WHY DOES THE AUSTRALIAN DOLLAR MOVE SO CLOSELY WITH THE TERMS OF TRADE?

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Abstract

The paper is motivated by two empirical results. Australia's terms of trade exhibit temporary fluctuations around a slowly declining trend, and movements in Australia's real exchange rate tend to follow those in the terms of trade. Together these results imply predictability in Australia's real exchange rate as well as the presence of predictable excess returns that are sometimes quite large.

Using a simple econometric model, with the terms of trade as the sole explanator, the paper demonstrates the forecastability of Australia's real exchange rate over horizons ranging from one to two years. It then quantifies the magnitude of the predictable excess returns to holding Australian dollar denominated assets over such horizons, finding them to be highly variable and sometimes quite large in magnitude. The results suggest a relative scarcity of forward-looking foreign exchange market participants with an investment horizon of a year or more.

JEL Classification Numbers C15, C22, F31.

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1. Introduction

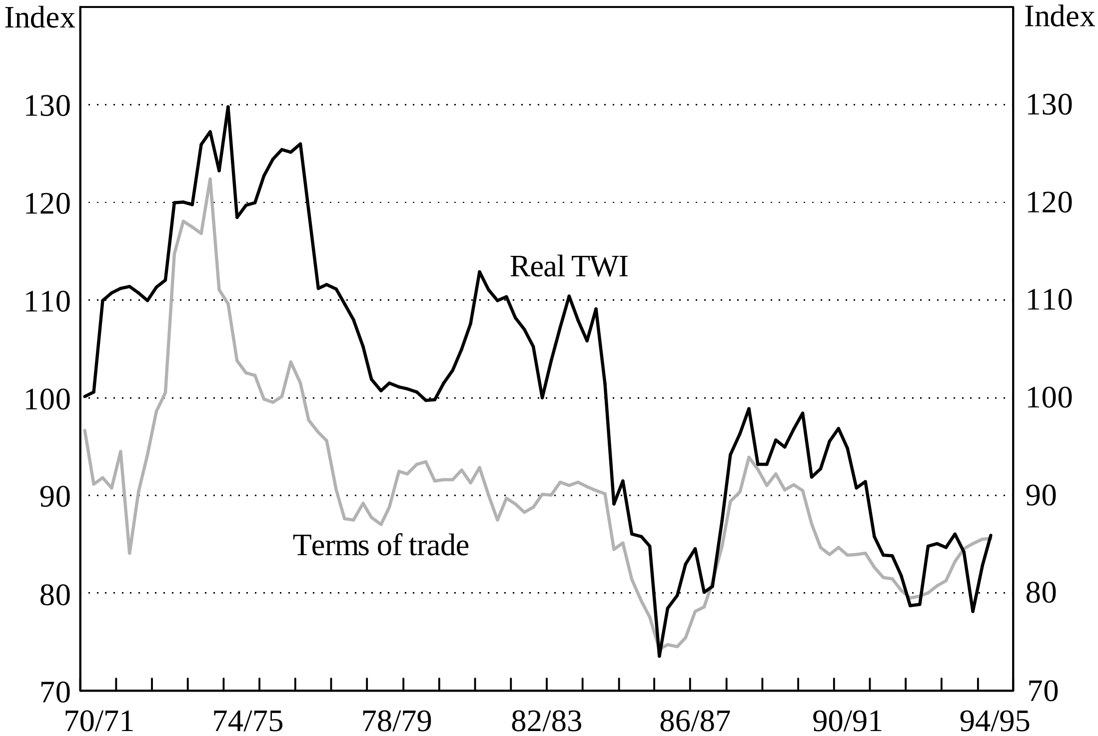
It has long been observed that there is a close relationship between commodity prices and the Australian dollar. When the world price of commodities rises, the Australian dollar tends to appreciate. When world commodity prices fall, the Australian dollar tends to fall. As a consequence, the Australian dollar is sometimes described as a 'commodity currency'.

Since commodities account for a large proportion of Australia's exports, this close link between commodity prices and the Australian dollar is also reflected in a close relationship between Australia's terms of trade and the real exchange rate, as shown in Figure 1 below. As the figure shows, the relationship has, if anything, become closer since the float of the Australian dollar in December 1983.

This relationship has been widely interpreted as evidence of a close link between the exchange rate and fundamentals. McKenzie (1986), Blundell-Wignall and Gregory (1990), Blundell-Wignall, Fahrer and Heath (1993) and Gruen and Wilkinson (1994) all argued that the terms of trade are a fundamental determinant of the real exchange rate for a small commodity-exporting country like Australia. However, while the direction of this link is consistent with economic theory, it is the contention of this paper that the extent to which the Australian real exchange rate responds to movements in the terms of trade provides evidence of short-sighted behaviour in the foreign exchange market. Australia's real exchange rate appears to *move too closely* with the terms of trade to be consistent with the actions of rational market participants with long investment horizons.

Our argument is easily summarised. As we show, there is substantial predictability of future movements in the terms of trade. This predictability, together with the fairly close co-movement of the Australian real exchange rate and the terms of trade often implies a large expected real exchange rate change. Furthermore, in general, this expected exchange rate change is not offset by an expected Australian/foreign real interest differential, implying the existence of substantial predictable excess returns to holding Australian assets. This observation, however, suggests a relative scarcity of rational investors with long-horizons, because if such investors were instead relatively numerous, they should move the *current* exchange rate to eliminate most of any predictable excess returns to holding Australian assets.

Figure 1: Terms of Trade and Real Exchange Rate (Real TWI)
June 1970 = 100



Note: The figure shows the terms of trade for goods and services and the 22 country trade-weighted Australian real exchange rate derived using CPIs to deflate nominal exchange rates. Underlying CPI is used where available.

We develop this argument in four steps. First, we show that Australia's terms of trade are well-described as fluctuating around a slowly declining trend. Deviations from this slow downward trend are quite long-lived, but do not appear to be permanent. Instead, there is strong evidence that the terms of trade revert to their historical trend over time. Thus, at times when the terms of trade are below trend, they can be expected to improve relative to trend, and when they are above trend, they can be expected to fall.

The second step is to estimate simple time-series models describing the terms of trade as well as the relationship between the terms of trade and the real exchange rate since the float of the Australian dollar. These models are used to generate forecasts of the change in the real exchange rate over horizons ranging from one to eight quarters ahead. The generated forecasts are truly *ex ante*; that is, they are based solely on the past behaviour of the real exchange rate and the terms of trade.

The third step is to compare these forecasts with the actual change in the real exchange rate over the forecast period. As we show, the forecasts provide no statistically-significant information about the change in the real exchange rate over short horizons – one, two or three quarters. By contrast, however, the forecasts provide significant, and apparently unbiased, estimates of the change in the real exchange rate over horizons from four to eight quarters ahead.

The final step is to combine the forecasts of the change in the real exchange rate with estimates of the Australian-foreign expected real interest differential. With these two ingredients, we can estimate the predictable excess return to holding Australian dollar-denominated assets.

The key point of the paper is that the link between the terms of trade and the real exchange rate is strong enough to generate quite large predictable excess returns to holding Australian assets over horizons of a year or more. At different times since the float, the predictable excess return to holding either one or two-year Australian bonds has varied in a range from about -15 to +15 per cent per annum. That is, at some times, an investor using our simple model would have expected a return of about 15 per cent per annum *less* holding either one or two-year Australian bonds than holding a portfolio of foreign bonds with the same maturity. At other times, the predictable excess return to holding Australian one or two-year bonds was about +15 per cent per annum.

The presence of such large predictable excess returns, sometimes positive, sometimes negative, suggests less than efficient processing of relevant information by the foreign exchange market. There appears to be a relative scarcity of forward-looking participants in the market with an investment horizon of a year or more. It may be that central banks are among the few active portfolio managers in the market with an investment horizon this long.

The next section of the paper, Section 2, examines the dynamic properties of Australia's terms of trade. It reports the results of three statistical tests which suggest that Australia’s terms of trade exhibit only temporary fluctuations around a slowly declining trend. A preferred time-series model of the terms of trade is also presented.

Section 3 describes two simple models of Australia’s trade-weighted real exchange rate over the post-float period. For both models, the terms of trade is the sole explanator of the real exchange rate. These models are used to generate *ex ante* forecasts of the change in Australia’s real exchange rate. Section 4 explains the forecasting procedure and examines the models’ out-of-sample forecasting performance over different time horizons. Estimates of the one and two-year expected excess returns to holding Australian dollar assets are also derived. Section 5 discusses the results and concludes.

2. The Dynamic Properties of the Terms of Trade

2.1 Statistical Tests

This section examines the dynamic properties of Australia’s terms of trade. At issue is whether shocks have a permanent effect on the level of the terms of trade (in which case the terms of trade process contains a unit-root), or only a temporary effect (implying that the terms of trade are stationary, possibly around a trend).

In previous empirical research on the Australian real exchange rate, the terms of trade was characterised as a unit-root process (Blundell-Wignall and Gregory (1990), Blundell-Wignall, Fahrer and Heath (1993) and Gruen and Wilkinson (1994)). This characterisation was primarily based, however, on statistical tests which did not allow for a trend in the terms of trade. Visual inspection of Figure 1 above suggests a slight but noticeable downward trend in the terms of trade, suggesting that such a trend should be allowed for when examining the time-series properties of the terms of trade. When this is done, strong evidence emerges that the terms of trade exhibit only temporary deviations from a slowly declining trend.

We present evidence from three statistical tests: the Augmented Dickey-Fuller (ADF) test, the Dickey-Fuller (1981) Φ_3 test and the Kwiatkowski, Phillips, Schmidt and Shin (1992) test. The ADF unit-root test is based on the regression:

$$\Delta tot_t = a + bt + (r-1)tot_{t-1} + \sum_{j=1}^{p} g_j \Delta tot_{t-j} + e_t \qquad (1)$$

where tot_t is natural log of the terms of trade in period t, Δ is the first difference operator, α is the constant or 'drift' term, and t is a linear time trend. The lag length on the lagged dependent variable, p, is chosen to eliminate serial correlation in the estimated residuals of equation (1). We set $p = 4$, and in our results in Table 1 below, show evidence of a lack of serial correlation in the residuals (based on the Ljung-Box Q statistic).[1]

For the entire sample period, Sept 1969 – June 1994, Table 1 shows that the null hypothesis that the terms of trade possesses a unit root is rejected at the 1 per cent level of significance on the basis of both the ADF statistic with constant and trend and the Dickey-Fuller Φ_3 statistic. Note, however, that the ADF statistic with constant term and no trend leads to rejection of the null hypothesis only at the 10 per cent level of significance. Thus, assuming no trend gives much weaker evidence against non-stationarity in Australia's terms of trade over the period.

Table 1: Augmented Dickey-Fuller and Dickey-Fuller Tests for Australia's Terms of Trade

	ADF statistic (with constant and trend)	ADF statistic (with constant and no trend)	D&F Φ_3 test statistic	Ljung-Box Q statistic
Sept 1969 – June 1994	-5.04***	-2.87*	13.76***	19.44
Sept 1969 – Dec 1986	-4.08**	-1.61	9.99***	16.14

Notes: ***, **, and *, signify rejection of the relevant null hypotheses at 1%, 5%, and 10% levels of significance. The Ljung-Box Q Statistic is for the first 23 autocorrelations of the residuals for the full sample, and 16 for the truncated sample.

1 See Appendix A for further details on all the statistics reported here.

Later in the paper, we generate forecasts of the real exchange rate beginning in 1987:Q1. Table 1 also reports the results of the statistical tests over a sample period ending before these forecasts begin. Estimation over this truncated sample period, Sept 1969 – Dec 1986, yields results that are qualitatively similar to those generated using the entire sample. The ADF statistic with constant and trend rejects the presence of a unit root at the 5 per cent significance level, while the Dickey and Fuller Φ_3 statistic rejects the null hypothesis of non-stationarity at the 1 per cent level of significance.

In contrast to the ADF unit-root test, the Kwiatkowski, Phillips, Schmidt and Shin (KPSS) test has either trend or level-stationarity as a null hypothesis. It thus complements the ADF unit-root test. The KPSS test for trend-stationarity involves regressing the variable under examination against a constant and time trend and calculating the following modified LM statistic:

$$\hat{h} = \frac{\sum_{t=1}^{T} S_t^2}{T^2 s^2(k)} \tag{2}$$

where $S_t = \sum_{i=1}^{t} e_i$ is the partial sum process of the regression residuals, e_i, $s^2(k)$ is a consistent estimator of the long run error variance based on a Bartlett window adjustment using the first k sample autocovariances as advocated by Newey and West (1987), and T is the sample size.

Table 2 presents KPSS test statistics, calculated over the two sample periods, Sept 1969 – Dec 1986, and Sept 1969 – June 1994, for different values of k. The null hypothesis of trend-stationarity in Australia's terms of trade is accepted at a 5 per cent level in all cases.

Thus, over either sample, the three statistical tests provide evidence that Australia's terms of trade exhibit only temporary fluctuations around a slowly declining trend.

Table 2: The KPSS Test for Australia's Terms of Trade

Lag truncation parameter k	KPSS test statistic Sept 1969 – Dec 1986	KPSS test statistic Sept 1969 – June 1994
2	0.139	0.114
4	0.095	0.077
6	0.081	0.063
8	0.077	0.059
10	0.078	0.059
12	0.082	0.061

Note: The null hypothesis of stationarity around a constant and trend is rejected for values of the KPSS statistic greater than 0.216, 0.146, and 0.119 at the 1%, 5% and 10% significance levels.

2.2 Modelling Australia's Terms of Trade

For the forecasting exercise to follow, we need a time-series model of the Australian terms of trade. For simplicity, we restrict our specification search to autoregressive models, AR(1) to AR(8). Given the evidence from the last section, each model includes a constant and linear trend. Thus, we estimate the following models:

$$tot_t = a + dt + \sum_{i=1}^{p} b_i tot_{t-i} + u_t \qquad \text{for } p=1,2, \ldots ,8 \qquad (3)$$

where tot_t is the log terms of trade in period t, α is a constant, t is a linear time trend, and u_t is a mean-zero error term. Estimation is over the truncated sample period, 1971:Q3 – 1986:Q4, so that the preferred specification is determined using data preceding the forecast period (which begins in 1987:Q1).

Summary statistics for the estimated models are presented in Table 3. On the basis of both the Akaike and Schwarz information criteria, the AR(5) model is the preferred specification for Australia's log terms of trade. We therefore use this specification to generate out-of-sample forecasts of the terms of trade throughout the forecast period.[2]

[2] While we keep the same AR(5) specification, we do re-estimate the model parameters as each new quarter of data becomes available in the forecast period.

Table 3: Autoregressive Time Series Models for the Terms of Trade

Model	Akaike information criterion (AIC)	Schwarz information criterion (SIC)	$\overline{R}^2$	Ljung-Box Q statistic
AR(1)	-161.90	-155.52	0.90	20.01
AR(2)	-160.45	-151.94	0.90	18.15
AR(3)	-162.02	-151.38	0.90	21.98**
AR(4)	-163.85	-151.09	0.90	19.41*
AR(5)	-175.21	-160.32	0.92	15.94
AR(6)	-174.51	-157.49	0.92	13.47
AR(7)	-172.53	-153.39	0.92	12.93
AR(8)	-172.78	-151.51	0.92	11.81

Note: The models are estimated over the period Sept 1971 – Dec 1986. **, and *, signify rejection of the null hypothesis at the 5% and 10% levels of significance. The Ljung-Box Q Statistic is for the first 15 autocorrelations of the residuals.

Table 4 presents coefficient estimates for this preferred AR(5) model:

$$tot_t = a + dt + \sum_{i=1}^{5} b_i tot_{t-i} + u_t \tag{4}$$

Note that the trend term is negative and highly significant, and that the sum of the coefficients on the lagged dependent variables, $\sum_{i=1}^{5} b_i = 0.72$, is less than unity, consistent with the earlier evidence that the terms of trade exhibit temporary fluctuations around a declining trend.

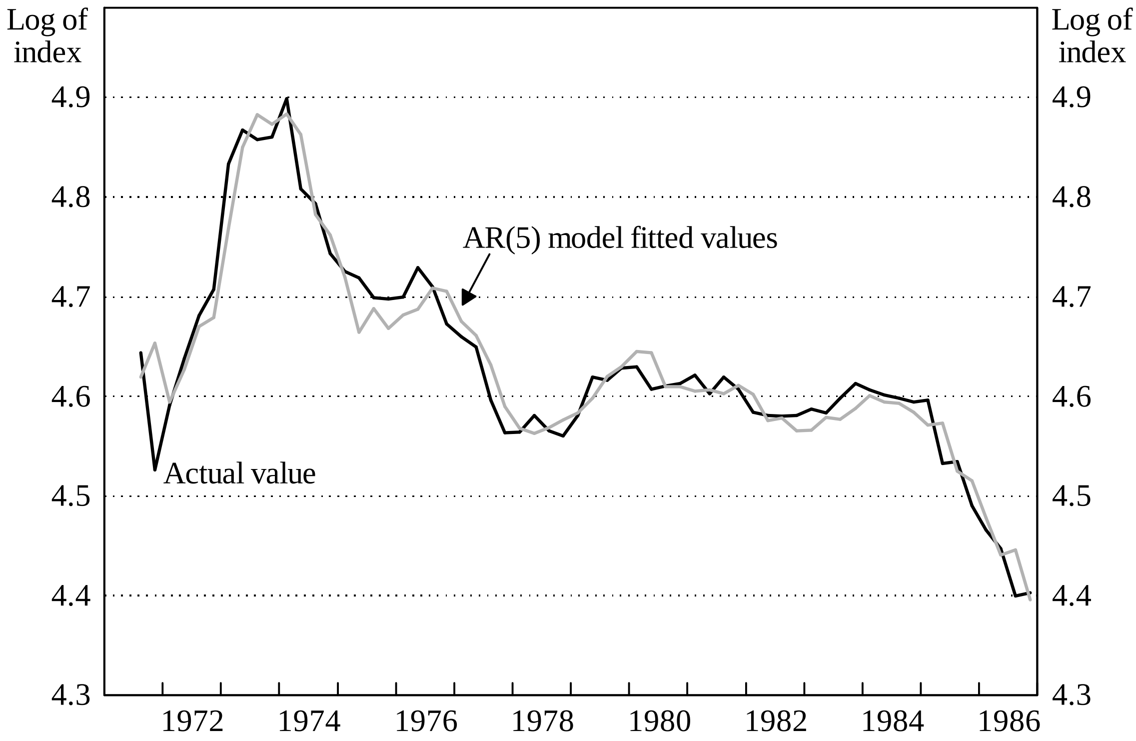
Table 4: Estimated AR(5) Model for Australia's Terms of Trade

	Constant	Trend	tot_{t-1}	tot_{t-2}	tot_{t-3}	tot_{t-4}	tot_{t-5}
coeff.	1.37	-0.0012	0.77	0.16	0.07	0.15	-0.43
(s.e.)	(0.29)	(0.0003)	(0.12)	(0.16)	(0.16)	(0.16)	(0.12)

Note: Estimation Period: Sept 1971 – Dec 1986.

Finally, Figure 2 shows the actual log terms of trade over the sample period, 1971:Q3-1986:Q4, and the estimates generated from the preferred AR(5) model. The in-sample fit of the preferred model is clearly quite good.

Figure 2: Actual and Fitted Log Terms of Trade



3. Modelling the Real Exchange Rate

We construct a measure of Australia's real exchange rate as a trade-weighted arithmetic average of the real exchange rates of Australia's five largest trading partners, using trade weights derived from average annual trade flows over the two financial years 1984/85 and 1985/86.[3] With out-of-sample forecasting beginning in 1987:Q1, this choice of trade weights again ensures that model-generated forecasts are truly *ex ante*. The measure of the real exchange rate, q_t, is therefore:

$$q_t \equiv \ln \sum_{j=1}^{5} w_j Q_{jt} \qquad \textit{where} \quad Q_{jt} \equiv \frac{E_{jt} \cdot P_t^{AUS}}{P_t^j} \tag{5}$$

w_j is the normalised country j trade weight, Q_{jt} is the Australia-country j real exchange rate, P_t^{AUS} and P_t^j are consumer price indices in Australia and country j,

3 Australia's five largest trading partners over the period, with normalised trade weights, were: Japan 0.4485, US 0.2896, UK 0.0969, W. Germany 0.0834, NZ 0.0816.

and E_{jt} is the price of the Australian dollar in country *j*'s currency on the last day of quarter *t*.

Two time-series models for this trade-weighted real exchange rate, q_t, are estimated.[4] The first, Model A, simply assumes that real exchange rate changes are determined by contemporaneous changes in Australia's terms of trade:

$$\Delta q_t \quad = \quad l.\Delta tot_t \quad + \quad e_t, \qquad\qquad l > 0 \tag{6}$$

To derive the second time-series model for Australia's real exchange rate, we begin with an unrestricted error-correction model (ECM):

$$\Delta q_t = a_0 + a_1 t + \sum_{i=0}^{4} g_i \Delta tot_{t-i} + \sum_{i=1}^{4} l_i \Delta q_{t-i} + c_1 tot_{t-1} + c_2 q_{t-1} + e_t \tag{7}$$

This specification allows for a longer-run relationship between the log-levels of the terms of trade and the real exchange rate. It also includes a time-trend to allow for the possibility that the real exchange rate and the terms of trade do not share the same longer-term trend. Since the relationship between the real exchange rate and the terms of trade may change over time, we generate preferred specifications over five sample periods, each with a starting date of 1984:Q1, but with end-dates extending in annual increments from 1989:Q4 to 1993:Q4.[5]

For each sample period, we use a general-to-specific modelling approach. We test sequentially larger sets of exclusion restrictions on the regressors of the unrestricted ECM, leading eventually to identification of the statistically-significant regressors to be included in the estimated equation for the real exchange rate (see Appendix B for further details).

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[4] We estimate the real exchange rate models over the post-float period; a sample so short that tests of non-stationarity generate ambiguous results. Tests on a longer sample of Australia's trade-weighted real exchange rate suggest it is stationary, possibly around a trend (Gruen and Shuetrim 1994).

[5] A preferred specification is assumed to remain the same for the three quarters following each end-date. Given the lack of degrees of freedom, we impose the model A specification given by equation (6) for sample periods ending before 1989:Q4. In these estimated equations, the coefficient on Δtot_t is always significant at the 5 per cent level.

Three preferred regression specifications are identified. Thus, real exchange rate model B is given by:

$$\Delta q_t \quad = \quad g_0 \Delta tot_t \quad + \quad e_t$$

for estimation periods ending 1986:Q4 to 1990:Q3,

$$\Delta q_t = a_0 + a_1 t + \sum_{i=0}^{3} g_i \Delta tot_{t-i} + \sum_{i=1}^{3} l_i \Delta q_{t-i} + c_1 tot_{t-1} + c_2 q_{t-1} + e_t$$

for estimation periods ending 1990:Q4 to 1991:Q3,

$$\Delta q_t = a_0 + g_0 \Delta tot_t + c_1 tot_{t-1} + c_2 q_{t-1} + e_t$$

for estimation periods ending 1991:Q4 to 1994:Q2. (8)

4. Predictable Real Exchange Rate Changes and Excess Returns

4.1 Forecasting the Real Exchange Rate

We now turn to the out-of-sample forecasting performance of the two exchange rate models. Almost without exception, out-of-sample forecasting evaluations by previous researchers use actual future values of the explanatory variables to generate exchange rate forecasts.[6] By contrast, as we have stressed, the forecasts we generate are truly *ex ante*: they use only information available at the time the forecasts are made.

We use recursive estimation to generate out-of-sample forecasts. This involves initially estimating the models over a sample period up to, but not including, the first quarter in which forecasts are made, 1987:Q1.[7] The estimated terms of trade model

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[6] The seminal references are Meese and Rogoff (1983a, 1983b). See Frankel and Rose (1994) and Taylor (1995) for recent reviews on empirical exchange rate research.

[7] Expectations formed at the end of quarter *t*, E_t, use all information available at the end of this quarter. While nominal exchange rates and interest rates are available, the period-*t* realisations of the terms of trade and domestic and foreign price levels are not. The period-*t* expectations $E_t tot_t$ and $E_t q_t$ are therefore generated using terms of trade and real exchange rate models up to and including quarter *t-1*.

is used to generate out-of-sample forecasts for the terms of trade from one to eight quarters ahead. These forecasts are then used with the preferred exchange rate models to generate forecasts of the real exchange rate change from one to eight quarters ahead.

The sample is then extended by one quarter to include data from 1987:Q1, and the models' parameters re-estimated. This leads to new out-of-sample forecasts for the terms of trade and the change in the real exchange rate from one to eight quarters ahead. Repeating this procedure for each subsequent quarter, up to and including 1994:Q2, generates 30 out-of-sample forecasts of the expected change in the trade-weighted real exchange rate for each forecast horizon from one to eight quarters ahead.

We can now evaluate the models' out-of-sample forecasting performance by regressing the actual, *ex post* real exchange rate change on its *ex ante* forecast:

$$q_{t+k} - q_t = a + b\, E_t^x (q_{t+k} - q_t) + e_{t+k,t}, \quad x = A, B \tag{9}$$

where E_t^x is the expectation at time-t based on real exchange rate model $x = A,B$. If the model contains no useful information for k-period-ahead forecasts of the exchange rate change, the coefficient estimate of b will be insignificantly different from zero. In this case, the exchange rate model forecast does not out-perform those of a random walk. If, however, b is significantly different from zero, then the model contains significant information about future movements of the real exchange rate. Further, if this coefficient is insignificantly different from unity, this suggests the model generates unbiased forecasts of the real exchange rate change k-quarters ahead.[8] In the results reported below, we test two null hypotheses, H_{01}: $b = 0$ and H_{02}: $b = 1$.

Before estimating equation (9) and conducting these hypothesis tests, however, there are two technical difficulties to be addressed. The first is that the OLS estimate of b may be biased in small samples, if the error terms in the terms of trade equation (equation 4) and in the process driving the real exchange rate are

8 A formal test of unbiasness also requires $a = 0$.

correlated (Stambaugh 1986). This problem is potentially serious as the sample size is indeed small, especially for longer-horizon forecasts.[9]

The second difficulty is that, for $k > 1$, the forecast horizon extends beyond the sampling interval which induces (k-1)th-order serial correlation in the regression residuals, $e_{t+k,t}$. This problem can be dealt with by using the Newey and West (1987) consistent estimate of the asympotic covariance matrix, and we report results based on this approach. Unfortunately, while it is valid asympotically, there is no guarantee that this approach gives accurate results in small samples.

To deal with both these small-sample problems, we therefore conduct Monte Carlo simulations, described in Appendix C, to derive estimates of the exact distributions of relevant statistics.

Tables 5a and 5b report the results for models A and B. They show OLS estimates of the coefficients in equation (9), R^2s, and adjusted values, $\hat{b}_{ADJ}$ and R^2_{ADJ}, derived by subtracting from the OLS estimates, $\hat{b}$ and R^2, median estimates from Monte Carlo simulations assuming the real exchange rate is unforecastable. For the hypothesis test H_{01}: $b = 0$, the tables show Newey-West t-statistics and their associated marginal significance levels against the alternative $b > 0$, denoted MSL_{NW}, as well as results derived from the Monte Carlo simulations. For H_{02}: $b = 1$, the Tables show only marginal significance levels from the Monte Carlo simulations.

We begin with results for model A. Table 5a shows that both the OLS coefficient estimate, $\hat{b}$, and the explanatory power of the regression, measured by R^2, rises as the forecast horizon is lengthened – a pattern suggesting that the explanatory

9 One to eight-quarter-ahead forecasts are made in each period, 1987:Q1 to 1994:Q2 while the actual exchange rate data ends in 1994:Q4. Hence, the sample size for the equation (9) regressions is 30 for k = 1,2, but progressively less for forecasts further ahead (for k = 8, it is only 24).

Table 5a: Model A Out-Of-Sample Forecasting Performance

Regressions of actual on expected change in the log real exchange rate

$$q_{t+k} - q_t = a + b\, E_t^A(q_{t+k} - q_t) + e_{t+k,t}$$

						$H_{01}: b = 0$				$H_{02}: b = 1$
k	$\hat{a}$	$\hat{b}$	$\hat{b}_{ADJ}$	R^2	R^2_{ADJ}	$t_{NW}(\hat{b})$	MSL_{NW}	$t_{OLS}(\hat{b})$	MSL_{MC}	MSL_{MC}
(i)	*(ii)*	*(iii)*	*(iv)*	*(v)*	*(vi)*	*(vii)*	*(viii)*	*(ix)*	*(x)*	*(xi)*
1	-0.0026	0.30	0.39	0.01	0.00	0.61	0.272	0.61	0.239	0.378
2	0.0009	0.51	0.60	0.07	0.04	1.19	0.116	1.45	0.134	0.423
3	-0.0006	0.51	0.58	0.11	0.06	1.52	0.064	1.83	0.132	0.437
4	0.0095	0.59	0.64	0.19	0.12	2.34	0.010	2.45	0.110	0.448
5	0.0190	0.64	0.65	0.22	0.14	2.84	0.002	2.63	0.122	0.457
6	0.0140	0.62	0.60	0.22	0.12	2.68	0.004	2.61	0.153	0.468
7	0.0110	0.62	0.59	0.22	0.10	2.63	0.004	2.56	0.181	0.484
8	0.0110	0.62	0.55	0.20	0.07	2.64	0.004	2.37	0.214	0.499

Note: Column *(i)* reports the forecast horizon, k, while columns *(ii)*, *(iii)* and *(v)* report the OLS coefficient estimates, $\hat{a}$, $\hat{b}$ and the regression R^2. Columns *(iv)* and *(vi)* report adjusted values $\hat{b}_{ADJ}$ and R^2_{ADJ}, derived by subtracting from the OLS estimates, median estimates derived from Monte Carlo simulations assuming the real exchange rate is unforecastable (that is, assuming data generating process, DGP_1 as defined in Appendix C). Columns *(vii)* and *(viii)* report the asymptotically-valid Newey-West *t*-statistics for $\hat{b}$ and the Newey-West marginal significance levels based on these *t*-statistics, MSL_{NW} (the probability of accepting $H_{01}: b = 0$, rather than the alternative, $b > 0$). Columns *(ix)* and *(x)* report the OLS *t*-statistics for $\hat{b}$ and the Monte Carlo derived marginal significance levels based on these *t*-statistics, MSL_{MC} (again, the probability of accepting $H_{01}: b = 0$, rather than the alternative, $b > 0$). Assuming that the estimated relationship between the terms of trade and the real exchange rate over the full post-float sample is the true data generating process (that is, assuming DGP_2), column *(xi)* reports the proportion of Monte Carlo trials for which $\hat{b} > 1$. See Appendix C for details of the Monte Carlo simulations.

power of model A rises with forecast length. Unfortunately, however, this conclusion is premature. This pattern of rising estimates for longer forecast horizons is not shared by $\hat{b}_{ADJ}$ and R^2_{ADJ}, implying that there is no clear improvement in the explanatory power of model A as the horizon is lengthened.

Table 5b: Model B Out-Of-Sample Forecasting Performance

Regressions of actual on expected change in the log real exchange rate

$$q_{t+k} - q_t = a + b\, E_t^{\mathrm{B}}(q_{t+k} - q_t) + e_{t+k,t}$$

						$H_{01}: b=0$				$H_{02}: b=1$
k	$\hat{a}$	$\hat{b}$	$\hat{b}_{ADJ}$	R^2	R^2_{ADJ}	$t_{NW}(\hat{b})$	MSL_{NW}	$t_{OLS}(\hat{b})$	MSL_{MC}	MSL_{MC}
(i)	*(ii)*	*(iii)*	*(iv)*	*(v)*	*(vi)*	*(vii)*	*(viii)*	*(ix)*	*(x)*	*(xi)*
1	-0.0048	0.05	0.11	0.00	-0.02	0.13	0.447	0.13	0.389	0.137
2	-0.0011	0.41	0.48	0.06	0.03	1.26	0.104	1.34	0.129	0.191
3	0.0004	0.47	0.53	0.13	0.09	1.88	0.030	2.01	0.078	0.225
4	0.0140	0.59	0.64	0.24	0.19	3.06	0.001	2.87	0.046	0.256
5	0.0290	0.67	0.72	0.31	0.24	4.08	0.000	3.33	0.040	0.278
6	0.0340	0.69	0.73	0.35	0.27	4.25	0.000	3.60	0.042	0.308
7	0.0440	0.75	0.79	0.40	0.31	4.33	0.000	3.94	0.040	0.329
8	0.0580	0.80	0.83	0.41	0.32	4.73	0.000	3.94	0.050	0.341

Note: Column *(i)* reports the forecast horizon, *k*, while columns *(ii)*, *(iii)* and *(v)* report the OLS coefficient estimates, $\hat{a}$, $\hat{b}$ and the regression R^2. Columns *(iv)* and *(vi)* report adjusted values $\hat{b}_{ADJ}$ and R^2_{ADJ}, derived by subtracting from the OLS estimates, median estimates derived from Monte Carlo simulations assuming the real exchange rate is unforecastable (that is, assuming data generating process, DGP_1 as defined in Appendix C). Columns *(vii)* and *(viii)* report the asymptotically-valid Newey-West *t*-statistics for $\hat{b}$ and the Newey-West marginal significance levels based on these *t*-statistics, MSL_{NW} (the probability of accepting $H_{01}: b=0$, rather than the alternative, $b>0$). Columns *(ix)* and *(x)* report the OLS *t*-statistics for $\hat{b}$ and the Monte Carlo derived marginal significance levels based on these *t*-statistics, MSL_{MC} (again, the probability of accepting $H_{01}: b=0$, rather than the alternative, $b>0$). Assuming that the estimated relationship between the terms of trade and the real exchange rate over the full post-float sample is the true data generating process (that is, assuming DGP_2), column *(xi)* reports the proportion of Monte Carlo trials for which $\hat{b}>1$. See Appendix C for details of the Monte Carlo simulations.

Turning to the null hypothesis H_{01}: $b=0$, Table 5a shows that the marginal significance levels derived from the Newey-West *t*-statistics, MSL_{NW}, and from the Monte Carlo simulations, MSL_{MC}, are often very different. The samples are apparently small enough to render the Newey-West asymptotic results extremely inaccurate. As a consequence, we must rely on the Monte Carlo evidence to assess the performance of the model. Based on this Monte Carlo evidence, at all forecast horizons, the null hypothesis H_{01}: $b=0$ cannot be rejected against the alternative

$b > 0$, even at a 10 per cent level of significance. Model A cannot significantly outperform a random walk.

For model B, both the OLS coefficient estimate, $\hat{b}$, and the regression R^2, rise as the forecast horizon is lengthened. This pattern is repeated by $\hat{b}_{ADJ}$ and R^2_{ADJ}, implying that the explanatory power of model B does rise with forecast length; in contrast to the results for model A.

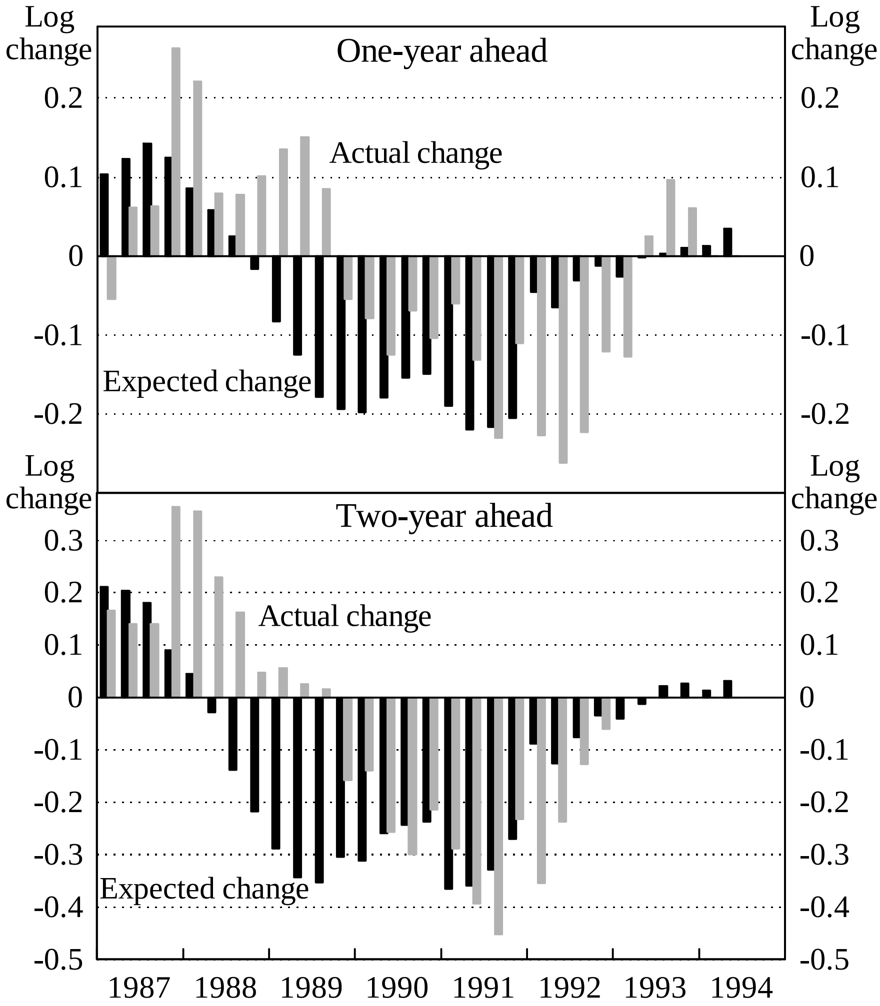
As for model A, there is a substantial difference between the marginal significance levels for the null hypothesis, $b = 0$, based on Newey-West *t*-statistics and those based on the Monte Carlo simulations, again leading us to rely on the Monte Carlo results to derive inferences about model performance. These Monte Carlo results reveal a marked improvement in out-of-sample forecasting performance as the forecast horizon lengthens. Point estimates of *b* for one, two, and three-quarters-ahead forecasts, are positive, but insignificant. Over these shorter forecast horizons, model B cannot outperform a random walk.

The contrast with forecasts over time horizons longer than three quarters is striking. For these longer horizons, the out-of-sample forecasting performance improves considerably and the coefficient estimate of *b* is positive and significant at a five per cent level. Thus, model B contains significant information about future movements of the real exchange rate for horizons ranging from one to two years.

The results in Table 5b also show that the null hypothesis, H_{02}: $b = 1$, cannot be rejected, implying that there is no evidence of bias in the model forecasts of the change in the real exchange rate.

Figure 3 shows the actual log changes in the real exchange rate against the expected changes from Model B, over one and two year horizons. The positive relationship between the expected and actual changes over each horizon emerges clearly from the figure.

Figure 3: Actual Versus Expected Log Real Exchange Rate Changes



4.2 Expected Excess Returns

We now turn to estimates of the one and two-year-ahead expected excess return to holding Australian bonds rather than a trade-weighted basket of foreign bonds.

Excess returns can be expressed either in terms of nominal appreciation and the nominal domestic/foreign interest differential, or real appreciation and the real interest differential. Using real variables, the excess return (in per cent) to holding a one-year (*4*-quarter) Australian dollar bond, $ER_{t,4}$, is approximately:

$$ER_{t,4} \approx 100 \times (q_{t+4} - q_t) + {}_4r_t^{AUS} - {}_4r_t^{*} \tag{10}$$

where ${}_4r_t^{AUS}$ and ${}_4r_t^*$ are time-t real interest rates on the domestic and trade-weighted foreign basket of one-year bonds, also in per cent. ${}_4r_t^*$ is defined by:

$$ {}_4r_t^* \equiv \sum_{j=1}^{5} w_j \, {}_4r_t^j \tag{11}$$

where w_j is the normalised country j trade weight, defined earlier, and ${}_4r_t^j$ is country j's real interest rate, ${}_4r_t^j \approx {}_4i_t^j - 100 \times (p_{t+4}^j - p_t^j)$, where ${}_4i_t^j$ is the one-year nominal interest rate and p_t^j, the log consumer price index in country j. The expected excess return on the Australian one-year bond, $E_t(ER_{t,4})$, is therefore:

$$E_t(ER_{t,4}) \approx 100 \times E_t^B(q_{t+4} - q_t) + E_t({}_4r_t^{AUS} - {}_4r_t^*) \tag{12}$$

where we assume model B is used to generate the expectations $E_t^B(q_{t+4} - q_t)$. For all countries, we also make the simple assumption that the expected inflation rate from period t to t+4 is equal to the most recently published annual inflation rate available in quarter t, that is the inflation rate from period $t-5$ to $t-1$, so that $E_t({}_4r_t^j) \approx {}_4i_t^j - 100 \times (p_{t-1}^j - p_{t-5}^j)$ for country j.

The two-year-ahead excess return, $ER_{t,8}$, in per cent is approximately:

$$ER_{t,8} \approx 100 \times (q_{t+8} - q_t) + 2 \times ({}_8r_t^{AUS} - {}_8r_t^*) \tag{13}$$

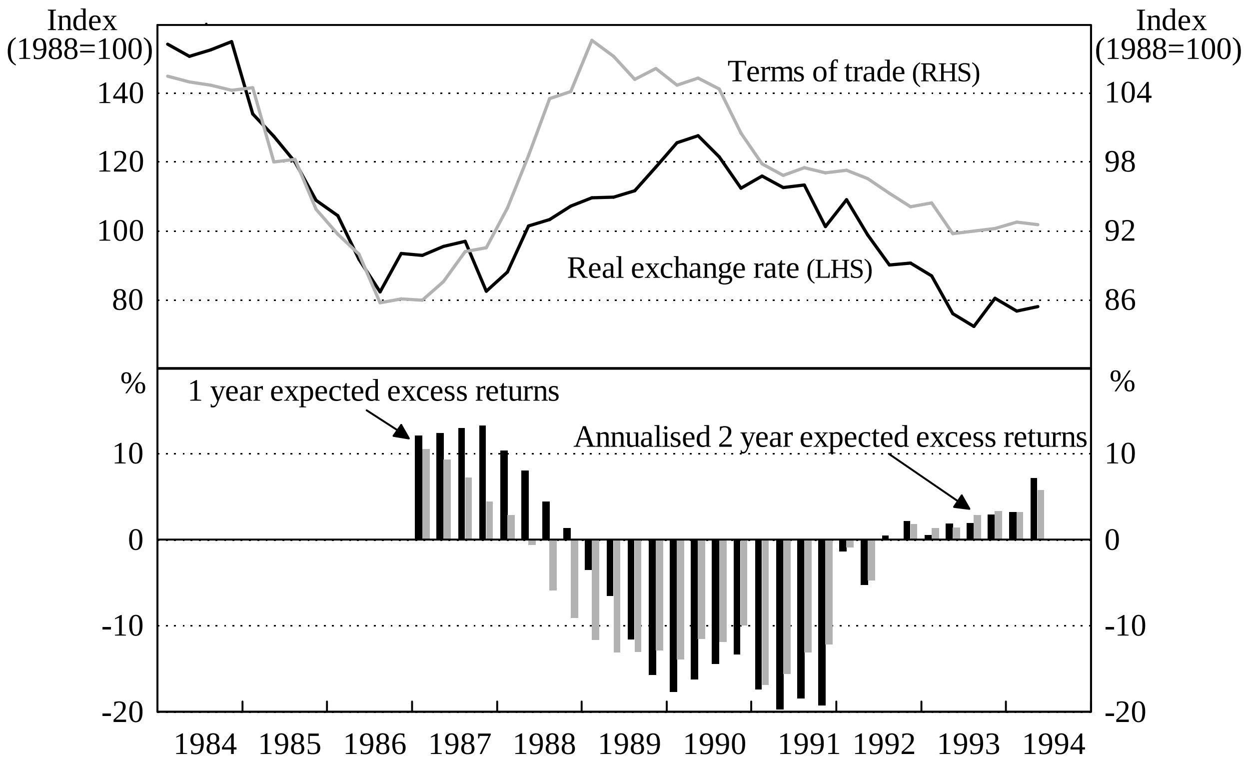
where ${}_8r_t^{AUS}$ and ${}_8r_t^*$ are the time-t real interest rates in per cent per annum on the domestic and trade-weighted foreign basket of two-year bonds, with the latter defined analogously to equation (11). For country j, ${}_8r_t^j$ is given by ${}_8r_t^j \approx {}_8i_t^j - (p_{t+8}^j - p_t^j) \times 100/2$. The expected excess return on the Australian two-year bond, $E_t(ER_{t,8})$, is therefore approximately:

$$E_t(ER_{t,8}) \approx 100 \times E_t^B(q_{t+8} - q_t) + 2E_t({}_8r_t^{AUS} - {}_8r_t^*) \tag{14}$$

where model B is used to generate the real exchange rate expectations, $E_t^B(q_{t+8} - q_t)$, and $E_t({}_8r_t^j) \approx {}_8i_t^j - 100 \times (p_{t-1}^j - p_{t-5}^j)$ for each country *j*.

Figure 4 shows the levels of the terms of trade and the five-country trade-weighted Australian real exchange rate since the float. Also shown, over the period 1987:Q1 to 1994:Q2, are the one-year expected excess returns, $E_t(ER_{t,4})$, and the annualised two-year expected excess returns, $AE_t(ER_{t,8})$, defined by $AE_t(ER_{t,8}) = E_t(ER_{t,8}) / 2$. Both these expected excess returns are highly variable over time and often large in magnitude. Thus, for example, the one-year expected excess return ranges from plus 13.2 per cent in 1987:Q4 to minus 19.7 per cent in 1991:Q2.

Figure 4: Terms of Trade, Real Exchange Rate and Expected Excess Returns



Both the size and variability of the expected excess returns is primarily due to large changes in Australia's expected real exchange rate, with little offset from the Australian-foreign expected real interest differential (see Figure 5 for a demonstration of this point for the one-year results). For both one and two-year expected returns, the correlation between the expected real exchange rate change and the expected real interest differential is negative but statistically insignificant.

Figure 5: One Year Expected Excess Returns

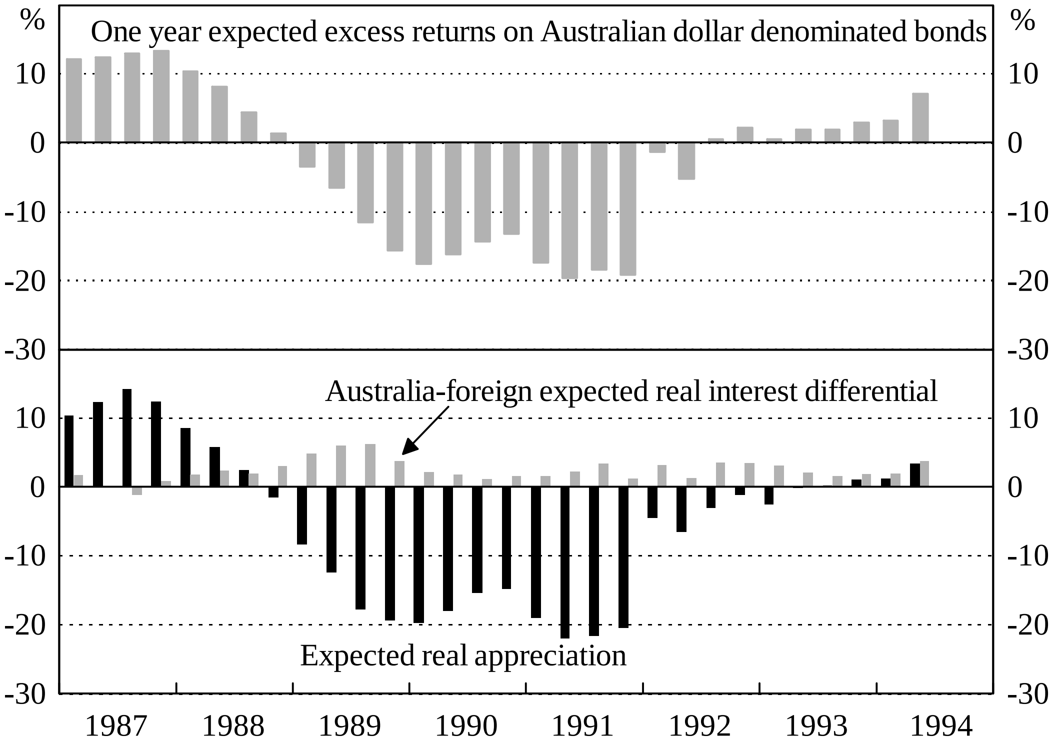


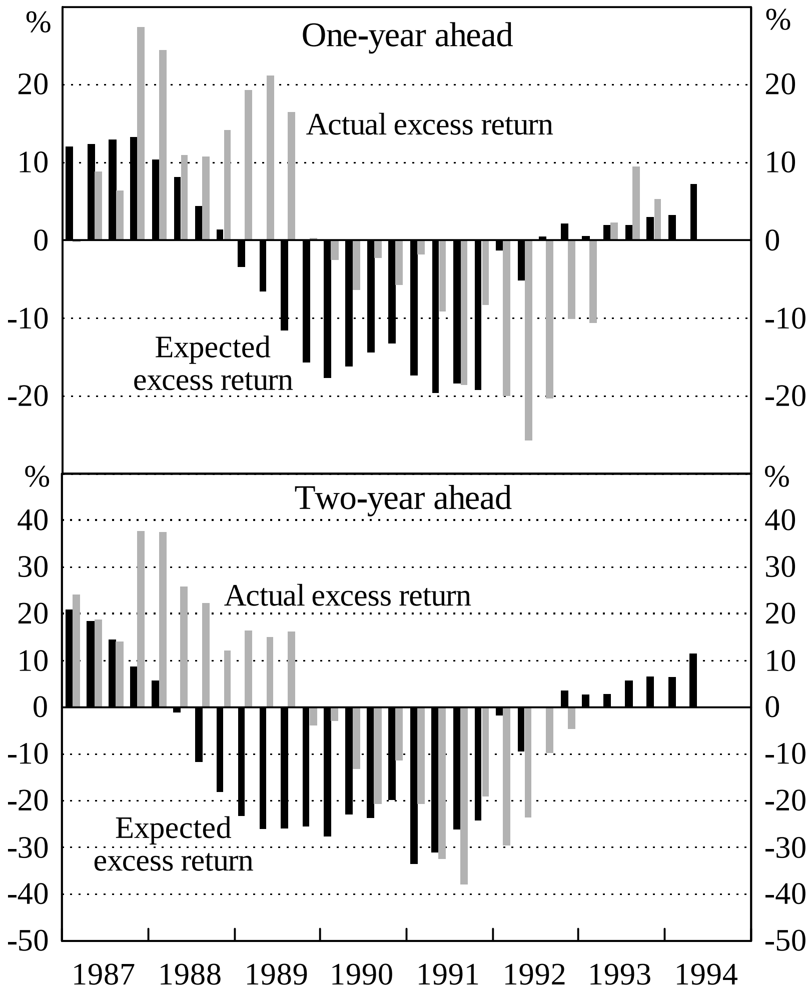
Figure 6 shows a comparison of actual and expected excess returns over both one and two year horizons. Given model B's capacity to predict real exchange rate changes over one and two-years, one might expect the model to have some predictive power for excess returns. As a formal test, we regress actual against expected excess returns to holding Australian dollar-denominated assets:

$$ER_{t,k} = a + b E_t(ER_{t,k}) + e_t, \quad k = 4,8, \tag{15}$$

and present the results, as well as relevant hypothesis tests, in Table 6. As before, we report results based on both Newey-West *t*-statistics and Monte Carlo simulations.

At both forecast horizons, the OLS coefficient estimate, $\hat{b}$, is positive and the equations have some explanatory power, judged by the R^2. The Newey-West *t*-statistics, $t_{NW}(\hat{b})$, suggest that the null hypothesis, $H_{01}: b = 0$, can be rejected against the alternative $b > 0$ at a one per cent significance level for both forecast horizons (that is, $MSL_{NW} < 0.01$). Again, however, the samples are so small that we must discount these results and rely instead on the Monte Carlo results. These results imply that the null hypothesis, $H_{01}: b = 0$, can be rejected against the alternative $b > 0$ at a significance level of only 16 or 17 per cent.

Figure 6: Actual and Expected Excess Returns



While the estimated expected excess returns on Australian-dollar assets are often quite large in magnitude (see Figure 4), the Monte Carlo results in Table 6 imply that we can only have a moderate degree of confidence that these expected excess returns help to predict actual excess returns. The small sample implies that the Monte-Carlo-estimated distributions of the coefficient estimate, $\hat{b}$, are broad enough to render statistical inference difficult. Only when a longer sample becomes available, will it be possible to be more definitive.

Table 6: Regression of Actual on Expected Excess Returns
$ER_{t,k} = a + b\, E_t(ER_{t,k}) + e_t, \quad k = 4,8$

				H_{01}: $b=0$			H_{02}: $b=1$
k	$\hat{a}$	$\hat{b}$	R^2	$t_{NW}(\hat{b})$	MSL_{NW}	MSL_{MC}	MSL_{MC}
(i)	*(ii)*	*(iii)*	*(iv)*	*(v)*	*(vi)*	*(vii)*	*(viii)*
4	3.08	0.54	0.18	2.86	0.00	0.16	0.25
8	8.65	0.71	0.28	3.42	0.00	0.17	0.31

Note: Column *(i)* reports the forecast horizon, *k*, while columns *(ii)*, *(iii)* and *(iv)* report the OLS coefficient estimates, $\hat{a}$, $\hat{b}$ and the regression R^2. Columns *(v)* and *(vi)* report asymptotically-valid Newey-West *t*-statistics for $\hat{b}$ and the Newey-West marginal significance levels based on these *t*-statistics, MSL_{NW} (the probability of accepting $H_{01}: b = 0$, rather than the alternative, $b > 0$). The Monte Carlo-based results reported in columns *(vii)* and *(viii)* assume data generating processes based on the estimated relationship between the terms of trade and the real exchange rate over the full post-float sample, and between the nominal interest differential and the quarterly inflation differential (that is, assuming DGP_3 and DGP_4 as defined in Appendix C). Column *(vii)* reports the proportion of Monte Carlo trials for which $\hat{b} < 0$, while column *(viii)* reports the proportion of trials for which $\hat{b} > 1$. See Appendix C for details of the Monte Carlo simulations.

5. Discussion and Conclusions

The paper has documented the predictability of the Australian real exchange rate over horizons of between one and two years. It has also presented results suggesting the presence of quite large and variable expected excess returns to holding Australian dollar assets. How can we explain these results?

One potential explanation is the presence of a time-varying risk premium. That is, the excess returns we document may not represent unexploited profit opportunities but, instead, the compensation demanded by risk-averse investors for bearing risk. Under such an interpretation, the positive (negative) excess returns displayed in Figure 4, would represent the risk-premium (discount) to holding Australian dollar assets relative to the trade-weighted portfolio of foreign bonds – ie, the additional compensation demanded by investors for holding the riskier Australian (trade-weighted foreign) bond. If this explanation is valid, then Australian bonds must be viewed by investors as much riskier than foreign bonds when Australia's terms of trade are low but expected to improve, and much less risky when the terms of trade are high but anticipated to fall. But the idea that the risk premium on

Australian bonds could be as large and volatile as suggested by the profile of expected excess returns shown in Figure 4, seems implausible to us.

A second possible explanation is that rational foreign exchange market participants only *gradually* learn of the predictability of exchange rate changes. It may take time for the market to learn of the existence of considerable longer-run predictability in Australia's terms of trade, and that this in turn renders medium and longer-run changes in Australia's real exchange rate forecastable. In which case, over time, one would expect to see the current level of the exchange rate increasingly reflect the market's implicit forecast of its level into the future. If so, the strong association between fluctuations in Australia's terms of trade and real exchange rate should decline over time.

A third possible explanation for our results is 'short-termism' of foreign exchange market participants. Recall that forecasting model B does not provide any statistically-significant information on exchange rate changes over short horizons, but does provide significant and apparently unbiased forecasts over horizons of one to two years. This indicates that in order to profitably exploit the real exchange rate – terms of trade link, it is necessary for investors to have a trading horizon of at least a year. However, it may be difficult for key market participants, such as fund-managers and institutional investors, to take long-term open positions in a foreign currency. For institutional reasons – related, for example, to the time-period over which their performance is assessed – their relevant trading horizon may necessarily be a few months, rather than the one or two years required to exploit the inherent predictability of the Australian currency. With short-term investment horizons apparently widespread in the foreign exchange market, it may not be so surprising to observe a relationship between the real exchange rate and the terms of trade that can only be exploited by (relatively scarce) long-horizon investors.

Finally, our paper has implications for empirical modelling of the Australian exchange rate. Most existing Australian empirical macroeconomic models, such as the Murphy model, the MSG2 model, and the Treasury's TRYM model, use the uncovered interest parity condition, sometimes allowing for a constant risk-premium, combined with the assumption of rational (or quasi-rational) expectations, as the central relationship determining exchange rate outcomes. Our results suggest that this is inconsistent with a key feature of the data. Predictable changes in Australia's terms of trade provide significant information about movements in the Australian real exchange rate over horizons of one to two years. Over these horizons, there appear to be quite large, and variable, expected excess returns on Australian dollar denominated assets – contrary to the predictions of uncovered interest parity.

Appendix A: Stationarity Tests

The Augmented Dickey-Fuller (ADF) statistic, with constant and trend, is the value of the *t*-statistic for the estimated coefficient (ρ-1) from the regression:

$$\Delta tot_t = a + bt + (r-1)tot_{t-1} + \sum_{j=1}^{p} g_j \Delta tot_{t-j} + e_t$$

Critical values for the null hypothesis, (ρ-1) = 0, in the presence of a constant and trend, at the 1 per cent, 5 per cent, and 10 per cent levels of significance are -4.15, -3.50 and -3.18 for sample size 50 and -4.04, -3.45, and -3.15 for sample size 100.

The ADF statistic, with constant and no trend, is the value of the *t*-statistic for the estimated coefficient (ρ-1) from the regression:

$$\Delta tot_t = a + (r-1)tot_{t-1} + \sum_{j=1}^{p} g_j \Delta tot_{t-j} + e_t$$

Critical values for the null hypothesis, (ρ-1) = 0, in the presence of a constant and no trend, at the 1 per cent, 5 per cent, and 10 per cent levels of significance, are -3.58, -2.93 and -2.60 for sample size 50 and -3.51, -2.89, and -2.58 for sample size 100.

The Dickey-Fuller Φ_3 test statistic, used to test the joint null hypothesis, β = 0 and $(r-1)=0$, is defined by $\Phi_3 = (RSS_r - RSS_u).T/2.RSS_u$, where RSS_r is the residual sum of squares from the restricted regression when β = 0 and $(r-1)=0$ are imposed jointly, RSS_u is the residual sum of squares from the unrestricted regression and T is the total number of observations in the sample. Critical values for the joint null hypothesis at the 1 per cent, 5 per cent, and 10 per cent levels of significance are 9.31, 6.73, 5.61 for sample size 50 and 8.73, 6.49, 5.47 for sample size 100.

The Ljung-Box Q statistic for M lags is given by $Q(M) = T(T+2)\sum_{j=1}^{M} \hat{r}_j^2/(T-j)$ where $\hat{r}_j$ is the sample autocorrelation of the residuals at lag *j*, *j = 1,2, ..., M*. The null hypothesis under the Ljung-Box Q test is that the first *M* autocorrelations of the residuals are zero, and under this null, the statistic is distributed $c^2_{(M)}$.

Kwiatkowski, Phillips, Schmidt and Shin (1992) decompose the time series $\{y_t\}$ into a linear trend *t*, a random walk r_t, and a stationary error term e_t. Thus, $y_t = xt + r_t + e_t$ where $r_t = r_{t-1} + u_t$ and $u_t \sim i.i.d(0, s^2_u)$. The assumption of trend-stationarity in the series $\{y_t\}$ implies the absence of the random walk component, or equivalently, that the variance of *u*, s^2_u is zero. This is the null hypothesis in the KPSS unit-root test.

Appendix B: Real Exchange Rate Model B

This appendix explains how the preferred specifications for real exchange rate model B are determined. For a given sample period, we begin with the unrestricted error-correction model defined by equation (7) in the text, reproduced below as model (B1). We then define eight restricted versions of this model derived by imposing sequentially larger sets of exclusion restrictions. Specifically, we have the following real exchange rate models:

$$\Delta q_t = a_0 + a_1 t + \sum_{i=0}^{5-j} g_i \Delta tot_{t-i} + \sum_{i=1}^{5-j} l_i \Delta q_{t-i} + c_1 tot_{t-1} + c_2 q_{t-1} + e_t, \qquad \text{(Bj), j = 1,…, 4}$$

$$\Delta q_t = a_0 + a_1 t + g_0 \Delta tot_t + c_1 tot_{t-1} + c_2 q_{t-1} + e_t \qquad \text{(B5)}$$

$$\Delta q_t = a_0 + g_0 \Delta tot_t + c_1 tot_{t-1} + c_2 q_{t-1} + e_t \qquad \text{(B6)}$$

$$\Delta q_t = a_0 + g_0 \Delta tot_t + e_t \qquad \text{(B7)}$$

$$\Delta q_t = g_0 \Delta tot_t + e_t \qquad \text{(B8)}$$

$$\Delta q_t = e_t \qquad \text{(B9)}$$

Note that model (B8) is simply exchange rate model A. For a given sample period, let ${}_iR_k$ be the set of exclusion restrictions that reduces model (Bi) to model (Bk), for i = 1, …, 9 and k = i +1, … , 9. Then, model (Bp) is the preferred specification, when at the 10 per cent level of significance, all of the exclusion restrictions ${}_lR_m$, l = 1, 2, … , p-1, m = l + 1, …, p are accepted and at least one of the exclusion restrictions ${}_lR_{p+1}$, l = 1, 2, … , p, is rejected. The exclusion restrictions are tested assuming the OLS variance-covariance matrix.

Appendix C: Monte Carlo Simulations

This appendix outlines the Monte Carlo simulations used to generate results reported in Tables 5a, 5b and 6 in the text. Simulations are conducted assuming four different data generating processes (DGPs), which we examine in turn.

C.1 The Real Exchange Rate

To begin, we test the null hypotheses that our models of the real exchange rate have no explanatory power, that is, that $b_k^x = 0$ in each of the equations:

$$q_{t+k} - q_t = a + b_k^x \, E_t^x(q_{t+k} - q_t) + e_{t+k,t}, \quad x = A, B, \quad k = 1, \ldots, 8 \qquad \text{(C1)}$$

To test these null hypotheses, we assume a data generating process (DGP$_1$) with the log terms of trade, tot_t, following the preferred specification identified in Section 2.2, and embodying the hypothesis that the real exchange rate is unforecastable:

$$tot_t = a + bt + \sum_{i=1}^{5} c_i \, tot_{t-i} + e_{1,t} \qquad \text{(C2)}$$

$$\Delta q_t = d + e_{2,t} \qquad \text{(C3)}$$

Let $e_t = (e_{1,t}, e_{2,t})'$, $V = E(e_t e_t')$ and the estimates be $(\hat{a}, \hat{b}, \hat{c}_1, \hat{c}_2, \hat{c}_3, \hat{c}_4, \hat{c}_5, \hat{d}, \hat{V})$. Both equations are estimated by OLS, with equation (C2) estimated over the full terms of trade sample, 1969:Q3 to 1994:Q2, and (C3) estimated over the post-float period, 1984:Q1 to 1994:Q2. The off-diagonal elements of V are derived from the correlation coefficient, $\hat{r}_{1,2}$, of the estimated errors, $\hat{e}_{1,t}$ and $\hat{e}_{2,t}$ over the post-float period 1984:Q1 to 1994:Q2. The estimates are:

$$\hat{a} = 1.06,\ \hat{b} = -0.00061,\ \hat{c}_1 = 0.90,\ \hat{c}_2 = 0.14,\ \hat{c}_3 = -0.06,\ \hat{c}_4 = 0.12,\ \hat{c}_5 = -0.33,\ \hat{s}_{e_1} = 0.02 \qquad \text{(C4)}$$

and

$$\hat{d} = -0.015,\ \hat{s}_{e_2} = 0.07 \text{ and } \hat{r}_{1,2} = 0.38 \qquad \text{(C5)}$$

The Monte Carlo distributions are generated by running 5,000 trials with each trial (i = 1, …, 5,000) proceeding as follows.

1. Draw a vector sequence of observations $\{e_j^i\}_{j=1}^{n+T}$ from a bivariate normal distribution with mean 0, covariance matrix $\hat{V}$, n = 58 and T = 44. (n is the length of the pre-float terms of trade sample; T is the post-float period, 1984:Q1 to 1994:Q4.)

2. Generate sequences of observations $\{tot_j^i\}_{j=6}^{n+T}$, $\{\Delta q_j^i\}_{j=6}^{n+T}$ according to

   $$tot_j^i = \hat{a} + \hat{b}\,j + \sum_{l=1}^{5} \hat{c}_l\, tot_{j-l}^i + e_{1,j}^i, \qquad \Delta q_j^i = \hat{d} + e_{2,j}^i,$$

   using $(tot_5, tot_4, tot_3, tot_2, tot_1)$ to start the autoregression.

3. For time t, 71 ≤ t ≤ 100 (t = 71 corresponds to 1987:Q1; t = 100 to 1994:Q2) use the sequence of synthetic terms of trade data, $\{tot_j^i\}_{j=1}^{t-1}$ to estimate the terms of trade model (C2) and thereby generate a vector of parameter estimates $(\hat{a}_t^i, \hat{b}_t^i, \hat{c}_{1,t}^i, \hat{c}_{2,t}^i, \hat{c}_{3,t}^i, \hat{c}_{4,t}^i, \hat{c}_{5,t}^i)$, and hence derive $E_t(tot_{t+k}^i)$, k = 0,…, 8, t = 71,…, 100.

4. For time t, 71 ≤ t ≤ 100, use the sequences of synthetic data, $\{tot_j^i\}_{j=n+1}^{t-1}$ and $\{\Delta q_j^i\}_{j=n+1}^{t-1}$ to estimate a preferred real exchange rate model. For model A,

this is equation (6) in the text, $\Delta q_j^i = l.\Delta tot_j^i + e_j$. For model B, it is one of the three specifications in equation (8), depending on the time, *t*. Use this preferred real exchange rate model, together with estimates of $E_t(tot_{t+k}^i)$, derived above, to generate $E_t^x(q_{t+k}^i - q_t^i)$, $x = A, B$, $k = 1,\ldots, 8$, $t = 71,\ldots, 100$.

5. Estimate the regressions $q_{t+k}^i - q_t^i = a + b_k^{x,i}\, E_t^x(q_{t+k}^i - q_t^i) + e_{t+k,t}^i$, $t = 71,\ldots,100$, by OLS for all values $k = 1,\ldots, 8$, and for the two models and generate the OLS *t*-statistics, $t_{OLS}(b_k^{x,i}) = \hat{b}_k^{x,i} / s.e(\hat{b}_k^{x,i})$, $x = A, B$.[10]

These 5,000 observations of $t_{OLS}(b_k^{A,i})$ and $t_{OLS}(b_k^{B,i})$ form the Monte Carlo distributions under the null hypothesis that the models of the real exchange rate have no explanatory power. Based on these Monte Carlo distributions, the results of the hypothesis tests, $H_{01}: b = 0$, for $k = 1,\ldots, 8$, are shown in column *(x)* in Table 5a for model A and in Table 5b for model B.

We turn now to the second data generating process (DGP_2) which assumes that the real exchange rate model, estimated over the post-float period 1984:Q1 to 1994:Q2, is the true model. For this data-generating process, we again assume that the terms of trade follow model (C2) with parameter values (C4). The real exchange rate model (C3) is, however, replaced by:

$$\Delta q_t = d + e\Delta tot_t + f\, tot_{t-1} + g\, q_{t-1} + e_{2,t} \qquad \text{(C6)}$$

with parameter values, estimated by OLS over the post-float period 1984:Q1 to 1994:Q2, given by:

$$\hat{d} = -1.95,\ \hat{e} = 1.45,\ \hat{f} = 0.42,\ \hat{g} = -0.14,\ \hat{s}_{e_2} = 0.06,\ \hat{r}_{1,2} = 0.003 \qquad \text{(C7)}$$

---

[10] As explained earlier, the actual exchange rate data ends in 1994:Q4, and hence the sample size is 30 for $k = 1,2$, but progressively less for forecasts further ahead (for $k = 8$, it is only 24). This pattern is replicated for the synthetic data.

The Monte Carlo distributions under this data generating process are again generated by running 5,000 trials with each trial (i = 1, …, 5,000) proceeding as before, but with the following modifications. In step 2, the sequences of observations $\{tot_j^i\}_{j=6}^{n+T}$, $\{\Delta q_j^i\}_{j=6}^{n+T}$ are now derived from $tot_j^i = \hat{a} + \hat{b}\,j + \sum_{l=1}^{5} \hat{c}_l\, tot_{j-l}^i + e_{1,j}^i$, and $\Delta q_t = \hat{d} + \hat{e}\Delta tot_t + \hat{f}\,tot_{t-1} + \hat{g}\,q_{t-1} + e_{2,t}$.

In step 5, after running the regressions $q_{t+k}^i - q_t^i = a + b_k^{x,i}\, E_t^x(q_{t+k}^i - q_t^i) + e_{t+k,t}^i$, x = A,B, k = 1,...,8 we now collect the coefficients, $b_k^{x,i}$.

These 5,000 observations of $b_k^{x,i}$ for the two models, x = A,B and for each value of k form the Monte Carlo distributions under the assumption that the real exchange rate is described by model (C6) with parameter values (C7). They are used to test the null hypotheses, $H_{02}\!:b=1$, for k = 1,…, 8, with the results shown in column *(xi)* of Tables 5a and 5b.

C.2 Excess Returns

The third data generating process, DGP$_3$, combines the terms of trade and real exchange rate models defined by equations (C2) and (C6) with parameter values (C4) and (C7) – used for DGP$_2$ – with models for the Australian-foreign quarterly inflation differential, p_t^{diff}, and the one-year (four-quarter) Australian-foreign nominal interest differential, ${}_4i_t^{diff}$. p_t^{diff} is defined by $p_t^{diff} = p_t^{AUS} - p_t^*$, where $p_t^{AUS} = 100\times(p_t^{AUS} - p_{t-1}^{AUS})$ and p_t^* is the trade-weighted foreign quarterly inflation rate, $p_t^* = 100\times\sum_{j=1}^{5} w_j(p_t^j - p_{t-1}^j)$, and ${}_4i_t^{diff}$ is defined analogously.

Simple bi-variate auto-regressive time-series models are fitted for p_t^{diff} and ${}_4i_t^{diff}$ by OLS using data from 1987:Q1 to 1994:Q4 and allowing up to six lags for each variable. Eliminating insignificant lags leads to these models:

$$p_t^{diff} = -0.114 + 0.206 p_{t-4}^{diff} - 0.173 p_{t-6}^{diff} + 0.499 \; {}_4i_{t-2}^{diff} - 0.395 \; {}_4i_{t-3}^{diff} + 0.511 e_{3,t} \tag{C8}$$

$$\begin{aligned} {}_4i_t^{diff} &= 2.619 + 1.372 p_{t-3}^{diff} + 0.871 p_{t-4}^{diff} + 0.929 p_{t-5}^{diff} + 0.706 p_{t-6}^{diff} \\ &\quad + 0.559 \; {}_4i_{t-1}^{diff} - 0.502 \; {}_4i_{t-5}^{diff} + 0.727 e_{4,t} \end{aligned} \tag{C9}$$

where $e_{3,t}$ and $e_{4,t}$ are uncorrelated i.i.d. N(0,1) random variables.

Assuming this data generating process, DGP$_3$, Monte Carlo distributions are again generated by running 5,000 trials with each trial *i* proceeding as for DGP$_2$, but with the following modifications. As well as generating sequences of observations $\{tot_j^i\}_{j=6}^{n+T}$, $\{\Delta q_j^i\}_{j=6}^{n+T}$, we now also generate sequences of observations $\{p_j^{diff,i}\}_{j=6}^{n+T}$ and $\{{}_4i_j^{diff,i}\}_{j=6}^{n+T}$ using (C8) and (C9).

As for DGP$_2$, for time *t*, $71 \le t \le 100$, we use the sequences of synthetic data, $\{tot_j^i\}_{j=n+1}^{t-1}$ and $\{\Delta q_j^i\}_{j=n+1}^{t-1}$ to estimate real exchange rate model B (which involves estimating one of the three specifications in equation (8), depending on the time, *t*). With this real exchange rate model, together with estimates of $E_t(tot_{t+4}^i)$, we generate $E_t^B(q_{t+4}^i - q_t^i)$, *t* = 71,…, 100.

At time *t*, the expected and actual excess returns to holding a one-year Australian dollar bond rather than a trade-weighted basket of foreign bonds, are $E_t(ER_{t,4}) \approx 100 \times E_t^B(q_{t+4} - q_t) + E_t({}_4r_t^{AUS} - {}_4r_t^*)$ and $ER_{t,4} \approx 100 \times (q_{t+4} - q_t) + {}_4r_t^{AUS} - {}_4r_t^*$ which are equations (12) and (10) in the text. By construction, the expected real interest differential using backward-looking inflationary expectations is $E_t({}_4r_t^{AUS} - {}_4r_t^*) = {}_4i_t^{diff} - \sum_{j=1}^4 p_{t-j}^{diff}$, while the actual real interest differential is ${}_4r_t^{AUS} - {}_4r_t^* = {}_4i_t^{diff} - \sum_{j=1}^4 p_{t+j}^{diff}$. Thus, for trial *i*, the expected and actual excess returns are (approximately) $E_t(ER^i_{t,4}) = 100 \times E^B_t(q^i_{t+4} - q^i_t) + {}_4i^{diff,i}_t - \sum_{j=1}^4 p^{diff,i}_{t-j}$ and $ER^i_{t,4} = 100 \times (q^i_{t+4} - q^i_t) + {}_4i^{diff,i}_t - \sum_{j=1}^4 p^{diff,i}_{t+j}$.

For trial i, we use the values of $E_t(ER^i_{t,4})$ and $ER^i_{t,4}$, $71 \le t \le 98$, to run the regression $ER^i_{t,4} = a^i + b^i E_t(ER^i_{t,4}) + e_t$. The 5,000 values of b^i form the Monte Carlo distribution with which the hypotheses, $H_{01}\!: b = 0$ and $H_{02}\!: b = 1$ are tested. The results are shown in the first row of Table 6.

The fourth data generating process, DGP$_4$, combines the terms of trade and real exchange rate models defined by equations (C2) and (C6) with parameter values (C4) and (C7), with models for the Australian-foreign quarterly inflation differential, p^{diff}_t , and the two-year (eight-quarter) Australian-foreign nominal interest differential, ${}_8i^{diff}_t$.

As before, simple bi-variate auto-regressive time-series models are fitted for p^{diff}_t and ${}_8i^{diff}_t$ by OLS using data from 1987:Q1 to 1994:Q4 and allowing up to six lags for each variable. Eliminating insignificant lags leads to these models:

$$p^{diff}_t = -0.308 + 0.629\ {}_8i^{diff}_{t-2} - 0.293\ {}_8i^{diff}_{t-3} - 0.307\ {}_8i^{diff}_{t-4} + 0.125\ {}_8i^{diff}_{t-6} + 0.526 e_{3,t} \tag{C10}$$

$$\begin{aligned} {}_8i^{diff}_t = {} & 3.400 + 1.364 p^{diff}_{t-3} + 0.861 p^{diff}_{t-4} + 0.929 p^{diff}_{t-5} + 0.620 p^{diff}_{t-6} \\ & + 0.480\ {}_8i^{diff}_{t-2} - 0.712\ {}_8i^{diff}_{t-5} + 0.514 e_{4,t} \end{aligned} \tag{C11}$$

where, again, $e_{3,t}$ and $e_{4,t}$ are uncorrelated i.i.d. N(0,1) random variables.

The Monte Carlo distributions are derived as for DGP$_3$, with the following modifications. The expected and actual excess returns to holding a two-year Australian dollar bond are $E_t(ER_{t,8}) \approx 100 \times E^B_t(q_{t+8} - q_t) + 2E_t({}_8r^{AUS}_t - {}_8r^*_t)$ and $ER_{t,8} \approx 100 \times (q_{t+8} - q_t) + 2 \times ({}_8r^{AUS}_t - {}_8r^*_t)$ which are equations (14) and (13) in the text. The expected real interest differential using backward-looking inflationary expectations is $E_t({}_8r_t^{AUS} - {}_8r_t^*) = {}_8i_t^{diff} - \sum_{j=1}^{4} p_{t-j}^{diff}$, while the actual real interest differential is ${}_8r_t^{AUS} - {}_8r_t^* = {}_8i_t^{diff} - \sum_{j=1}^{8} p_{t+j}^{diff} / 2$. Thus, for trial i, the expected and actual excess returns are (approximately) $E_t(ER_{t,8}^i) = 100 \times E_t^B(q_{t+8}^i - q_t^i) + 2 \times ({}_8i_t^{diff,i} - \sum_{j=1}^{4} p_{t-j}^{diff,i})$ and $ER_{t,8}^i = 100 \times (q_{t+8}^i - q_t^i) + 2 \times {}_8i_t^{diff,i} - \sum_{j=1}^{8} p_{t+j}^{diff,i}$.

For trial i, we use the values of $E_t(ER_{t,8}^i)$ and $ER_{t,8}^i$, $71 \leq t \leq 94$, to run the regression $ER_{t,8}^i = a^i + b^i\, E_t(ER_{t,8}^i) + e_t$. Again, the 5,000 values of b^i form the Monte Carlo distribution with which the hypotheses, $H_{01}: b = 0$ and $H_{02}: b = 1$ are tested. The results are shown in the second row of Table 6.

Appendix D: Data

The terms of trade, defined as the ratio of the implicit price deflators for exports of goods and services to imports of goods and services, are from Table H.3 in the Reserve Bank of Australia (RBA) *Bulletin* Database.

Nominal bilateral exchange rates are end of quarter 4 pm (Sydney) quotations, from Table F.9 of the RBA *Bulletin* Database.

Consumer Price Index (CPI) data is of quarterly frequency and obtained from the International Monetary Fund's International Financial Statistics (IFS) database. For Australia, the 'Medicare adjusted' CPI series is used.

Nominal one and two year interest rates at the end of quarter t are the daily rates prevailing on the next day (ie, the first day of quarter $t+1$).[11]

The following one-year nominal interest rates are obtained from Datastream. Japan: 1 year London euro-yen rate; United States: 1 year treasury bill rate;

[11] As reported above, we use nominal exchange rates at the end (4pm) of the last day of quarter t. Since one and two year interest rates change very little from one day to the next, interest rates prevailing on the next day, which are readily available, are satisfactory for our purposes.

United Kingdom: 1 year interbank rate; New Zealand: 1 year interbank rate; Germany: 1 year treasury bond rate; Australia: 1 year interbank rate.

The two-year nominal interest rates are defined as follows. Japan: linear interpolation of the 7 year Japanese bond rate and 1 year London euro-yen rate; United States: 2 year US treasury bond rate; UK: linear interpolation of the 5 year gilt rate and 1 year interbank rate; NZ: 2 year government bond rate (from 1987:Q1-1992:Q3), thereafter, a linear interpolation of the 5 year government bond and 1 year interbank rate; Germany: 2 year German bond rate; Australia: 2 year Commonwealth Government bond rate.

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