results from estimation on a quarterly and on a monthly basis respectively.

Test statistics using quarterly data again show much stronger evidence of cointegration when long real interest differentials are included in the relationship than when short ones are (Table 3). The evidence is strongest using TOT(C) and LONG as regressors. In this case, both the ADF and Z(t) tests reject the null of no cointegration at a 15% level of significance, while the H(p,q) test accepts the null of cointegration at a 10% level.[16] Coefficient estimates on both the terms of trade and long real interest differentials are lower than the independent estimates given in Table 2. This reflects the fact, highlighted in Table 1, that the terms of trade and the long real interest differential are highly positively correlated after the float.

Using the Johansen procedure to analyse the same data gives mixed evidence. Anywhere between 0 and 3 cointegrating relationships are identified. The existence of three cointegrating relationships implies that each series in the estimation is stationary. With the possible exception of SHORTM(B12), we have already established to our satisfaction that this is not the case. However, there may be more than one cointegrating relationship. We discuss this case later in this section.

Before doing that, we report results of monthly estimation after the float in Table 4. The evidence for a stable long run relationship between the real

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[15] Both graphical and econometric evidence suggest that the relationship between the variables is not stable for several months after the float. As a consequence, our monthly analysis is from 1984:12 to 1990:9.

[16] Note, however, that two of the three test statistics show stronger evidence of cointegration with LONG as the sole explanator, than with both LONG and TOT(C) as explanators (see Tables 2 and 3).

exchange rate and a single other variable is again stronger with the long real interest differential than with the short real differential, or with any of the measures of the terms of trade. Note that, in all the regressions which use a single explanatory variable (i.e. the first five), the explanatory variable is of the expected sign and highly significant (judged by the t-statistics on the Phillips-Hansen regressions). The coefficient estimates in these regressions are very similar to those estimated on quarterly data after the float (Table 2).

The results of single equation estimation when both the terms of trade and real interest differentials are used as explanatory variables are reported as the last six regressions in Table 4. The evidence in favour of a stable long run relationship is mixed – though it is again more favourable with long rather than short real interest differentials. The coefficient estimates are always of the expected sign.[17] With the exception of two of the regressions with the short real differential, both explanatory variables in the Phillips-Hansen regressions are highly significant.

Applying the Johansen technique to these data once again gives mixed results. Anywhere between 1 and 3 cointegrating relationships are identified. As before, we do not further analyse the case when three cointegrating relationships are identified (e.g. using TOTM(C) and LONGM as explanators). We recognise that these results contradict our original analysis of the data and are another indication of the low power of the time series tests in short runs of data.

Both panels of Graph 5 show the **level** of the monthly real exchange rate series (RERM) from December 1983 to March 1991 (i.e. since the float of the \$A). The upper panel also shows the long run real exchange rate derived

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[17] Since our real interest differentials are expressed as percent per annum, the theoretical restriction on α, the coefficient on the real interest differential in equation (2), is $\alpha > 0.01$. Based on the P-H estimates post-float, this restriction is accepted in all but one equation where long real interest differentials are used as an explanator (the exception is the first regression in Table 3). However, the restriction is accepted only in one case where short real interest differentials are used along with a terms of trade explanator. With time measured in years, the implied speed of adjustment parameter is in the range $0.5 \leq \theta \leq 0.8$. This is comparable to the Meese and Rogoff (1988) and Campbell and Clarida (1987) results.

from the Phillips-Hansen estimated cointegrating relationship (P-H ECR) with the terms of trade (TOTM(X)), while the lower panel shows the long run real exchange rate derived from the P-H ECR with the long real interest differential (LONGM). Graph 6 shows the level of the monthly real exchange rate as well as the long run real exchange rate derived from the P-H ECR with **both** the terms of trade (TOTM(X)) and the long real interest differential (LONGM). The three cointegrating regressions used for Graphs 5 and 6 are reported in Table 4. Note that they are all estimated for the period 1984:12 to 1990:9, i.e. the period between the vertical lines on the graphs.

The cointegrating relationship used for Graph 6 is chosen because of its appealing properties. At a 5% level of significance, the Z(t) statistic rejects the null of no cointegration, the H(p,q) test accepts the null of cointegration and both tests from the Johansen method accept the hypothesis of a single cointegrating relationship between these two variables and the real exchange rate.[18] Finally, the t-statistics for the regression imply that both variables are highly significant.

Interestingly, in Graph 6, the eight months after October 1987 stand out as the longest time in the estimation period during which the real exchange rate deviates in one direction from its long run estimate. An obvious reason for this behaviour of the exchange rate was the world stockmarket crash. At the time, this was widely expected to lead to a world-wide recession and hence a fall in commodity prices. In fact, the event preceded a world-wide boom in 1988.

For completeness, Graph 7 shows two Phillips-Hansen estimated cointegrating relationships using SHORTM(B12). Note that the evidence from Table 4 suggests that the relationship shown in the bottom panel of Graph 7 **is not** a stable long run relationship.

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[18] Unfortunately, the ADF statistic accepts the null of no cointegration, but you can't have everything.

Table 3 [a]: Dependent Variable: Real Exchange Rate (quarterly 1984:1 - 1990:4)

	SINGLE EQUATION ESTIMATES						JOHANSEN[b]			
	Coefficient Estimates[c]			Test Statistics[d]			Coefficient Estimates		Tests for the number of cointegrating relationships	
	E-G	P-H	t-statistic	ADF	Z(t)	H(p,q)	1st Vector	2nd Vector	Lambda Max 5% (10%)	Trace 5% (10%)
TOT LONG	0.95 0.007	0.97 0.004	2.73 0.21	-3.40*	-2.82	7.32	1.30 0.04	-0.41 0.06	0 (3)	3 (3)
TOT(C) LONG	0.43 0.029	0.43 0.031	1.21 1.66	-3.42*	-3.39*	8.37*	-1.42 0.20	3.39 -0.14	2 (3)	2 (3)
TOT SHORT (B12)	1.23 -0.014	1.24 -0.019	6.26 -2.20	-2.02	-2.08	5.84	1.32 -0.003	1.01 -0.037	2 (2)	2 (2)
TOT(C) SHORT (B12)	0.84 -0.003	0.95 -0.010	3.11 -0.75	-2.09	-2.47	5.99	0.49 0.047	2.57 -0.119	3 (3)	3 (3)

a. In this Table and in Table 4, the relationships between the real exchange rate and its determinants are estimated using three methods. In the first part of the Table are the single equation estimates - based on the Engle-Granger (E-G) and Phillips-Hansen (P-H) approaches. In the second part of the Table are the Johansen estimates - based on the estimation of the full system of equations.

b. The Johansen procedure is applied to the quarterly data with 4 lags, and to the monthly data with 12 lags. These long lag lengths are necessary to remove problems of non-normality from most of the equations. Only the first two cointegrating relationships are reported. In the results of two different tests for the number of cointegrating relationships, "0" indicates the test statistic rejects the null hypothesis of one cointegrating relationship, and hence accepts the alternative of less than one. Similarly, "1" indicates that two cointegrating relationships can be rejected, but one cannot be.

c. Two coefficient estimates are reported. E-G, the Engle-Granger estimates are OLS coefficient estimates. P-H, the Phillips-Hansen estimates are calculated by making a non-parametric adjustment to the OLS estimates, as described in Hansen (1990). This adjustment ensures that the t-statistics have an asymptotic t-distribution. 5 lags in the Bartlett window are used when deriving the long run variances used in the Phillips-Hansen estimates.

d. The ADF and Z(t) test statistics are again based on the E-G regressions, while the H(p,q) test statistics are based on the P-H regressions. For the ADF and the Z(t) tests, the null hypothesis is **no cointegration**. If the test statistic is **more negative** than the critical value, the null is rejected, i.e. cointegration is accepted. Critical values also depend on the number of explanators in the regression. 5%, 10% and 15% critical values for these test statistics given one (two) explanators are: -3.37 (-3.77), -3.07 (-3.45), and -2.86 (-3.26). (Phillips and Ouliaris (1990)). For the H(p,q) statistic the null hypothesis is **cointegration**. This statistic is asymptotically distributed as a chi-squared with q-p = 5 degrees of freedom. 5%, 10% and 15% critical values for this test statistic are: 11.07, 9.24 and 8.12.

***, **, * indicates the null is rejected at a 5%, 10%, 15% level of significance.

Table 4: Dependent Variable: Real Exchange Rate (monthly 1984:12 - 1990:9)

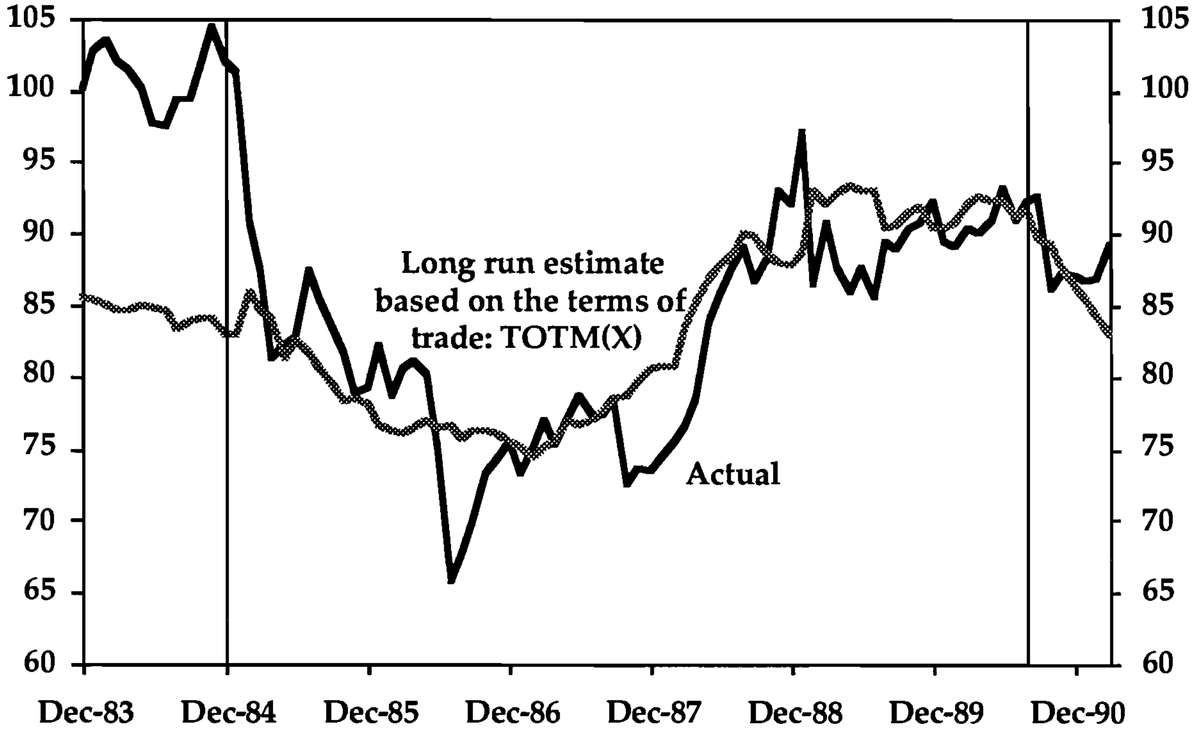
	SINGLE EQUATION ESTIMATES						JOHANSEN			
	Coefficient Estimates			Test Statistics			Coefficient Estimates		Tests for the number of cointegrating relationships	
	E-G	P-H	t-statistic	ADF	Z(t)	H(p,q)	1st Vector	2nd Vector	Lambda Max 5% (10%)	Trace 5% (10%)
TOTM(I)	0.93	0.94	8.67	-2.46	-4.22***	9.75**				
TOTM(C)	0.74	0.85	6.29	-1.85	-3.91***	14.61***				
TOTM(X)	0.78	0.82	7.73	-2.69	-4.54***	13.79***				
LONGM	0.040	0.054	5.98	-2.77	-3.45***	2.60				
SHORTM (B12)	0.025	0.038	3.25	-1.53	-3.30**	6.20				

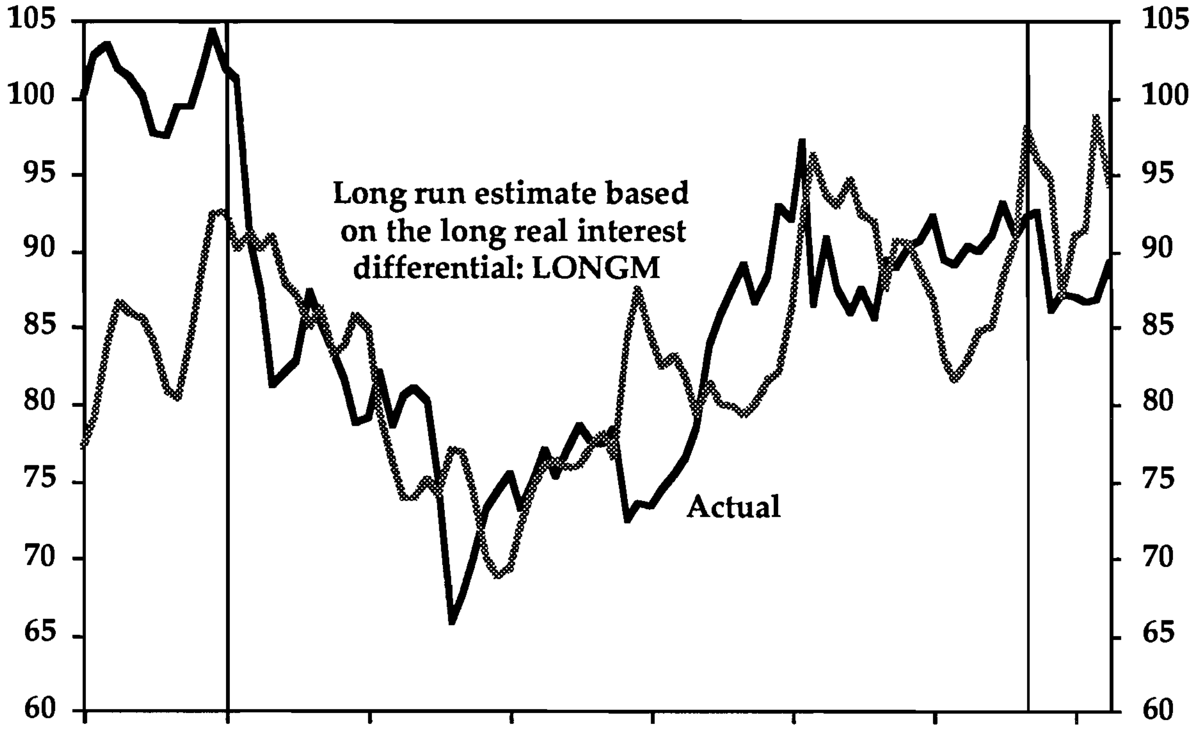
Table 4 (continued)

	E-G	P-H	t-statistic	ADF	Z(t)	H(p,q)	1st Vector	2nd Vector	Lambda Max 5% (10%)	Trace 5% (10%)
TOTM(I)	0.86	0.70	4.27				0.21	1.59		
LONGM	0.004	0.015	1.79	-2.59	-4.20***	7.21	0.042	-0.014	2 (2)	2 (2)
TOTM(C)	0.43	0.35	2.43				0.30	0.14		
LONGM	0.023	0.033	4.22	-2.67	-3.90***	14.43***	0.048	0.030	3 (3)	3 (3)
TOTM(X)	0.62	0.51	3.79				0.34	1.26		
LONGM	0.013	0.024	3.02	-2.96	-4.39***	10.02**	0.033	-0.002	1 (2)	1 (2)
TOTM(I)	0.91	0.90	6.91				0.49	3.55		
SHORTM (B12)	0.001	0.002	0.41	-2.50	-4.23***	9.93**	0.048	-0.178	2 (2)	2 (2)
TOTM(C)	0.63	0.68	4.86				0.12	2.16		
SHORTM (B12)	0.011	0.014	2.04	-2.28	-4.07***	16.52***	0.068	-0.063	3 (3)	3 (3)
TOTM(X)	0.72	0.71	6.30				0.05	0.91		
SHORTM (B12)	0.006	0.009	1.55	-2.93	-4.63***	12.63***	0.069	-0.019	3 (3)	3 (3)

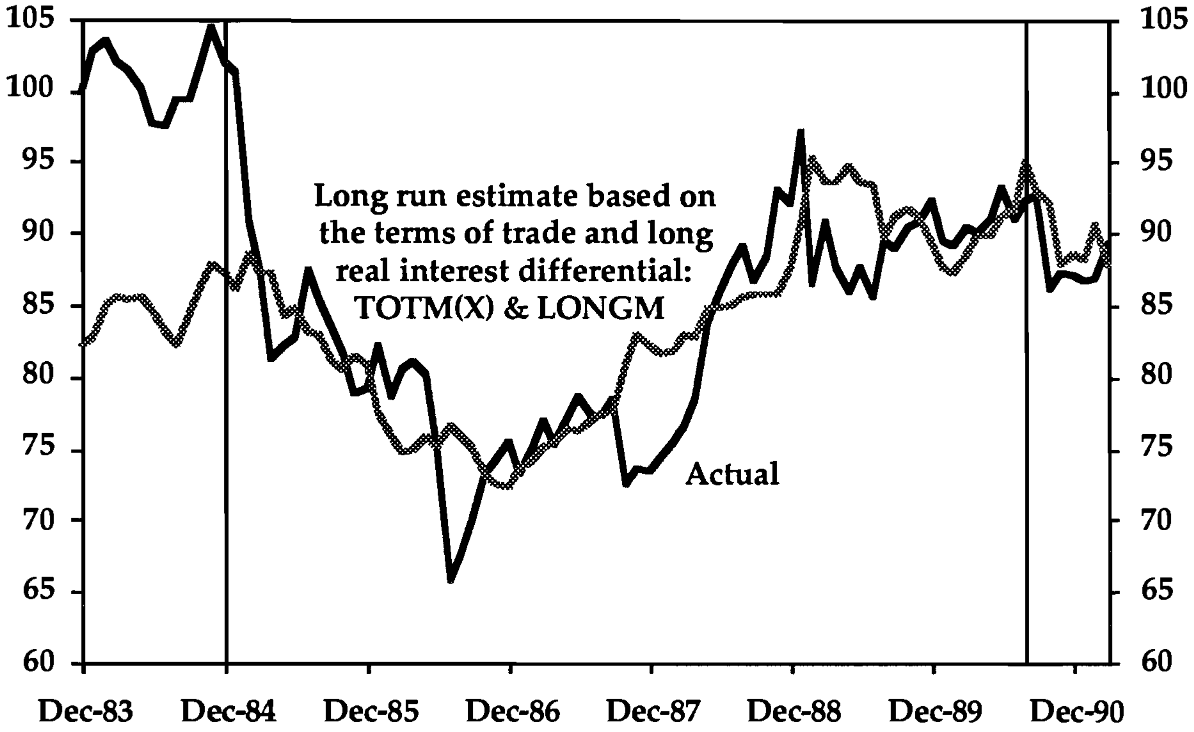
For an explanation of this Table see the notes to Table 3.

Graph 5: The Real Exchange Rate

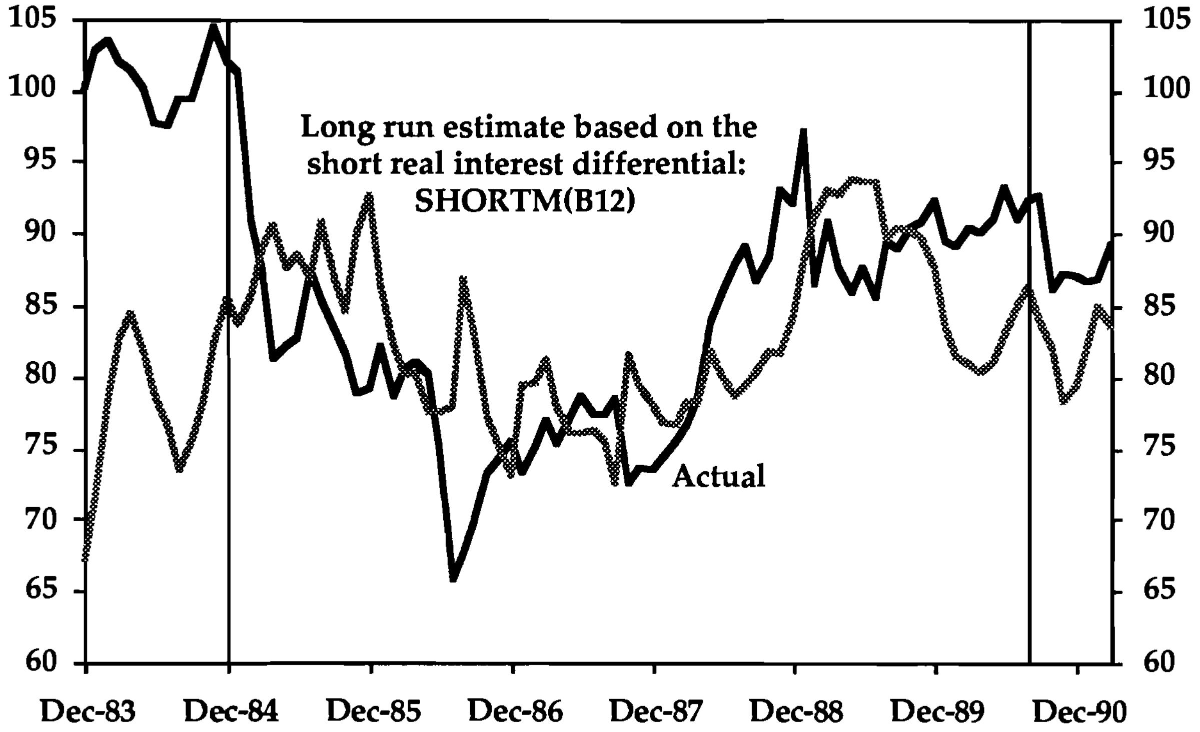


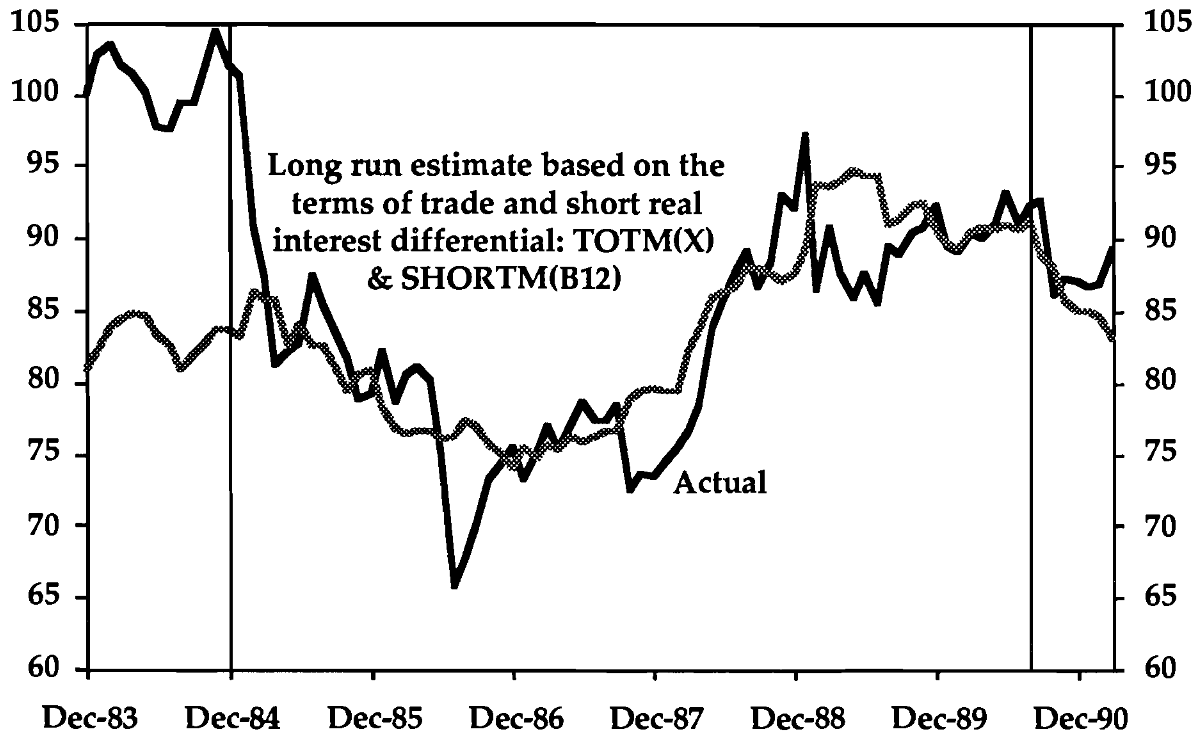


Graph 6: The Real Exchange Rate



Graph 7: The Real Exchange Rate





We now briefly discuss the out-of-sample behaviour of the estimated relationships used for Graphs 5, 6 and 7. For the year 1983:12 to 1984:11 (before the estimation period), all the estimated long run real exchange rate series very substantially underestimate the actual real exchange rate.

Interestingly, they share this property with the cointegrating relationship estimated between the terms of trade and the real exchange rate for the period 1969 to 1990 and shown in Graph 4. Thus, during the first year after the float, all our evidence suggests that the Australian real exchange rate was well above the level consistent with either the terms of trade or the relative level of real interest rates.

Table 5: Johansen Coefficient Estimates

Explanators	Original Estimates		Transformed Coefficient Estimates	
	First Vector	Second Vector	First Vector	Second Vector
Quarterly 84:1 - 90:4				
TOT(C)	-1.42	3.39	1.41	0.00
LONG	0.20	-0.14	0.00	0.099
TOT	1.32	1.01	1.35	0.00
SHORT(B12)	-0.003	-0.037	0.00	-0.148
Monthly 84:12 - 90:9				
TOTM(I)	0.21	1.59	1.25	0.00
LONGM	0.042	-0.014	0.00	0.051
TOTM(X)	0.34	1.26	1.21	0.00
LONGM	0.033	-0.002	0.00	0.046
TOTM(I)	0.49	3.55	1.14	0.00
SHORTM(B12)	0.048	-0.178	0.00	0.084

In the six months 1990:10 to 1991:3 (after the estimation period), it appears that the cointegrating relationship based on both the terms of trade and the long real interest differential has performed better than either of the relationships using a single explanator (compare Graph 6 with Graph 5) or than the relationships estimated with the short real differential (Graph 7).

There are five sets of variables from Tables 3 and 4 for which the Johansen method suggests that there may be two cointegrating relationships. In these cases, interpretation is not straightforward because the estimated cointegrating relationships are not unique – any linear combination of two cointegrating vectors also represents a cointegrating relationship (see footnote 6). One way to interpret the results is to transform the vectors to relationships involving only one of the explanators. In this way, the coefficient estimates can be compared with the independently derived single equation estimates presented in Tables 2 and 4. For these five cases, the original Johansen coefficient estimates and the transformed estimates are reported in Table 5. It is encouraging that, with one exception, the transformed coefficient estimates are all of the expected sign. They are of the same order of magnitude as the single equation estimates from Tables 2 and 4 (although, in all cases, the transformed estimates of the terms of trade are larger).

5. DISCUSSION AND CONCLUSIONS

Four broad conclusions can be drawn from our results. First, they confirm the results of other studies (e.g. Blundell-Wignall and Gregory (1990), Corbae and Ouliaris (1991)) and accept the hypothesis that the Australian real exchange rate is non-stationary, rather than deviating only temporarily from purchasing power parity. However, we should again emphasise the low power of our statistical tests. For all the sample periods examined, **both** the null hypothesis of non-stationarity (using the ADF and Z(t) tests) and the null of stationarity (using the G(p,q) test) are accepted by the data.

Second, graphical analysis over the period 1969 – 1990 supports the evidence of Blundell-Wignall and Gregory (1990) in suggesting that there is a stable long run relationship between the Australian real exchange rate and the terms of trade (Graph 4). Perhaps surprisingly, the econometric tests provide only weak evidence supporting the existence of this stable long run relationship (Table 2). A possible explanation for this result is that other non-stationary variables are missing from the relationship. Plausible candidates are relative productivity growth and net foreign asset accumulation – both of which should have a longer-term impact on the real exchange rate. We briefly discuss these influences at the end of this section.

Third, over the period 1969 – 1990, the evidence does not suggest that short or long-term real interest differentials contribute to any stable long run relationship with the real exchange rate (Table 2).

Fourth, after the float of the $A and the world-wide deregulation of financial markets, there is evidence that real interest differentials do contribute to the behaviour of the real exchange rate. Although there is some evidence that short-term real differentials contribute to a stable relationship with the real exchange rate, the evidence is much stronger that long-term real interest differentials make a contribution.

After the float, three pieces of evidence which suggest that long-term real interest differentials contribute to a stable relationship with the real exchange rate are:

(i) There is stronger evidence for a cointegrating relationship between the real exchange rate and the long real interest differential on its own than between the real exchange rate and either short real differentials or the terms of trade on their own.

(ii) There is good evidence of one (and sometimes more than one) stable relationship between the real exchange rate, the terms of trade and the long real interest differential. For most specifications, the t-statistics associated with the Phillips-Hansen estimates imply that the coefficient on the real long interest differential is statistically significant.

(iii) For the six months after the end of the monthly estimation period, the long run relationship estimated using both TOTM(X) and LONGM seems to better explain the real exchange rate than either of these variables on their own or than SHORTM(B12) – see Graphs 5 – 7.

After the float, the evidence suggests that both the terms of trade and long real interest differentials contribute to a stable relationship with the real exchange rate. To estimate the magnitude of their influence on the real exchange rate, we use our preferred estimated relationship – with TOTM(X) and LONGM (Table 4 and Graph 6), and examine both Phillips-Hansen and Johansen estimates. Best estimates are that a 1% improvement in the terms of trade leads to an appreciation of the Australian real exchange rate of about 0.3 to 0.5%, while an increase of 1 percentage point in the differential between Australian and world long real interest rates is associated with an appreciation of the Australian real exchange rate of about 2 to 3 1/2%.

Table 6: Comparison of Changes to the Real Exchange Rate – Actual and Long Run Estimate [a]

	Dec 84 - Dec 86	Dec 86 - Sep 90
Actual Real Exchange Rate (% change)	-25.9	22.7
Long run estimate (% change)	-17.1	28.5
Percentage points of change contributed by:[b]		
Terms of trade (TOTM(X))	-5.8	11.3
Long real interest differential (LONGM)	-12.0	15.5

a. Our preferred estimated relationship – with TOTM(X) and LONGM as explanators – is used in this simulation.

b. The percentage points of change contributed by an explanator is calculated by simulating the long run model while holding the other explanator fixed at its initial level. As the model is non-linear in levels, the sum of individual contributions is not equal to the total estimated long run change.

Table 6 shows the size of changes in both the real exchange rate and simulations of our preferred long run relationship over the estimation period for the monthly data. As the Table illustrates, with this specification, long real interest differentials contributed more than the terms of trade to changes in the real exchange rate over both its depreciation and subsequent appreciation. However, an examination of Tables 3 and 4 demonstrates that different specifications yield significantly different coefficient estimates and hence relative contributions to exchange rate changes.

It is widely recognised that distinguishing between borderline stationary and non-stationary variables is a difficult exercise which is best attempted with long runs of data (see, for example, Frankel and Meese (1987)). From a theoretical perspective, in a world with deregulated financial flows, it is hard to understand how either short or long real interest differentials could be non-stationary. This would imply that real interest differentials **should not appear** in the long run relationship with the real exchange rate. The "true" long run (cointegrating) relationship would then be:

$$q_t = a + bTOT_t + v_t, \tag{2a}$$

where a and b are positive constants and v_t is a stationary error.

If equation (2a) represents the true long run model (with real interest differentials stationary), OLS estimation of equation (2a) will generate a "super-consistent" estimate of b. Under the same assumptions, the OLS estimate $\hat{\beta}$ derived from estimation of equation (2)

$$q_t = \alpha({}_kR_t - {}_kR_t^*) + \gamma + \beta TOT_t + u_t \tag{2}$$

is also a super-consistent estimate of b.[19] However, in small samples, there is no guarantee that equations (2) and (2a) will generate similar estimates of b. This point is highlighted by a comparison of the coefficient estimates on the terms of trade from estimates of the two equations. For example, from Table 4, estimation of equation (2a) using TOTM(X) gives $\hat{b} = 0.78$ (E-G) or 0.82 (P-H), while estimation of equation (2) using TOTM(X) and LONGM gives $\hat{\beta} = 0.62$ (E-G) or 0.51 (P-H) or 0.34

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19 As long as $({}_kR_t - {}_kR_t^*)$ is uncorrelated with u_t, OLS estimation of (2) also yields a consistent estimate of α.

(Johansen). Thus, over the sample period, including the long real interest differential in the regression substantially reduces the estimated influence of the terms of trade on the real exchange rate.

As discussed in Section 3(a) of this paper, economic theory implies that equation (2) is the correct specification. Given the short run of data, this leads us to have more confidence in the estimates derived from this equation than from equation (2a).

One can accept both the theoretical argument that long real interest differentials are stationary, as well as the empirical evidence that shocks to the long real interest differential persist for long enough to make the series appear non-stationary. The economic relevance of this argument is that while the real interest differential should not have a permanent effect on the level of the real exchange rate, its effect can last for an extended period – long enough to influence resource allocation between the traded and non-traded sectors of the economy.[20]

Note that the empirical observation that long real interest differentials appear non-stationary is not peculiar to Australia nor to our analysis being in trade-weighted terms. Bilateral studies of large OECD countries in Meese and Rogoff (1988) and Blundell-Wignall and Browne (1991) also come to this conclusion. There are two reasons why long real interest differentials may exhibit such strong persistence. The first relies on the Dornbusch (1976) argument that goods prices are sticky. Secondly, if monetary policy changes are not fully credible and/or if expectations are partly backward-looking, long real interest rates (as we measure them) may take considerable time to adjust to a change in the underlying inflation rate.

There is quite a strong correlation between short and long real interest differentials (with a correlation coefficient of 0.66 over the post-float period – see Table 1). So, high short real differentials are mostly associated with high long real differentials. Despite this fact, the relationships

[20] Meese and Rogoff (1988) point out that even if the series are borderline stationary, cointegration tests can still be meaningful, since they essentially test whether the large variance components of the different series effectively cancel each other, leaving a residual with only a small variance.

estimated in this paper are much more convincing with long rather than with short real interest differentials.

A possible explanation for the unsatisfactory results using short real differentials is that short-term nominal (and hence in a world with sticky inflation, real) interest rates are set by the authorities to achieve domestic economic objectives which change over time. Macfarlane and Tease (1989) point out that as well as having a medium-term inflation objective, short-term interest rates are used as a counter-cyclical stabilization tool and at times they have been used explicitly to support the exchange rate. These different roles for short interest rates presumably make it very difficult to uncover a stable relationship between short real interest differentials and the real exchange rate.[21]

The results in this paper are also consistent with those of Meese and Rogoff (1988) who find that their regressions (in first-difference form) are better with long real differentials rather than with short ones. Finally on this point, note that if inflationary expectations are rational, SHORT(F3) is an unbiased estimate of the expected short-term real interest differential. Since both our evidence and that of Meese and Rogoff is that SHORT(F3) is a stationary variable, it cannot (at least not on its own) form a stable long run relationship with the non-stationary real exchange rate.

To conclude, we briefly mention two further important determinants of the real exchange rate. First, over the longer run, inter-country differences in productivity growth make a profound difference to bilateral real exchange rates (see, for example, Dornbusch (1988)). The evidence of Broadbent (1991) and Lowe (1991) implies that labour productivity growth in the Australian traded sector in the 1970s and 1980s was significantly slower than the average labour productivity growth in the traded sectors of our major trading partners. Other things equal, this lower productivity growth implies a secular decline in the Australian real exchange rate.

Second, theory implies that other things equal, an increase (decrease) in a country's net holdings of foreign assets leads to an appreciation (a

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[21] Simes (1988) points out that this policy reaction function leads to a bias to the OLS coefficient estimate on the real interest differential.

depreciation) of the domestic real exchange rate (see Dornbusch and Fischer, 1980, Meese and Rogoff, 1988). Hence, the increase in the ratio of Australia's net external liabilities to GDP from about 20% to about 40% over the 1980s should have put some downward pressure on the real exchange rate.

Empirically however, the link from a country's net foreign asset position to its real exchange rate appears to be a weak one. In the regressions run by Meese and Rogoff (1988), the estimated coefficient on the cumulated trade balance is of the wrong sign in four cases out of six (and always statistically insignificant). By contrast, the results of Blundell-Wignall and Browne (1991) are more encouraging – with the estimated coefficient on the cumulated current account of the correct sign in all cases. Interestingly for our purposes, during the financially deregulated 1980s, the cumulated current account had **less than half** the effect on the real exchange rate as in the more financially regulated 1970s. From their estimates, the increase – by 20% of GDP – in the Australian cumulated current account deficit during the 1980s should have been associated with a real depreciation of 4.4%.[22] The results in Table 6 imply that other influences – that is, terms of trade and real interest rate changes – had a substantially larger effect on the real exchange rate during the 1980s than this. Hence, it may prove difficult to isolate the effect of the cumulated current account deficit on the real exchange rate for Australia over the medium term.

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22 Derived as an average of results for Japan/US, Germany/US and Germany/UK from Table 4 of Blundell-Wignall and Browne (1991).

DATA APPENDIX

1. Quarterly Data

The Real Exchange Rate

The real exchange rate is a trade-weighted exchange rate adjusted ratio of the Australian "Medicare adjusted" Consumer Price Index (CPI) to the CPIs of its 22 major trading partners. Trade weights have been calculated as an average of annual trade flows over the period from 1980 to 1989.

All exchange rates are quarterly averages. Exchange rate and CPI data has primarily been collected from the IMF's International Financial Statistics (IFS) with a few major exceptions. For Australia, a "Medicare adjusted" CPI series is used. Data for Taiwan is collected from Financial Statistics, Taiwan District, Republic of China. Exchange rate data for Hong Kong is collected from the International Department's Dealing Room, Reserve Bank of Australia (RBA). Where CPI data is not available (for example, in the most recent quarters for some of the smaller countries) estimates have been taken from a variety of sources.

Terms of Trade

Export and import implicit price deflators for goods and services are taken from Balance of Payments, Australia, Quarterly ABS Publication, Catalogue No. 5302.0.

Quarterly averages of the Reserve Bank Commodity Price Index, published in the RBA Bulletin, are used in calculating TOT(C).

Real Interest Rates

All real interest differentials used in the paper are calculated as the difference between Australian real interest rates and an arithmetic average of real interest rates in the US, UK, Japan and Germany. Real interest rates are calculated by adjusting annualised nominal interest rates for annualised inflationary expectations. Inflationary expectations calculations are based on CPIs.

Short Nominal Interest Rates
US, UK and Australia - quarterly averages of monthly average three month treasury bill rates published in the RBA Bulletin.

Germany - quarterly averages of the end-month 3mth Fibor rate from the OECD Main Economic Indicators (MEI).

Japan - quarterly average of end-month 3mth Gensaki rate from OECD MEI.

Because the Japanese short nominal interest rate is only available from 1978, the short real interest differentials before 1978 compare Australia with only the US, the UK, and Germany.

Long Nominal Interest Rates
Quarterly averages of monthly average data are used for all countries except Australia where quarterly averages of the last trading day of each month are used. All data is taken from the RBA Bulletin Database. The series used are:

US - Government security yields greater than 10 years
UK - Government security yields of 10 years
Germany - Public sector bond yields 7-15 years
Japan - Central government bond yields
Australia - 10 year bonds.

2. Monthly Data

Real Exchange Rate
The monthly real exchange rate is calculated using the same methodology, data sources and trade weights as the quarterly real exchange rate. Note, however, that end month exchange rates are used.

Monthly consumer prices series are taken from the IFS for all countries except:
• Australia where the ABS published "Medicare adjusted" CPI series is interpolated to a monthly series.

• PNG and NZ where quarterly series from IFS are interpolated to derive monthly series.
• Taiwan where monthly data is taken from the Financial Statistics, Taiwan District, Republic of China.
• China where the annual series from IFS is interpolated to derive a monthly series.

For recent months, these IFS CPI statistics have been updated from a variety of sources for many of the Asian countries.

Terms of Trade
The RBA Commodity Price Index, all items, in $A is taken from the RBA Bulletin Database.

The Export Price Index, published in ABS Cat. No. 6405.0, is based 1989/90. Prior to July 1989, this series is spliced with 1979/80 based series.

Quarterly export and import implicit price deflators for goods and services (see reference above) are interpolated to form monthly series. For the most recent three months, the series are linearly smoothed.

Short Nominal Interest Rates
US, UK, Japan and Germany - 3 month Eurocurrency rates taken at last trading day of month from International Department, RBA.

Australia - 90 day bill rate - taken at last trading day from Domestic Markets Department, RBA.

Long Nominal Interest Rates
Taken from RBA Bulletin. Monthly average data is used for all countries except Australia where last trading day of month is used.

US - Government security yields greater than 10 years
UK - Government security yields of 10 years
Germany - Public sector bond yields 7-15 years
Japan - Central government bond yields
Australia - 10 year bonds.

APPENDIX

1. Unit Root Tests

The following tables contain the results of the unit root tests, divided up according to time period and frequency. The data from each series is assumed to be drawn from a model of the form

$$y_t = \alpha + \beta^* y_{t-1} + \varepsilon_t ,$$

where ε_t has an ARMA(l,m) distribution. If $\beta < 1$, the series is stationary, while if $\beta = 1$, it is non-stationary (integrated of order 1). Three tests on each series are reported.

ADF is the augmented Dickey-Fuller test, as described in Said and Dickey (1984). Four lags on the differenced series (4 AR corrections) are included to absorb serial correlation. The null hypothesis for this test is non-stationarity.

Z(t) is the test proposed in Phillips (1987). This test involves making non-parametric adjustments to the ADF test. The null hypothesis for this test is also non-stationarity.

Since both the ADF and the Z(t) tests are widely recognised as having low power, we follow the recommendation of Pagan and Wickens (1989) and quote the coefficient estimate $\hat{\beta}$ as well as its test statistic, to give an indication of the estimated structural relationship.

G(p,q) is the test proposed by Park and Choi (1988). Unlike the other two tests, the null hypothesis for this test is stationarity. For all the tests in this paper, p = 0, and q = 5. Asymptotically, the G(p,q) statistic has a $\chi^2(q-p)$ distribution.

For both the Z(t) and the G(p,q) test, we follow Phillips (1987) in using a Bartlett Window when estimating the long run variance. 10 lags on this window are used for both the quarterly and monthly data over the full sample periods, while only 5 lags were used for the quarterly over the shorter sample period.

***, **, * indicates the null is rejected at a 1%, 5%, 10% level of significance. (The relevant critical values are -3.43, -2.86, and-2.57 (Fuller, 1976) for the ADF and Z(t) tests and 15.09, 11.07 and 9.24 for the G(p,q) tests.)

QUARTERLY 1969:4 - 1990:4

SERIES	RER		TOT		LONG	
	Coeff Est	**Test Stat**	**Coeff Est**	**Test Stat**	**Coeff Est**	**Test Stat**
Level						
ADF	0.93	-1.54	0.87	-3.23**	0.83	-1.98
Z(t)	0.94	-1.87	0.93	-2.14	0.79	-3.02**
G(p,q)		6.30		5.53		7.65
First Difference						
ADF	0.26	-3.12**	0.38	-3.70***	-0.52	-5.18***
Z(t)	0.02	-8.88***	0.24	-7.26***	-0.08	-11.12***
G(p,q)		0.81		2.53		3.85

SERIES	SHORT (B3)		SHORT (F3)		SHORT (B12)	
	Coeff Est	**Test Stat**	**Coeff Est**	**Test Stat**	**Coeff Est**	**Test Stat**
Level						
ADF	0.48	-2.47	0.48	-2.76*	0.70	-2.72*
Z(t)	0.17	-7.98***	0.21	-7.87***	0.69	-3.91***
G(p,q)		7.21		7.47		6.88
First Difference						
ADF	-2.25	-5.81***	-2.15	-5.88***	-0.68	-4.90***
Z(t)	-0.28	-22.95***	-0.38	-26.09***	-0.04	-11.95***
G(p,q)		7.69		4.42		4.79

QUARTERLY 1984:1 - 1990:4

	RER		TOT		TOT (C)		LONG	
	Coeff Est	Test Stat	Coeff Est	Test Stat	Coeff Est	Test Stat	Coeff Est	Test Stat
Level								
ADF	0.76	-1.71	0.81	-2.48	0.75	-2.48	0.78	-1.11
Z(t)	0.82	-2.10	0.93	-1.60	0.90	-1.66	0.75	-1.93
G(p,q)		6.40		5.74		5.86		6.07
First Difference								
ADF	0.40	-1.69	0.64	-1.51	0.46	-1.52	-0.36	-2.29
Z(t)	0.02	-5.02***	0.49	-3.03**	0.30	-3.81***	-0.06	-5.65***
G(p,q)		5.17		5.74		6.62		4.22

	SHORT (B3)		SHORT (F3)		SHORT (B12)	
	Coeff Est	Test Stat	Coeff Est	Test Stat	Coeff Est	Test Stat
Level						
ADF	0.31	-1.41	0.35	-1.57	0.57	-1.97
Z(t)	0.10	-4.55***	0.32	-3.66***	0.66	-2.59*
G(p,q)		7.14		7.75		6.05
First Difference						
ADF	-2.25	-2.80*	-2.08	-2.89**	-0.53	-2.89**
Z(t)	-0.39	-8.84***	-0.24	-8.01***	-0.17	-6.21***
G(p,q)		12.14**		5.12		5.96

MONTHLY 1984:12 - 1990:9

SERIES	RERM		TOTM (I)		TOTM (C)		TOTM (X)	
	Coeff Est	Test Stat	Coeff Est	Test Stat	Coeff Est	Test Stat	Coeff Est	Test Stat
ADF	0.95	-0.95	0.98	-1.30	0.95	-1.22	0.98	-0.70
Z(t)	0.89	-2.53	1.00	-1.03	0.96	-1.40	0.99	-0.74
G(p,q)		7.49		6.97		6.96		6.98
First difference								
ADF	-0.25	-4.34***	0.74	-2.11	0.15	-3.13**	0.44	-2.56
Z(t)	-0.02	-8.47***	0.61	-4.29***	-0.02	-8.51***	0.09	-7.85***
G(p,q)		6.95		6.85		8.57		6.96

SERIES	LONGM		SHORTM (B3)		SHORTM (F3)		SHORTM (B12)	
	Coeff Est	Test Stat	Coeff Est	Test Stat	Coeff Est	Test Stat	Coeff Est	Test Stat
Level								
ADF	0.91	-1.81	0.66	-3.17**	0.66	-3.03**	0.87	-1.83
Z(t)	0.94	-1.67	0.74	-2.88**	0.74	-2.88**	0.86	-2.43
G(p,q)		7.45		5.60		7.02		6.77
First difference								
ADF	0.06	-3.87***	-0.41	-4.83***	-0.32	-3.99***	-0.27	-3.71***
Z(t)	0.25	-6.31***	0.21	-7.26***	0.11	-8.28***	-0.01	-8.57***
G(p,q)		5.52		1.23		5.72		4.47

2. Interpolation of Quarterly-Period-Average Data

Define x(t) as the estimate for the series in quarter t; a(t), b(t), and c(t) as the consecutive monthly interpolated estimates in quarter t.

Define d(t,t+1) as the average of the current and next quarterly estimates,

i.e. $$d(t,t+1) \equiv \frac{x(t) + x(t+1)}{2}. \qquad \text{(A1)}$$

Assumptions:

1) The arithmetic average of the interpolated three months is the quarterly estimate,

i.e. $$a(t) + b(t) + c(t) = 3x(t); \qquad \text{(A2)}$$

2) At the end of each quarter, the line joining the monthly interpolated estimates passes through d(t,t+1); and

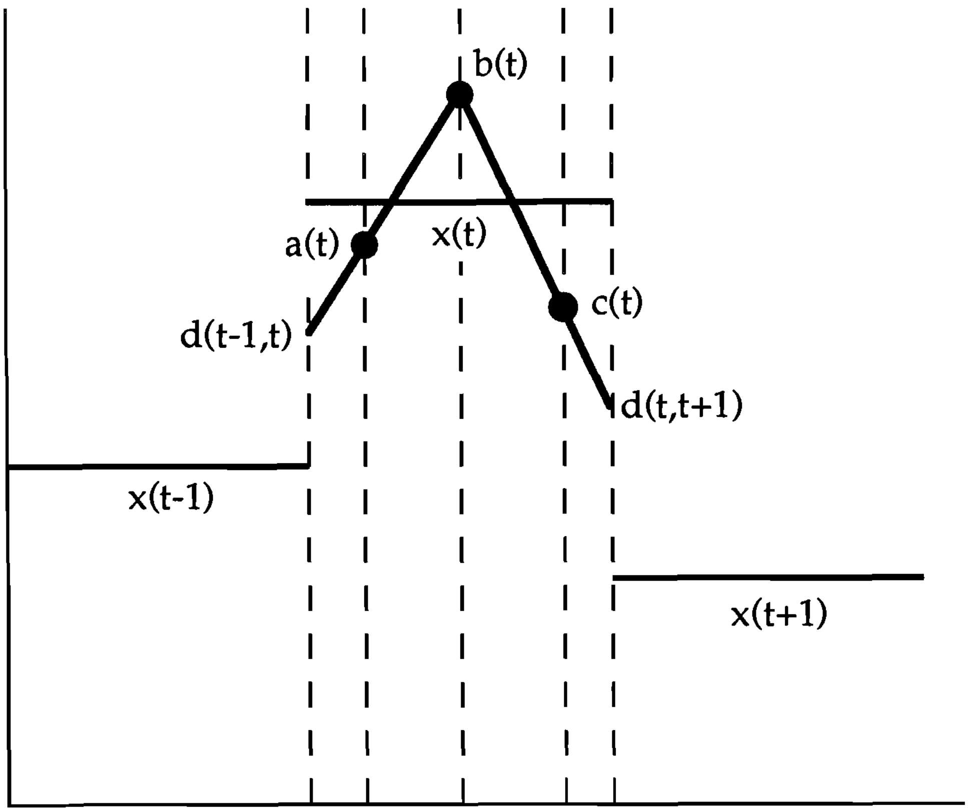
3) Each line segment d(t-1,t) b(t) and b(t) d(t,t+1) is a straight line segment.

Noting that the line segments d(t-1,t) a(t) and a(t) b(t) are in a ratio of 1:2, as are the line segments c(t) d(t,t+1) and b(t) c(t), equations (A1) and (A2) imply that:

$$\begin{aligned} a(t) &= {}^{1}/_{15}\,(12x(t) + 4x(t-1) - x(t+1)) \\ b(t) &= {}^{7}/_{5}\,x(t) - {}^{1}/_{5}\,x(t+1) - {}^{1}/_{5}\,x(t-1) \\ c(t) &= {}^{1}/_{15}\,(12x(t) - x(t-1) + 4x(t+1)) \end{aligned}$$

This is represented diagrammatically by Figure 1.

Figure 1: Interpolation of Quarterly-Period-Average-Data



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