A RANDOM WALK AROUND THE $A: EXPECTATIONS, RISK, INTEREST RATES AND CONSEQUENCES FOR EXTERNAL IMBALANCE

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ABSTRACT

Given essentially perfect capital mobility, Australian interest rates and the expected exchange rate change should satisfy international arbitrage conditions. We examine an arbitrage condition for a US investor, with a view to explaining the large short-term real interest differential between Australia and the US since late 1984. We have some evidence for a risk premium until late 1985. Since then, we explain the differential as a result of foreign exchange market inefficiency or as a consequence of the market having continually and rationally expected significant real devaluation of the $A. We provide evidence for both these explanations and draw implications for the current debate on Australia's external imbalance.

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I. INTRODUCTION

Australia is a small open economy operating with essentially no impediments to the movement of capital into or out of the country. As a consequence, Australian interest rates should satisfy international arbitrage conditions. The arbitrage condition for a representative US investor can be expressed either in terms of nominal or real interest rates. Thus,

$$i^A = i^{US} - E(\Delta S / S) + rp \tag{1}$$

or

$$r^A = r^{US} - E(\Delta S^R / S^R) + rp^R, \tag{2}$$

where i and r denote the nominal and real interest rates for some asset, S is the nominal \$US/\$A exchange rate, ΔS is the change in S over the life of the asset, and the superscript R denotes real. The risk premia, rp and rp^R, are the excess returns demanded by a US investor to hold the Australian denominated asset.[1]

This paper presents a detailed examination of these two equations. Our almost exclusive focus is on short-term nominal assets with the horizon of our analysis ranging from one week to three months. The paper is laid out as follows. Section II presents evidence on the forward rate and on survey market

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[1] The term 'risk premium' is often used loosely to mean the excess return demanded by investors to compensate them for the 'risk' of an exchange rate fall. In this paper, we use the term in its technical sense. For a given expectation of the return on an asset, the risk premium is the excess return required because of the expected distribution of that return which may be summarised by the expected higher moments of the return on the asset. Equations (1) and (2) are approximations because, for example, $(1 + i^A)/(1 + i^{US})$ is only approximately equal to $1 + i^A - i^{US}$. In the paper, the exact expressions are used when required.

expectations as predictors of the future spot exchange rate. In section III, after drawing implications from the survey on the size of the risk premium, two theoretical models are presented of the risk premium necessary to induce a representative US consumer-investor to hold a small proportion of assets in short-term nominal Australian assets. Section IV discusses the behaviour over the past fifteen years of consumer price inflation and short-term nominal and real interest rates for several OECD economies. The fifth section demonstrates that since Dec 83, the $A, unlike all the other currencies we examine, has exhibited significant skewness because of many large rapid unpredictable depreciations. The final section discusses the results of the paper in terms of either (i) a time-varying risk premium, or (ii) an inefficient foreign exchange market, or (iii) a peso problem (see definition on page 39) for the $A. Since late 85, all our evidence is that the risk premium has been much too small to explain the short-term real interest differential between Australia and the US, and so we focus on the latter two explanations. Evidence on the inefficiency of the foreign exchange market can provide a rationalization of the results – but a puzzle remains. As an alternative, we provide evidence that the $A suffers from a peso problem because of a market perception that, in the longer run, the real economy must adjust to put Australia on a sustainable net external debt/ GDP path – with a lower real and nominal exchange rate during the adjustment process.

II. EXCHANGE RATE EXPECTATIONS AND THE FORWARD RATE

To begin, we define the notation to be used in the paper for spot and forward exchange rates. Let S_t [a/b] be the spot price of currency 'a' measured in currency 'b' at time t (so that an increase in S_t [a/b] represents an appreciation of currency b with respect to currency a) and let $F_{t,k}$ [a/b] be the price at time t, k periods forward. Further, let lower case exchange rate variables denote the natural log of upper case exchange rate variables (s_t [a/b] $\equiv \ln S_t$ [a/b] and $f_{t,k}$ [a/b] $\equiv \ln F_{t,k}$ [a/b]).

The forward rate

A standard way to test whether the forward rate is a biased predictor of the future spot rate (see e.g., Hodrick (1987), Goodhart (1988)) is to estimate the equation

$$\Delta s_{t+k} = \alpha + \beta\, fd_{t,k} + \eta_{t+k}, \qquad (3)$$

where $\Delta s_{t+k} = s_{t+k} - s_t$, $fd_{t,k} = f_{t,k} - s_t$ is the current k-period forward discount[2] and η_{t+k} is the error term. If the change at time t predicted by the (log) forward discount is an unbiased predictor of the actual (log) change in the spot exchange rate, then $\alpha = 0$ and $\beta = 1$. By contrast, if $\beta = 0$, the forward rate (or forward discount) tells us nothing about the future movement of the spot rate. Finally, if $\beta < 0$, on average over the next k periods, the deviation from trend of the spot rate is in the opposite direction to the deviation predicted by the forward discount.

Table 1 reports regressions of equation (3) for the \$A / \$US market over two different time periods for both a four week and a thirteen week horizon. The reason for the two different time periods and technical details concerning the regressions are discussed in Appendix A. Details about the exchange rate dataset (dataset A) are provided in the Data Appendix. In all cases in which data from overlapping time periods are used (Table 1 is the first example), the standard errors (SE) of the estimates are evaluated using the technique of Hansen (1982).

There is by now "considerable evidence for a variety of currencies and sample periods ... that indicates a strong rejection of the proposition that the forward rate is an unbiased predictor of the future spot rate" (Hodrick, 1987, p. 54). To give a recent example, Goodhart (1988) estimates equation (3) for nine datasets (using daily, weekly and monthly exchange rate data on four currencies against the \$US). Six of his point estimates of β are negative and in five cases his estimate is

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[2] Given our definitions of s and f, if for example s_t is the log \$US/\$A rate and $fd_{t,k}$ is positive, then in the forward market the \$US is trading at a discount compared to the spot market.

significantly (more than two standard errors) less than 1. By contrast, none of his estimates are significantly different from 0. That is, he cannot reject the hypothesis that the forward discount has no capacity to explain future movement of the spot rate.

TABLE 1

ESTIMATES OF COEFFICIENTS FOR EQUATION (3)

k (Weeks)	Intercept		Coefficient on Forward discount			
	Coefficient	SE	Coefficient	SE	R^2	DW
Four*	– 0.011	0.009	– 1.93	1.59	0.023	0.40
Four+	– 0.007	0.014	– 1.67	2.15	0.011	0.47
Thirteen*	– 0.028	0.016	– 0.35	1.22	0.003	0.12
Thirteen+	– 0.039	0.025	– 1.20	1.72	0.022	0.12

Dataset A: * 6 Jan 84 to 21 Apr 89; + 15 Feb 85 to 21 Apr 89.

For Australia, Thorpe et. al. (1988) estimate equation (3) over one month, three month and six month horizons using a trade-weighted exchange rate measure (formed from bilateral exchange rates with the US, Japan, West Germany and Britain). Their period of estimation ranges from Dec 83 – Sept 87 to Nov 84 – Jan 88, and each of their point estimates of β is negative – but the level of significance is low. Only for a one month horizon can they reject H_0: $\beta = 1$ at a 5% level of significance.[3] All the regressions in Table 1 are consistent with this pattern – negative point estimates of β – but again the level of significance is low. Only in the case of the first reported regression can the hypothesis H_0: $\beta = 1$ be rejected at the 5% level in favour of the alternative H_1: $\beta < 1$.

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[3] For this regression, they also reject $\beta = 0$ at a 5% level of significance.

Survey data on exchange rate expectations

Between Mar 85 and Sept 87, *The Australian* newspaper published the results of a market survey of expectations of the \$US/\$A conducted by Dr. Ben Hunt. For almost the whole sample period sixteen of the major companies involved in the foreign exchange market took part in the survey. Every Friday between 2pm and 5pm each company's chief foreign exchange dealer was asked his/her expectation of the spot value of the \$US/\$A at 3pm the following Friday and at 3pm in four weeks time (see Hunt, 1987). Further details of the survey are in the Data Appendix. In this sub-section we analyse the four week expectational data.

Figures 1 and 2 show comparisons of the four week exchange rate expectations with the one month forward rate, the four-weekly inflation differential between the US and Australia ($\pi^{us}_t - \pi^{A}_t$),[4] and the behaviour of the spot rate. Define s^e_{t+k} as the (log of the) mean of the market participants' expectations of the spot rate in k weeks,[5] and Δs^e_{t+k} by $\Delta s^e_{t+k} = s^e_{t+k} - s_t$. Judged by the root mean square error (RMSE) over the sample period, both the average market participants' four-week forecast (s^e_{t+4}) and the forward rate ($f_{t,1\ month}$) are marginally worse forecasts of the spot rate in four weeks than the no-change forecast ($s^e_{t+4} \equiv s_t$) one would use if one thought the exchange rate was a martingale or a random walk without drift.

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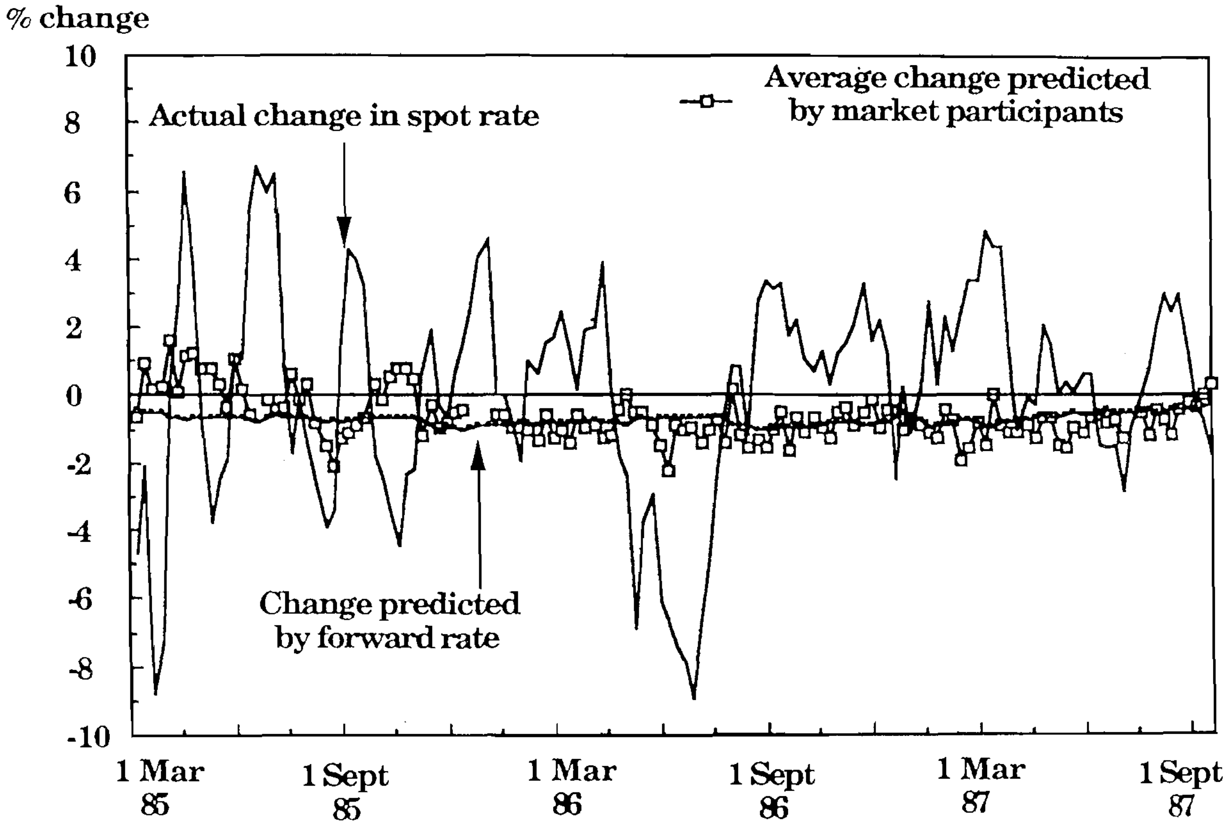
4 To derive inflation, we use price indices which had been published when the expectations were formed. Australian quarterly CPI numbers were always published by the end of the month immediately following the end of each quarter. For each Friday in month i we therefore define annual Australian inflation as $P_{i-4} / P_{i-16} - 1$, where P_i is the CPI for the quarter containing month i. For these numbers, we adjust for the 'Medicare effect' – which affects the estimates for the first twenty one weeks. Defining p_i as the monthly CPI for the US, annual US inflation is defined as $p_{i-3} / p_{i-15} - 1$. Four weekly inflation is derived from these annual inflation figures.

5 Much existing literature imagines that there is a single expectation that is homogeneously held by investors. We assume that this expectation is being measured by the mean of the responses. One can take the somewhat more sophisticated view that we measure the 'true investor expectation' with a measurement error (Froot and Frankel, 1989). Our econometric tests remain valid provided this measurement error is random (and hence uncorrelated with information available when the expectation is formed).

The root mean square errors are respectively, 3.3% and 3.2% compared with a RMSE for the no-change forecast of 3.1%.[6] Table 2 reports estimates of equation (3) for the time period of the expectational data as well as four other equations.

Figure 1

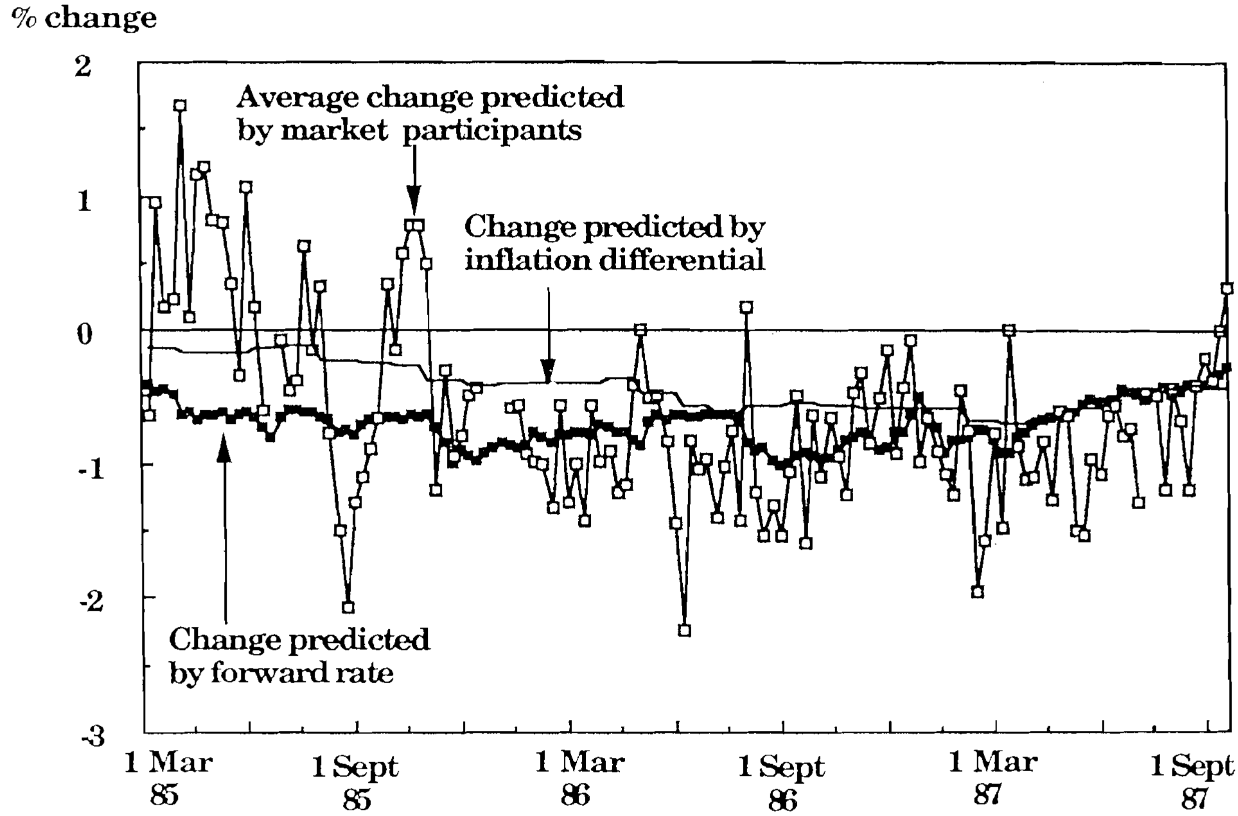
Average four week forecast, forward rate and change in spot rate for \$US/\$A



6 This result tallies with others who find that over a short horizon, market participants' forecasts of the future exchange rate are often worse, but never significantly better than a 'no-change forecast' – Lowe and Trevor (1986), Hunt (1987) and Manzur (1988).

Figure 2

Average four week forecast compared with the forward rate and the inflation differential



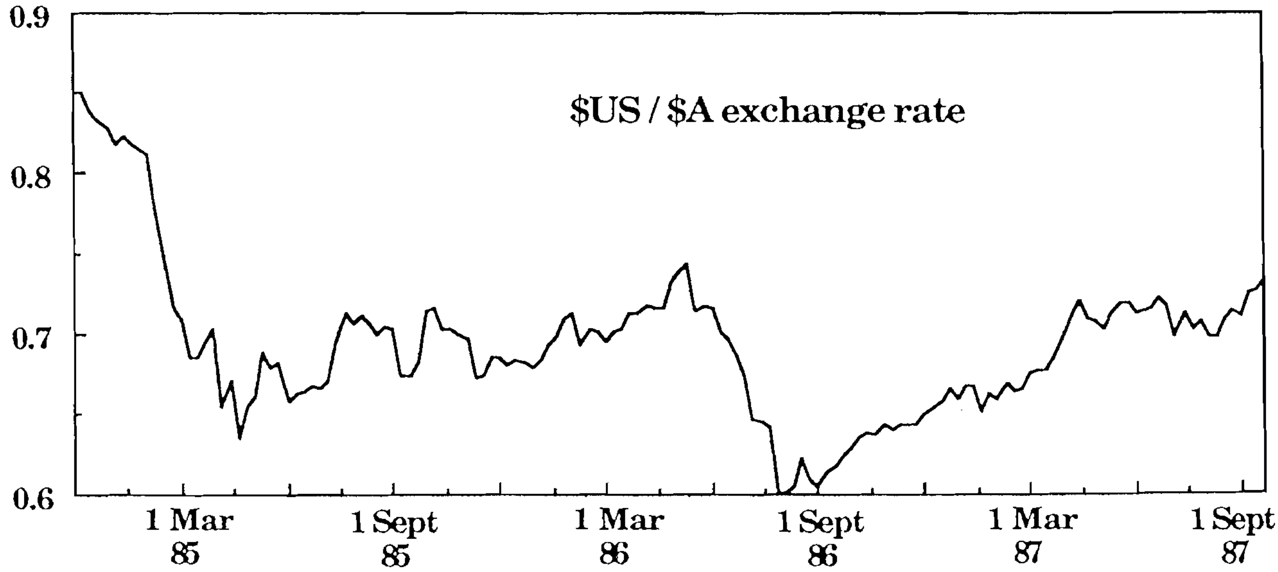


TABLE 2

ESTIMATES OF COEFFICIENTS FOR EQUATIONS (3 – 7)

					R^2	DW
Equation (3)	$\Delta s_{t+4} =$	α +	$\beta \; fd_{t,\,1\,month}$			
Estimates		– 0.044	– 6.18		0.098	0.47
Standard errors (SE)		(0.023)	(3.11)			
Equation (4)	$\Delta s_{t+4} =$					
	α +	$\beta \; fd_{t,\,1\,month}$	+ $\gamma \; (\pi^{us}_t - \pi^A_t)$			
Estimates	– 0.044	– 6.22	0.10		0.098	0.47
SE	(0.024)	(3.24)	(3.16)			
Equation (5)	$\Delta s^e_{t+4} =$	α +	$\beta \; fd_{t,\,1\,month}$			
Estimates		0.002	1.20		0.071	0.96
SE		(0.005)	(0.69)			
Equation (6)	$\Delta s^e_{t+4} =$					
	α +	$\beta \; fd_{t,\,1\,month}$	+ $\gamma \; (\pi^{us}_t - \pi^A_t)$			
Estimates	0.0065	0.514	2.10		0.309	1.25
SE	(0.0040)	(0.547)	(0.53)			
Equation (7)	$s^e_{t+4} -$	$s_{t+4} =$				
	α +	$\beta \; fd_{t,\,1\,month}$				
Estimates	0.047	7.38			0.131	0.47
SE	(0.023)	(3.14)				

Survey and exchange rate data: Weekly (Fridays), 8 Mar 85 to 18 Sept 87.

Estimation of equation (3) over the period of the exchange rate surveys again shows a negative point estimate of β (Table 2). The hypothesis H_0: $\beta = 1$ is rejected at the 5% level in favour of the alternative H_1: $\beta < 1$. Equation (5) shows that, on its own, the forward rate has significant explanatory power for the average exchange rate change expected by market participants. On the basis of estimates of this equation for their (fairly extensive) survey expectations data, Froot and Frankel (1989) conclude that "expectations seem to move very strongly with the forward rate", although they don't examine alternative explanators for expectations. The evidence from equation (6) is that our survey expectations seem to move strongly with the inflation differential rather than the forward rate. Importantly, this is in stark contrast with the actual exchange rate which shows no tendency to move in the direction predicted by either the forward discount or the inflation differential **either** over the sample period (equation 4) **or** during the whole period of the $A float (not shown). Equations (3) and (5) can be combined to give equation (7) which demonstrates that the average expectational error at time t exhibits statistically significant correlation with information available to the market at time t (the one month forward discount at t). This is evidence that either market expectations are not rational, or that over this period, the $A/$US market suffered from a 'peso problem'. These possibilities are examined in the discussion section.

Over the sample, the average annual inflation differential between Australia and the US was 5.4% p.a., the average annual rate of depreciation implied by the one month forward rate was 8.2% p.a., while market participants expected depreciation at an average rate of 7.8% p.a. On average, the actual rate **appreciated** at 2.7% p.a. Corresponding figures for the period Nov 85 – Sept 87 were 6.5% p.a. for the inflation differential, 8.5% p.a. for the one month forward rate, 10.5% p.a. for the average market participant compared to an average rate of **appreciation** of the actual rate of 2.8% p.a.

Thorpe et. al. (1988) report an extensive survey of market participants' exchange rate expectations at a one month horizon. The surveys were conducted weekly from Nov 84 to Mar 88 and used to estimate expectations of a trade-weighted $A exchange rate (TWA, to distinguish it from the Reserve Bank's TWI). Their study shows interesting similarities with our survey results. They fit equation (7) for their expectations data and conclude, as we do, that the average expectational error at time t is correlated with the forward discount at t. From Nov 84 to Dec 85, their market participants expected nominal appreciation of their TWA at an average rate of 4.7% p.a. This seems to correspond quite well with our expectations results for most of 1985 (see Figure 2). From Jan 86 to the end of their sample in Mar 88, their participants expected nominal depreciation of their TWA at an average rate of 8.6% p.a. – again somewhat comparable to our results.

An objection to the use of survey data to draw inferences about market expectations is this: "Consider the incentive problem of a trader who possesses private information that he has used to construct a portfolio of positions based on the deviations of his expectations from the current forward rates. When [someone] calls him for his expectations, will he reveal his information, or will he lie and quote something like the forward rate? Does he even know that the two are different?" (Hodrick, 1987 p. 135) Two points are worth making in response to this objection. Firstly, Goodhart (1988) provides evidence that foreign exchange traders – and the banks they work for – very seldom take open positions over a period as long as a month, which suggests that if traders have private information they very rarely think it sufficiently reliable to use it in the marketplace. Secondly, it is easy to understand the nature of private information in, for example, the stockmarket. In that market, there are obvious examples of potentially private information likely to affect a company's future profitability (e.g., the discovery of a valuable mineral ore deposit) and hence its stock price. But what is the nature of private information which, on average, helps to predict the value of the $US/$A exchange rate in four weeks time? If such information exists, it has been kept secret from macroeconomists for some time. Our best models of exchange rate determination over a four-week (or substantially longer) horizon do not perform significantly better than a random walk forecast (Meese and Rogoff, 1983). This second point sounds flippant, but is meant in all seriousness.

III. THE RISK PREMIUM

To begin, we briefly use some of the results from the previous section to provide measures of the nominal risk premium. Figure 3 shows a decomposition suggested by Froot and Frankel (1989) of the average market participant's expectational error, $s_{t+4} - s^e_{t+4}$, into the nominal risk premium, $rp = s^e_{t+4} - f_{t,1month}$, and the forward rate error, $fre = s_{t+4} - f_{t,1\,month}$,[7]

$$s_{t+4} - s^e_{t+4} = fre - rp.$$

For the estimate of the risk premium in Figure 3 to be accurate, it is necessary that the 'true investor expectation' is measured with **no** error – a requirement that seems rather stringent (see footnote 5). We can alternatively use equation (5) in Table 2 to test the null hypothesis that the risk premium is zero (equivalent to $\Delta s^e_{t+4} = fd_{t,1\,month}$; hence the null hypothesis is $\alpha = 0$, $\beta = 1$). Applying this test to our whole sample survey at a 5% level of significance, we accept this null of zero risk premium.[8]

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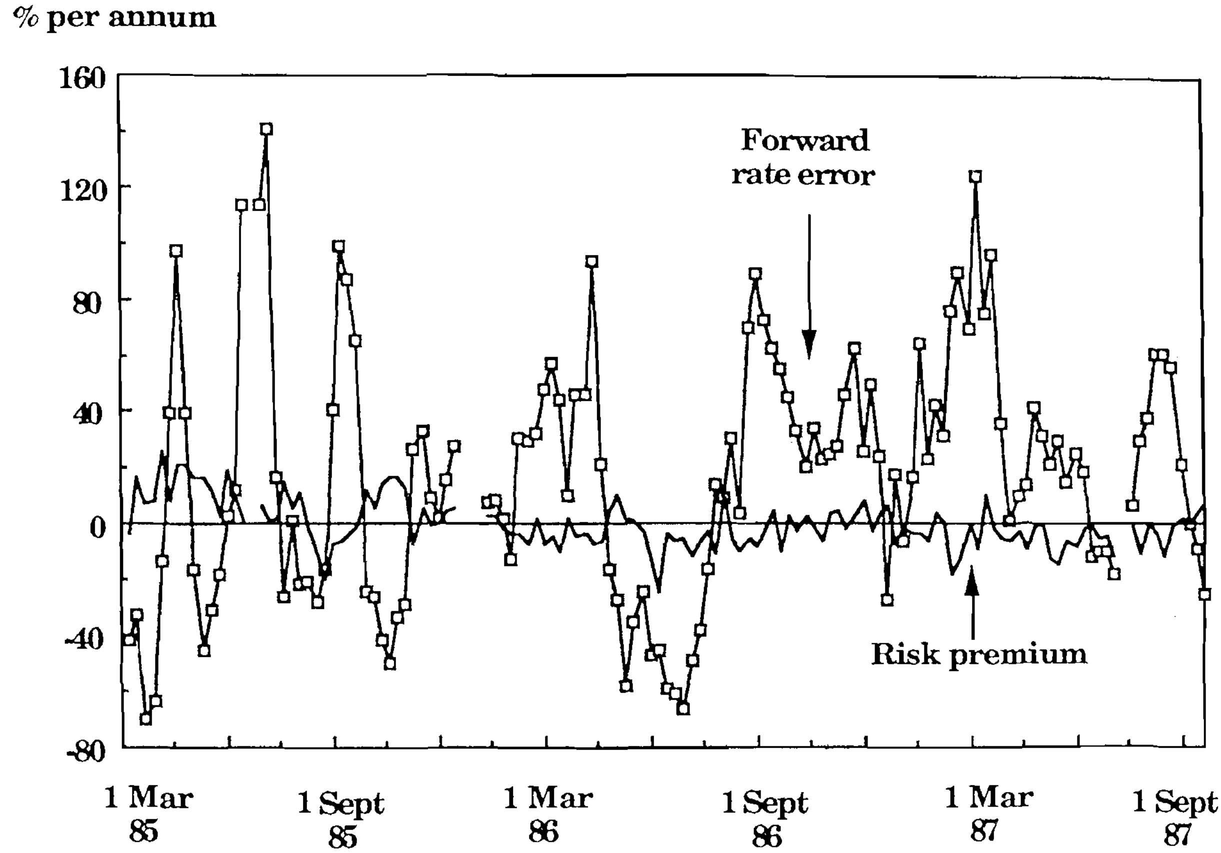
[7] It would make minimal difference to construct a four-week forward rate.

[8] Thorp et. al. (1988) estimate equation (5) for market expectations of their trade-weighted \$A. They find their point estimate $\beta = 3.52$ is very significantly different from $\beta = 1$, and hence they reject the null of zero risk premium. This result is dominated by their survey expectations from Nov 84 to late 85. For expectations formed from late 85 to the end of their survey in Mar 88, they accept the null of zero risk premium. Analysis of the sub-sample of our survey data from Mar to Oct 85 also reveals a statistically significant risk premium, which is not present in the sub-sample from Nov 85 to the end of the sample in Sept 87.

Figure 3

Nominal risk premium of the average market participant and the forward rate error



In the rest of this section we provide *a priori* estimates of the risk premium necessary to induce a US consumer-investor to hold a proportion of wealth in short-term Australian denominated interest-bearing bills. We present two simple single-period models in which a US consumer-investor maximizes end-of-period expected utility. This utility is derived from consumption of a basket of goods. The period of the model is four weeks.

In the first model, four weeks is assumed sufficiently short that the price index for the basket of goods at the end of this time, P^{us}_{t+1}, is known at the beginning of the period (i.e., it is not a source of uncertainty). The investor has an initial real wealth, W_t, and chooses to hold a proportion of this wealth, x, as an Australian bill with nominal return i^A with the remainder $(1 - x)$ held as a US bill with nominal return i^{us}. Real end-of-period wealth, W_{t+1}, is therefore

$$W_{t+1} = \frac{W_t}{P^{us}_{t+1} / P^{us}_t} [(1 - x).(1 + i^{us}) + x. (1 + i^A). (S_{t+1} / S_t)]$$

where, S is the spot exchange rate, S = \$US/\$A. Defining δ by

$$1 + \delta \equiv (S_{t+1} / S_t).(1 + i^A) / (1 + i^{us}), \tag{8}$$

end-of-period wealth is given by

$$W_{t+1} = W_t. (1 + i^{us}) . [1 + x\delta] / (P^{us}_{t+1} / P^{us}_t).$$

The investor is assumed to have a constant relative risk aversion, β, and hence seeks to maximize expected utility of the form

$$EU = E [c^{1-\beta} / (1 - \beta)],$$

where consumption, c, satisfies $c = W_{t+1}$. By construction, the only source of uncertainty arises from movements in the \$US/\$A exchange rate which lead to δ being a random variable. As we find evidence that changes in the \$US/\$A exchange rate exhibit both skewness and leptokurtosis (see section V), we consider the first four moments of the distribution of δ. Thus, expanding $f(\delta) = [1 + x\delta]^{1-\beta}$ as a Taylor expansion around f(0) and including terms up to δ^4, leads to an expression for expected utility, EU(x), as a function of the share of initial wealth held in Australia, x. Expected utility maximization implies that $dEU(x) / dx = 0$, and imposing this condition gives

$$E\,\delta = \beta x E\,\delta^2 - \frac{\beta(\beta + 1)}{2} x^2 E\,\delta^3 + \frac{\beta(\beta + 1)(\beta + 2)}{6} x^3 E\,\delta^4. \tag{9}$$

This equation gives the nominal risk premium, $E\,\delta$ – equivalent to rp in equation (1) – which the Australian bill must pay to induce the investor to hold a proportion x of wealth in Australia, as a function of the higher moments of the distribution of δ. The coefficient of relative risk aversion, β, is thought to be about two (see, for example, Friend and Blume (1975) or Newberry and Stiglitz (1981)). Quantitative estimates of the risk premium are derived from Dataset A using the period Jan 84 to Apr 89 and assuming $\beta = 2$. As mentioned above, the period of the model is four weeks.[9] Sample estimates of the unconditional four-week moments[10] of δ are: $E\delta^2 = 1.27\times10^{-3}$, $E\delta^3 = -1.99\times10^{-5}$, and $E\delta^4 = 5.69\times10^{-6}$. These numbers are used in equation (9) to calculate the nominal risk premium as a function of the fraction of the portfolio held in Australia and the results are displayed in Figure 4.[11]

An estimate of the proportion of Australian assets held in an international asset portfolio can be derived from a survey reported in *The Economist* March 25, 1989. Nine international investment banks were asked for their best mix of investments (equities, bonds and 'cash') over the next 12 months. Their suggestions for the geographical location of their equity holdings (but not their bond or cash holdings) were reported. Of six who suggested putting part of their portfolio in Australia, the range of recommended Australian portfolio shares was $0.1\% \leq x \leq 2.5\%$. When examining Figure 4, we may use these numbers as a guide to estimating portfolio

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9 Despite suggestions to the contrary (Juttner and Luedecke, 1988), covered interest parity seems a very good approximation for the Australian–US foreign exchange market after the float of the $A (Frankel and Froot, 1987; Smith, 1989a). We therefore use the approximation: $(1 + i_t^A) / (1 + i_t^{us}) = (S_t / F_{t, 1\,month})^{12/13}$. The exponent 12 / 13 is used to convert from one month forward to four weeks forward. δ is therefore defined by: $1 + \delta = (S_{t+4} / S_t).(S_t / F_{t, 1\,month})^{12/13}$, where the subscript t is defined in weeks. Of course, almost all the variation in δ arises from the term (S_{t+4} / S_t).

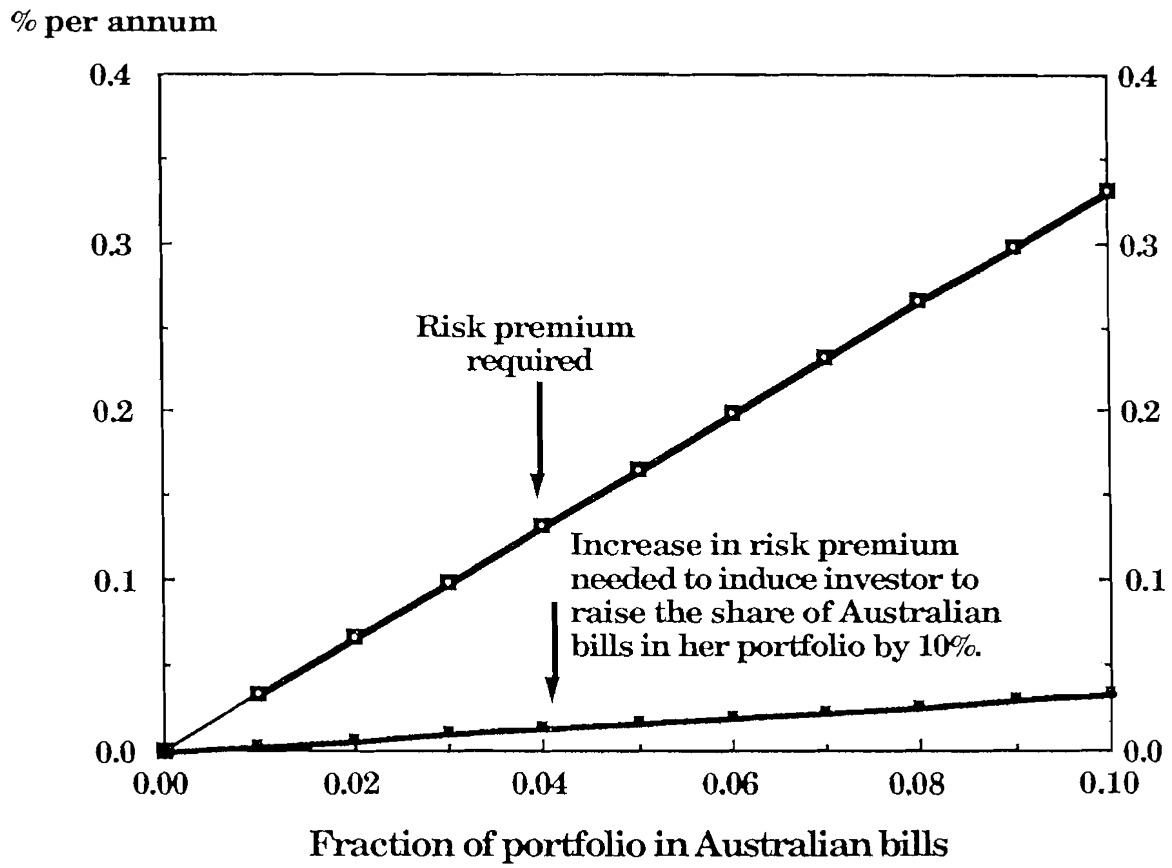
10 As Pagan (1988) stresses, the risk premium at any time is determined by the moments of δ **conditional** on information available at that time. The unconditional moments give an estimate of the average conditional moments over some extended period of time – a few weeks or a few months (see Frankel, 1988). Appendix B addresses this issue, but it does not change the conclusions presented here.

11 To stress the point: we do not evaluate the actual excess return on Australian bills over the sample period. Rather, we calculate the excess nominal return that a utility maximizing US investor requires because of the volatility – measured in $US – of the nominal return on her Australian asset. It is this excess return that constitutes the risk premium. We tested the sensitivity of the results to changes in some of the assumptions. Deriving the moments of δ using Dataset B, or alternatively, superimposing a constant depreciation of the $A of 10% p.a. (equivalent to 0.736% every four weeks) on the actual exchange rate data both lead to estimates of the risk premium almost identical to those shown in the Figure.

shares for Australian bills.[12] For $x \leq 2.5\%$, the risk premium required to induce such a holding is less than 0.09% p.a., and the increase in risk premium required to induce an investor to raise Australian asset holdings by 10% is less than 0.01%p.a.

Figure 4

Nominal risk premium required to hold short-term Australian bills



For the estimated values of the second, third and fourth moments of δ, and for values of x relevant to Australia, the effects of skewness and leptokurtosis on the risk premium are negligible – only the $E\delta^2$ term on the rhs of equation (9) contributes to the risk premium. Truncating the analysis beyond this term is

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[12] As will be discussed shortly, the representative consumer is probably not free to put all her financial resources into an international asset portfolio, and hence these numbers should be overestimates.

equivalent to applying mean-variance analysis. Such analysis can be conducted without some of the restrictive assumptions used above, and it forms the basis of our second simple model. Thus, we now (i) assume that the price index for the basket of goods at the end of the period, P^{us}_{t+1}, is unknown at time t, and (ii) allow our investor to choose nominal assets in Australia and four other countries (Canada, Japan, U.K. and West Germany) as well as the US. Let ρ^j be the real return on the nominal asset from country j, defined by

$$1+\rho^j \equiv \frac{1+i^j}{P^{us}_{t+1}/P^{us}_t} \cdot (S_{t+1}[j/\$US]/S_t[j/\$US])$$

where S [j / $US] is the price of currency j in US dollars (equal to one when j = US). Define ρ as the column vector of the five non-US real returns, ρ^j, $j \neq US$, and z as the column vector of real returns relative to the real return on the US asset:

$$z \equiv \rho - \iota\, \rho^{us}$$

where ι is a column vector of five ones.[13] Define x as the column vector of the five non-US portfolio shares – and thus the share allocated to the US asset is $(1 - x'\iota)$. Finally, define Ω as the variance-covariance matrix of the real excess return of the non-US assets

$$\Omega \equiv E\,(z-Ez).(z-Ez)'$$

where the expectation is formed at the beginning of the period under consideration. If the investor maximizes a function of the expected value and variance of her end-of-period wealth, then Frankel and Engel (1984) show that

$$E\,z = \beta \operatorname{cov}(z, \rho^{us}) + \beta\, \Omega\, x. \tag{10}$$

We now use equation (10) to derive quantitative estimates of E z^A, the expected excess real return on the Australian asset over the US asset, or equivalently, the real risk premium, rp^R from equation (2), which a US investor requires to hold a

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[13] z^A, the Australian element of z ($z^A \equiv \rho^A - \rho^{us}$), satisfies $z^A \equiv \delta.(1+i^{us}).P^{us}_t/P^{us}_{t+1}$, where δ is defined by equation (8).

fraction x^A of her portfolio in Australian bills. The first term on the rhs of equation (10) is zero if it is assumed that the price index P^{us}_{t+1} is known at time t. Even if the price index is unknown, this term should be very small, because inflation rates are so much less volatile than exchange rates. Sample estimates of the unconditional covariances and variance needed to calculate E z^A are provided in Table 3.[14] Appendix C describes how the estimates are derived.

TABLE 3

ESTIMATES OF VARIANCE AND COVARIANCES OVER FOUR WEEKS

var (z^A)	11.9×10^{-4} (2.55×10^{-4})	cov (z^A , $z^£$)	3.23×10^{-4} (1.60×10^{-4})
cov (z^A , ρ^{us})	0.249×10^{-4} (0.183×10^{-4})	cov (z^A , z^{DM})	2.76×10^{-4} (1.76×10^{-4})
cov (z^A , $z^¥$)	2.58×10^{-4} (1.58×10^{-4})	cov (z^A , z^C)	-0.90×10^{-4} (0.37×10^{-4})

To estimate E z^A, it is also necessary to have estimates of the portfolio share held outside the US. Remember that wealth, as defined here, is consumed at the end of the model. For the sake of realism, this wealth should include labour income, since this makes up a large proportion of average consumers' resources. Hence, for a representative consumer, we should expect that only a small proportion of wealth is available to be held as a portfolio of foreign assets.[15] As an illustrative calculation, assume that the total portfolio share held in countries other than the US or Australia is 0.10, and that this share is divided in proportion to GNP (so that

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[14] Again, it is the conditional moments which determine investor behaviour and the unconditional moments give an estimate of the average conditional moments over some extended period of time (a few weeks or a few months). See Appendix B.
[15] Consumers may not be able to make large borrowings against future labour income and/or may have substantial financial commitments (e.g., borrowing for a house) which limit their capacity to invest in foreign assets.

$x^{¥} = 0.053$, $x^{DM} = 0.024$, $x^{£} = 0.015$, $x^{C} = 0.008$).[16] Substituting these numbers and the point estimates in Table 3 into equation (10), and again assuming $\beta = 2$ gives

$$E\, z^{A} \;=\; [0.128 \;+\; 3.09\, x^{A}\,] \quad \%\text{ per annum.}$$

Thus, for example, to induce our investor to hold $x^{A} = 0.01$ of wealth in Australia (as well as 0.10 in other countries in the above proportions), a real return in Australia 0.16% p.a. higher than the real return in the US is required. The reader is invited to choose alternative portfolio shares, and evaluate the corresponding risk premia.

Finally, we estimate the increase in risk premium needed to induce our investor to raise the proportion of wealth held in Australia by Δx^{A}. It is no longer necessary to make specific assumptions about the portfolio shares held in other countries. Provided the increase in Australian holdings comes at the expense of assets held in the US, equation (10) implies

$$\Delta\, E\, z^{A} \;=\; 3.09\, \Delta\, x^{A} \quad \%\text{ per annum.}^{17} \qquad (11)$$

To induce the representative consumer to put an extra 0.01 (that is, 1%) of wealth in Australia, an extra average return in Australia of 0.031% p.a. is required. At the risk of appearing uncontroversial, this number seems rather small to us.

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[16] Based on 87 GNP using average exchange rates in Dec 87.
[17] With the minor exception of Canada, if the rise in Australian holdings is at the expense of other foreign holdings, the increase in risk premium is even smaller than equation (11) because of the positive values of cov (z^{A} , z^{j}), $j \neq C$.

IV. INFLATION AND INTEREST RATES

Figures 5 and 6 display, respectively, short-term nominal interest rates and 12 month ended CPI inflation rates for seven OECD countries over the period 1975 – 1989.[18] The expected short-term real interest rate differential between each country j and the US, r^j_{diff}, is given by

$$1 + r^{j}_{diff} = \frac{(1 + i^{j})/(1 + E\pi^{j})}{(1 + i^{us})/(1 + E\pi^{us})},$$

where $E\pi^j$ is expected inflation in country j. Figure 7 shows this differential using CPI inflation over the previous 12 months for $E\pi^j$.[19]

Over the 59 months Nov 84 –Sept 89, this measure of the short-term real interest differential between Australia and the US was positive **for all but four months**, and averaged +2.6% p.a. An alternative measure of expected real interest rates can be derived using the realized inflation rate during the ex post 3 month period as a proxy for expected inflation. By this measure, over the nineteen quarters from Dec Q 84 to Jun Q 89 the short-term real interest differential between Australia and the US was positive for sixteen quarters, and averaged +2.4% p.a.[20]

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[18] The interest rates are 3 month Treasury bill rates for Australia, US and Canada; the 6 month Treasury bill rate for Italy, and 3 month Eurocurrency rates for Japan, West Germany and United Kingdom. For Australia, the CPI numbers are quarterly, while for all other countries, they are monthly. The Australian numbers make adjustment for the 'Medicare effect', which increases the estimated Australian inflation rate from Mar Q 84 to Mar Q 85. See the data appendix for data sources for this section.

[19] To compare like with like, we use the 3 month US Eurocurrency interest rate for i^{us} when evaluating the real interest differential for Japan, West Germany and United Kingdom.

[20] Estimated in the same way over the same period, the real interest differential on 10 year government bonds between Australia and the US averaged – 0.6% p.a. Since we have some evidence for a risk premium in 84-85, but not since then, it is interesting to evaluate the ex post real interest rate differential between Australia and the US over the fourteen quarters Mar Q 86 to Jun Q 89. It averaged 2.8% p.a. for 3 month Treasury bills and 0.2% p.a. for 10 year bonds. McKibbin and Morling (1989) examine alternative measures of the real interest rate differential for 90 day

Frankel and MacArthur (1988) used this approach to measure the 3 month real interest differential between 24 countries and the US over the period Sept 82–Oct 86. Excluding the four closed economies in their sample (Bahrain, Greece, Mexico and South Africa), four of the remaining twenty countries (Austria, Denmark, Hong Kong and Switzerland) have an average real interest differential more than 2% p.a. below the US, while none have a real interest differential more than 2% p.a. above the US. As Frankel and MacArthur point out, this asymmetry occurs because of the high real interest rates in the US over this period. But the US continues to experience fairly high real interest rates.[21] Hence, this comparison with Frankel and MacArthur emphasises how unusually high Australian short term real interest rates have been since late 84.

A different perspective is provided by Table 4 which shows the excess nominal return a US investor would have achieved by investing (and continually rolling over the investment) in Australian 3 month Treasury bills rather than US bills. Of course, the results include the exchange rate change which occurs between the purchase and the sale of the bills.[22] Over thirty years, from Jan 60 to Sept 89, the average excess return from investing in Australian short-term bills has been

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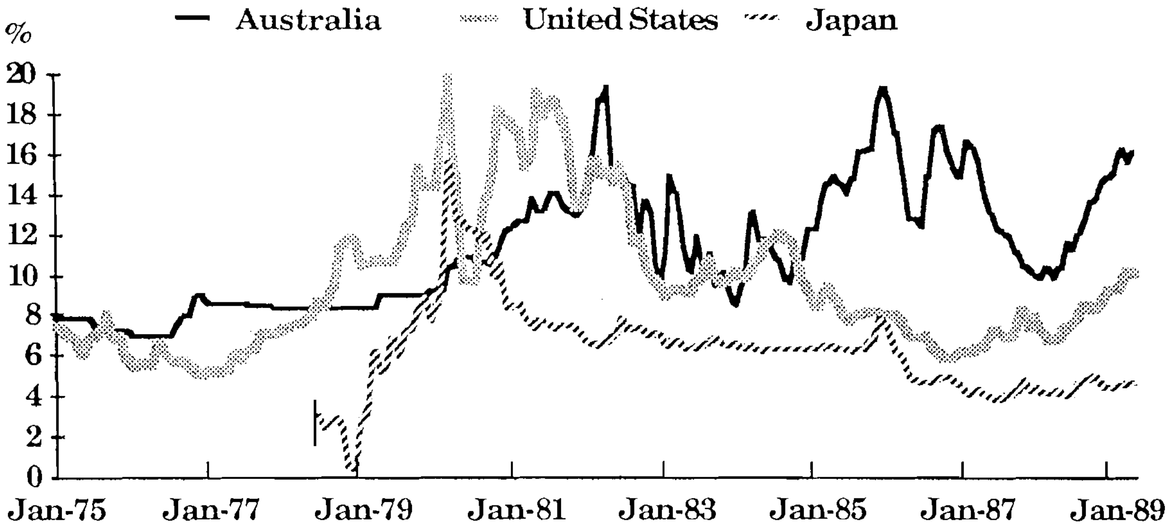
bank bills between Australia and US. From Dec Q 84 to Mar Q 89, using the quarterly change in the GDP (GNP) deflator for Australia (US) to estimate expected inflation, gives an average real interest rate differential of 2.3% p.a. Changing to a forecasting equation to estimate expected quarterly Australian GDP inflation (CPI inflation) leads to an average real interest differential over the period of 1.7% p.a. (2.5% p.a.). Finally, using a forecasting equation to estimate quarterly US GNP inflation and the forecasting equation to estimate expected quarterly Australian GDP inflation (CPI inflation) leads to an average real interest differential over the period of 1.9% p.a. (2.7% p.a.) – Steve Morling, personal communication.

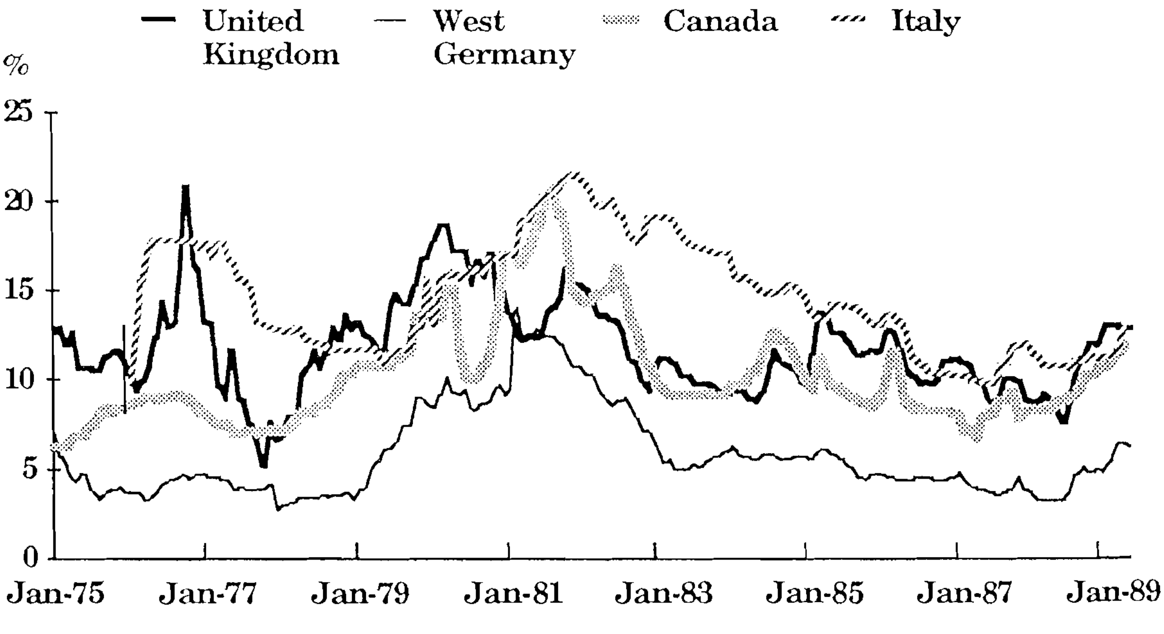
21 The ex post real US interest rate on 3 month Treasury bills averaged 5.4% p.a. from Sept 82 to Oct 86, 3.3% p.a. from Oct 84 to Jun 89, compared to an average of 2.0% p.a. from Jan 75 to Jun 89.

22 We are grateful to George Fane for suggesting this Table. It is constructed using end year exchange rates and annual average yields on 3 month Treasury notes from Tables S.7 and S.11 in Norton and Aylmer (1988) and from more recent RBA Bulletins.

Figure 5

Nominal Interest Rates

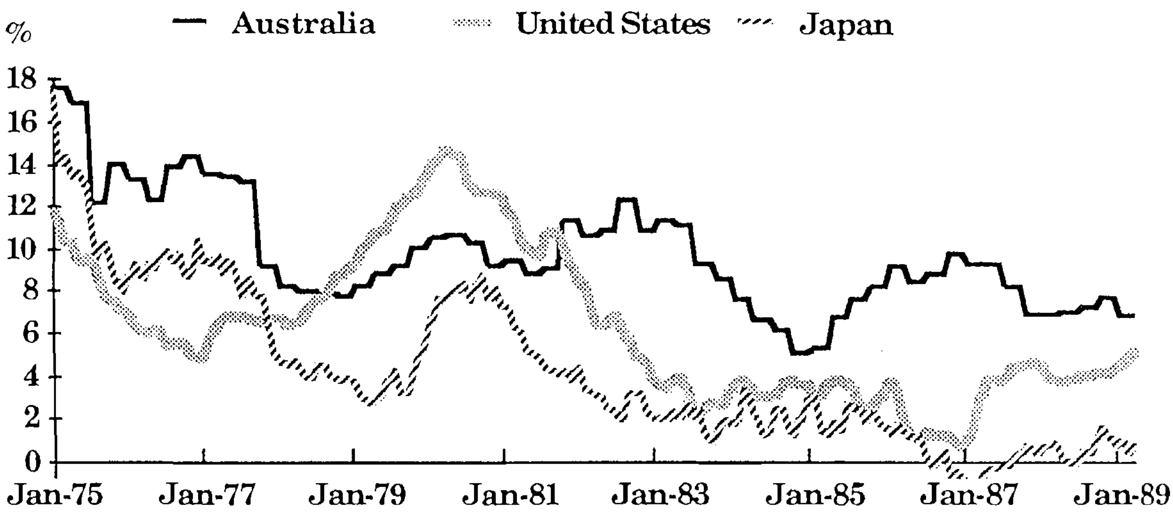


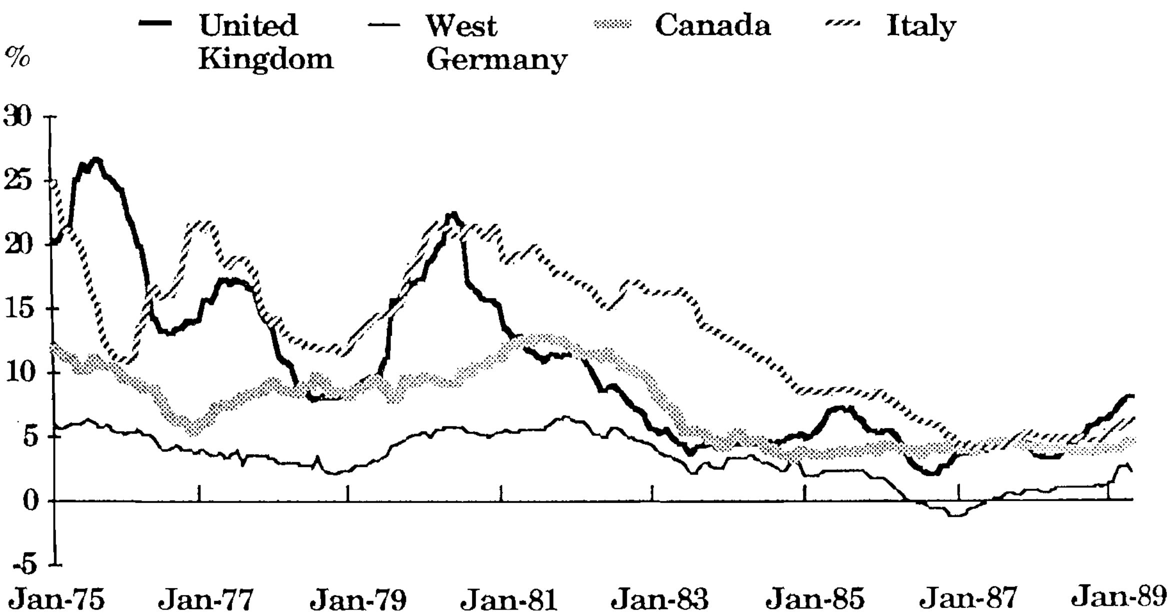


negligible (0.3% p.a.). But the Australian real interest premium since late 84 would have returned a US investor a nominal excess return of 5.7% p.a. on an investment from Jan 85 to Sept 89 – because over this time the $A experienced **real** appreciation against the $US at an average rate of 3.0% p.a. Even from Jan 84 –

Figure 6

12 Month Inflation Rate

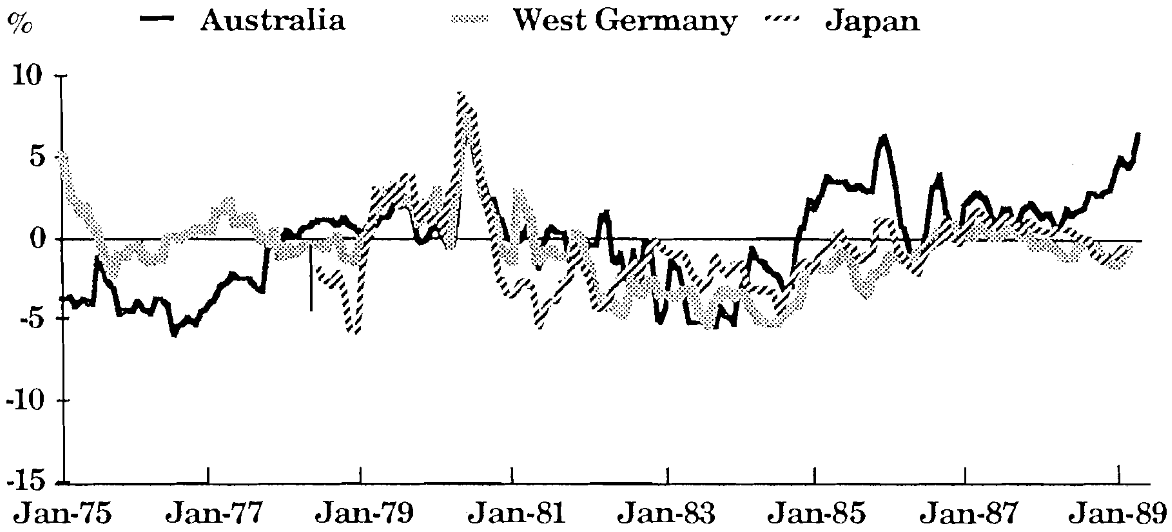




which includes virtually the whole period of the $A float – the excess return has averaged 3.3% p.a. as over this time the $A appreciated against the $US at an average **real** rate of 1.2% p.a.

Figure 7

Real Interest Rate Differential (cf U.S.)



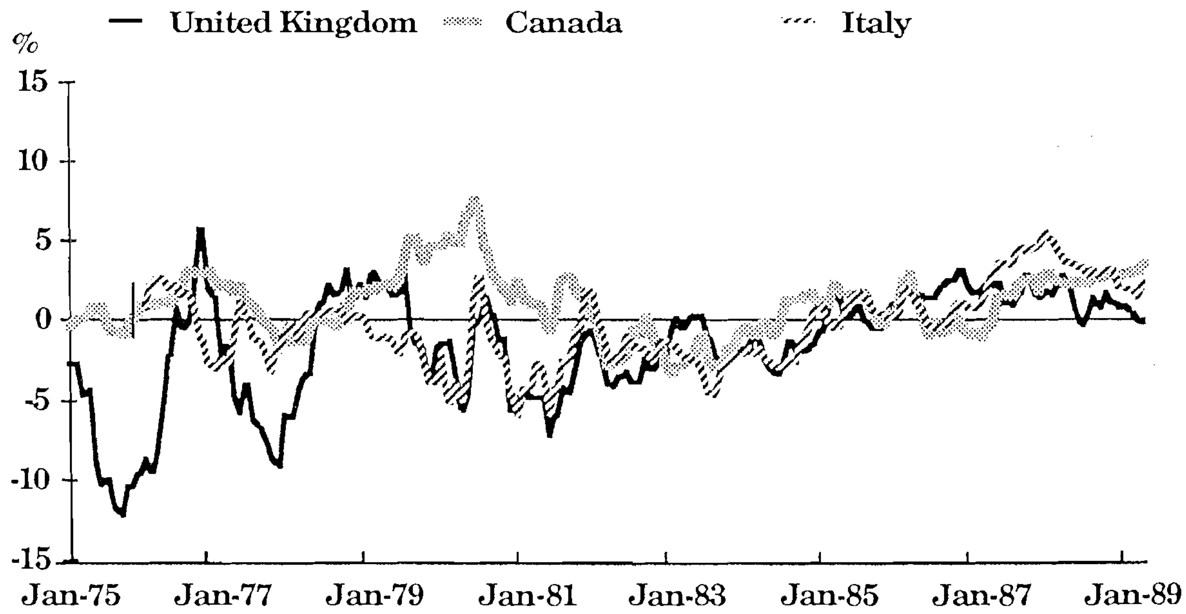


TABLE 4

NOMINAL EXCESS RETURN FROM INVESTING IN 3 MONTH AUSTRALIAN TREASURY BILLS RATHER THAN 3 MONTH U.S. TREASURY BILLS (EXPRESSED AS % P.A.)

	To the end of:									
	1964	**1969**	**1974**	**1979**	**1984**	**1985**	**1986**	**1987**	**1988**	**Sep 89**
From the start of: 1960	0.6	0.0	1.1	0.3	-0.7	-1.2	-0.9	-0.3	0.4	0.3
1965		-0.5	1.4	0.2	-1.0	-1.6	-1.2	-0.5	0.4	0.2
1970			3.4	0.6	-1.2	-1.9	-1.4	-0.5	0.7	0.4
1975				-2.2	-3.4	-4.2	-3.4	-2.0	-0.3	-0.6
1980					-4.6	-5.8	-4.2	-1.9	0.8	0.3
1984					-7.1	-9.4	-4.5	0.3	4.7	3.3
1985						-11.7	-3.1	2.8	7.9	5.7
1986							6.3	11.0	15.3	10.8
1987								15.9	20.1	12.5
1988									24.5	10.7
1989										-5.4

V. SKEWNESS

This section provides evidence about the skewness of the \$A against a range of currencies and against the trade-weighted index (TWI). There have been many studies examing the skewness of exchange rate returns over very short horizons (a day or less – see for example, Bewley et.al., (1987)). By contrast, our analysis examines skewness over a week and four weeks because we are interested in an horizon relevant to those investing in Australian assets for substantially longer than a day. In all cases we examine the behaviour of the random variable D_i [a/b], defined as

$$D_i\,[a/b] \;=\; 100 \times (\, s_{t+i}\,[a/b] - s_t\,[a/b]\,). \qquad (12)$$

There are two reasons for studying D_i [a/b]. Firstly, defining $\Delta_i\, S_t\,[a/b] \;=\; S_{t+i}\,[a/b] - S_t\,[a/b]$, it follows that

$$D_i\,[a/b] \;=\; 100 \times \;\ln(\,1 + \Delta_i\, S_t\,[a/b]\,/\,S_t\,[a/b]\,) \;\approx\; 100 \times \;\Delta_i\, S_t\,[a/b]\,/\,S_t\,[a/b],$$

provided $\Delta_i\, S_t\,/\,S_t$ is small compared to 1. Hence, D_i [a/b] is approximately the percentage return from converting currency a into currency b for i periods. Secondly, by definition, $D_i\,[b/a] \equiv -\,D_i\,[a/b]$ and so $-\,D_i$ [a/b] is approximately the percentage return from converting currency b into currency a for i periods.

TABLE 5: THIRD CENTRAL MOMENT OF D_i [a / \$A], i = ONE WEEK

	Currency a					
	\$US	**¥**	**£**	**DM**	**\$C**	**TWI**
Period						
Jan 79 – Dec 83	– 3.9	– 4.5	– 3.3	– 5.9	– 3.9	– 3.3
Jan 84 – Apr 89	– 6.1**	– 9.8**	– 8.8*	– 8.7**	– 5.3**	– 6.2**
Jan 86 – Apr 89	– 5.6*	– 12**	– 7.6	– 12*	na	– 6.4*

Periods 79 – 83 and 84 – 89: From dataset A. Period 86 – 89: From dataset B.

* Different from zero at 10% level of significance

** Different from zero at 5% level of significance

TABLE 6

THIRD CENTRAL MOMENT OF D_i [a / $ A]

Dataset A: weekly (Friday) data, Jan 84 – Apr 89.

Period	$ US	¥	£	DM	TWI
One week	– 6.11**	– 9.75**	– 8.83*	– 8.74**	– 6.22**
	(2.49)	(4.07)	(4.92)	(4.39)	(2.87)
Two weeks	– 10.85*	– 23.21**	– 16.99	– 16.93	– 13.75**
	(6.52)	(10.33)	(12.02)	(11.92)	(6.23)
Four weeks	– 40.10	– 79.20*	– 88.74	– 69.89	– 39.05*
	(31.70)	(46.33)	(69.64)	(48.16)	(22.22)

THIRD CENTRAL MOMENT OF D_i [a / $ A]

Dataset B: daily data, Jan 86 – Apr 89.

Period	$ US	¥	£	DM	TWI
One week	– 5.57*	– 12.04**	– 7.63	– 11.67*	– 6.42*
	(2.96)	(6.15)	(4.94)	(6.29)	(3.42)
Two weeks	– 15.25*	– 38.09*	– 25.67	– 38.51*	– 19.19*
	(8.67)	(20.06)	(18.46)	(22.52)	(10.59)
Month (22 working days)	– 40.61	– 119.76	– 51.69	– 114.17	– 54.97
	(32.52)	(90.94)	(60.64)	(87.27)	(38.64)

Standard errors in brackets

* Significantly different from zero at the 10% level

** Significantly different from zero at the 5% level

TABLE 7

CROSS COUNTRY COMPARISON OF TCM OF D_i [a / b], i = ONE WEEK

	Currency a					
	$A	$US	¥	£	DM	$C
Currency b						
$A	•	– 6.1**	– 9.8**	– 8.8*	– 8.7**	– 5.3**
$US	6.1**	•	– 1.9	– 2.0	– 1.4	0.0
¥	9.8**	1.9	•	0.7	0.0	1.6
£	8.8*	2.0	– 0.7	•	– 0.3	1.4
DM	8.7**	1.4	0.0	0.3	•	1.0
$C	5.3**	0.0	– 1.6	– 1.4	– 1.0	•

CROSS COUNTRY COMPARISON OF TCM OF D_i [a / b], i = FOUR WEEKS

	Currency a					
	$A	$US	¥	£	DM	$C
Currency b						
$A	•	– 40	– 79*	– 89	– 70+	– 22
$US	40	•	– 23+	– 27	– 10	0.9
¥	79*	23+	•	– 1.2	– 0.3	20
£	89	27	1.2	•	– 3.4	18
DM	70+	10	0.3	3.4	•	6.8
$C	22	– 0.9	– 20	–18	–6.8	•

Dataset A: weekly (Friday) data, Jan 84 – Apr 89.

\+ Significantly different from zero at 20% level (**only** used for four-weekly results)

* Significantly different from zero at the 10% level

** Significantly different from zero at the 5% level

TABLE 8

CROSS COUNTRY COMPARISON OF TCM OF D_i [a / b], i = ONE WEEK (FIVE WORKING DAYS)

	Currency a				
Currency b	$A	$US	¥	£	DM
$A	•	– 5.6*	– 12**	– 7.6	– 12*
$US	5.6*	•	– 1.3	– 0.2	– 0.7
¥	12**	1.3	•	0.1	– 0.1
£	7.6	0.2	– 0.1	•	– 0.3
DM	12*	0.7	0.1	0.3	•

CROSS COUNTRY COMPARISON OF TCM OF D_i [a / b], i = ONE MONTH (TWENTY TWO WORKING DAYS)

	Currency a				
Currency b	$A	$US	¥	£	DM
$A	•	– 41	– 120+	– 52	– 114+
$US	41	•	– 18	– 1.5	– 7.9
¥	120+	18	•	22	3.3
£	52	1.5	– 22	•	– 7.5
DM	114+	7.9	– 3.3	7.5	•

Dataset B: daily data, Jan 86 – Apr 89.

\+ Significantly different from zero at 20% level (**only** used for monthly results)

* Significantly different from zero at the 10% level

** Significantly different from zero at the 5% level

TABLE 9

NON-PARAMETRIC TEST OF SKEWNESS OF D_i [a / b], i = ONE WEEK (FIVE WORKING DAYS)

Currency b \ Currency a	$A	$US	¥	£	DM
$A	•	(–) 5**	(–) 3** 2*	(–) 2*	(–) 3** 1*
$US	(+) 5**	•	(–) 1*	0	(–) 1* (+) 1*
¥	(+) 3** 2*	(+) 1*	•	0	0
£	(+) 2*	0	0	•	0
DM	(+) 3** 1*	(+) 1* (–) 1*	0	0	•

Dataset B: daily data, Jan 86 – Apr 89.
Results of **one-sided** non-parametric test described in Appendix D. The sign (+) [(–)] indicates that significant positive [negative] skewness is found. The number indicates the number of working days of the week which exhibit significant skewness. The stars indicate the level of significance:

* significant at 5% level

** significant at 1% level

In Tables 5 – 8, we examine the sample values of the third central moment (tcm) of D_i [a/b], i.e., E { D_i [a/b] – E D_i [a/b] }^{3}, for a range of currencies and time periods. Table 5 demonstrates that for weekly changes, the $A may have been negatively skewed against most currencies since 1979, but that this skewness has become much larger and statistically significant since the $A was floated in Dec 83.[23] Since Jan 86, this skewness has remained large (and may have become larger – see the comparison of two week changes in Table 6).

Tables 7 and 8 show that the $A is much more skewed than other currencies. This is clear both from the size of the point estimates of the third central moment, and from their level of significance. These conclusions are supported by the results of a non-parametric test shown in Table 9. This test, which is described in Appendix D, has the advantage that under the null hypothesis, it is not necessary to assume that D_i [a/b] is normally distributed. The test is also approximately valid when the observations of D_i [a/b] come from different distributions. Figures 8 and 9 show the distributions of D_i [TWI/$A], D_i [¥/$A] and D_i [¥/$US] for i = one week, and for i = four weeks since Jan 86.[24] In Figure 9, D_i [¥/$US] is chosen because it is one of the most skewed distributions which does not involve the $A (see Table 8). In both Figures, the significant proportion of big depreciations of the $A is clearly not matched by an equal proportion of big appreciations.

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23 In common with others (e.g., Bewley et. al. 1987), we find evidence of leptokurtosis (which should be a disease, but in fact means a distribution with a larger fourth central moment than the normal distribution with the same variance). We find leptokurtosis in weekly log changes of the $A against most currencies since the float but **not** before it.

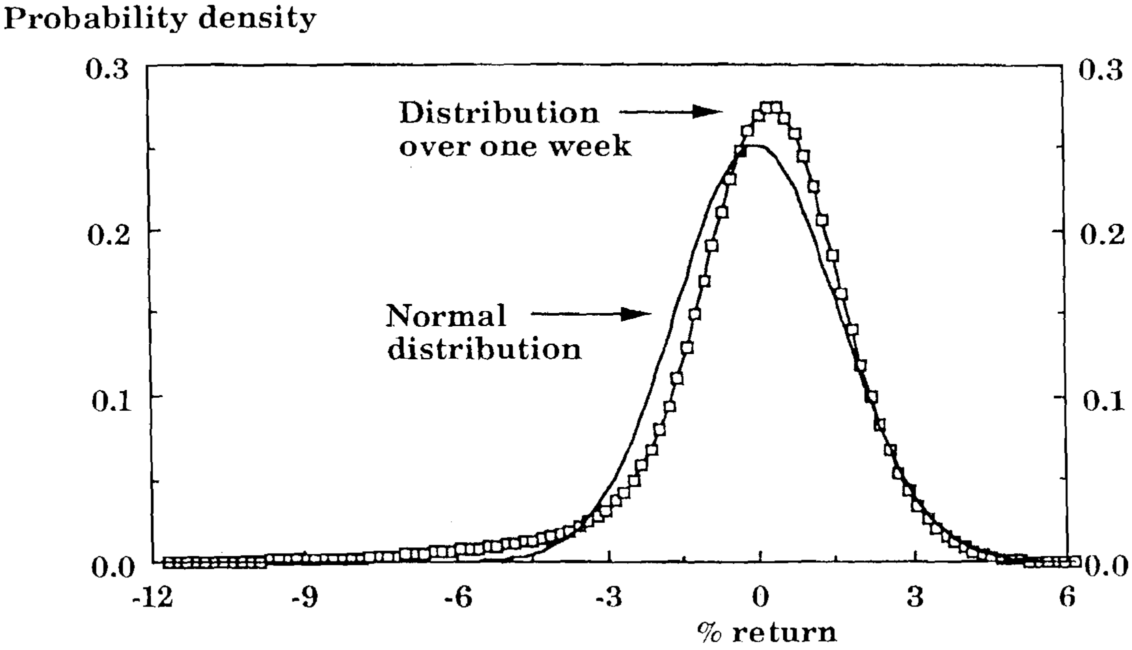
24 The distributions are derived by a non-parametric technique kindly suggested to us by Adrian Pagan. Given observations x_i, i = 1,...,N, from an unknown density function, f(x), the density function f(x*) at any point x* is estimated by

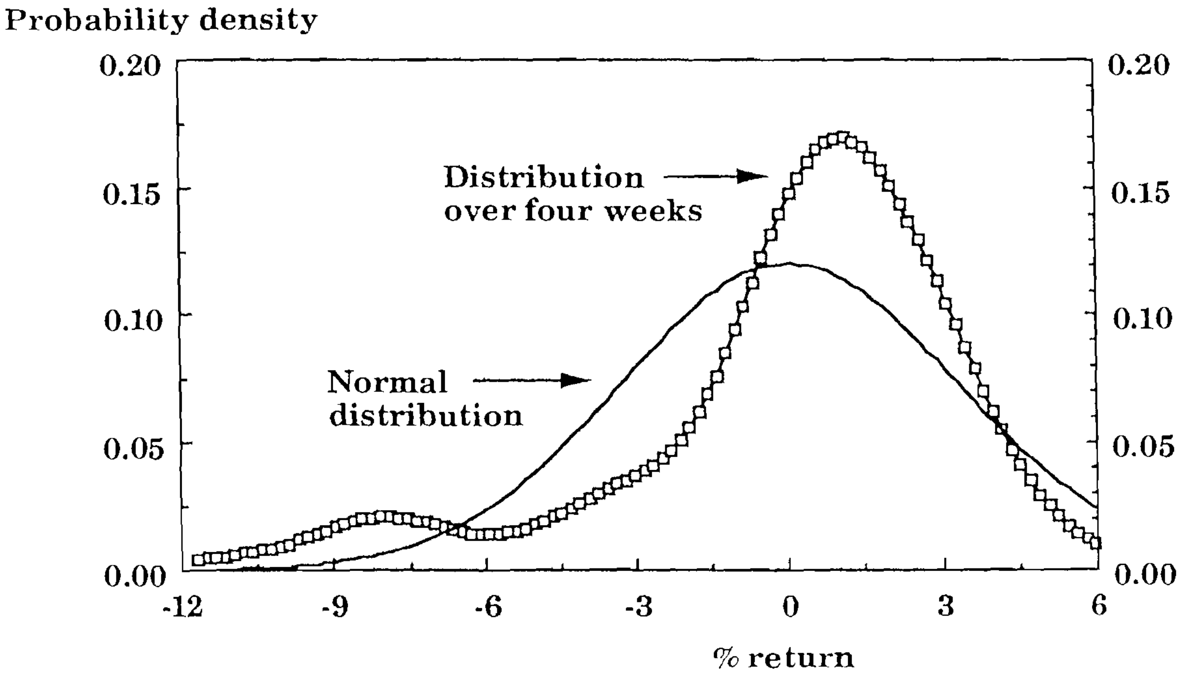
$$f(x^*) = \frac{1}{\sqrt{2\pi}\ Nh} \sum_{i=1}^{N} \exp\left(-\frac{1}{2}\left[\frac{x_i - x^*}{h}\right]^2 \right)$$

where $h = \sigma_x . N^{-1/5}$ and σ_x is the standard error of the observations x_i, i = 1,...,N.

Figure 8

Distribution of weekly and four weekly returns on the $A against the TWI

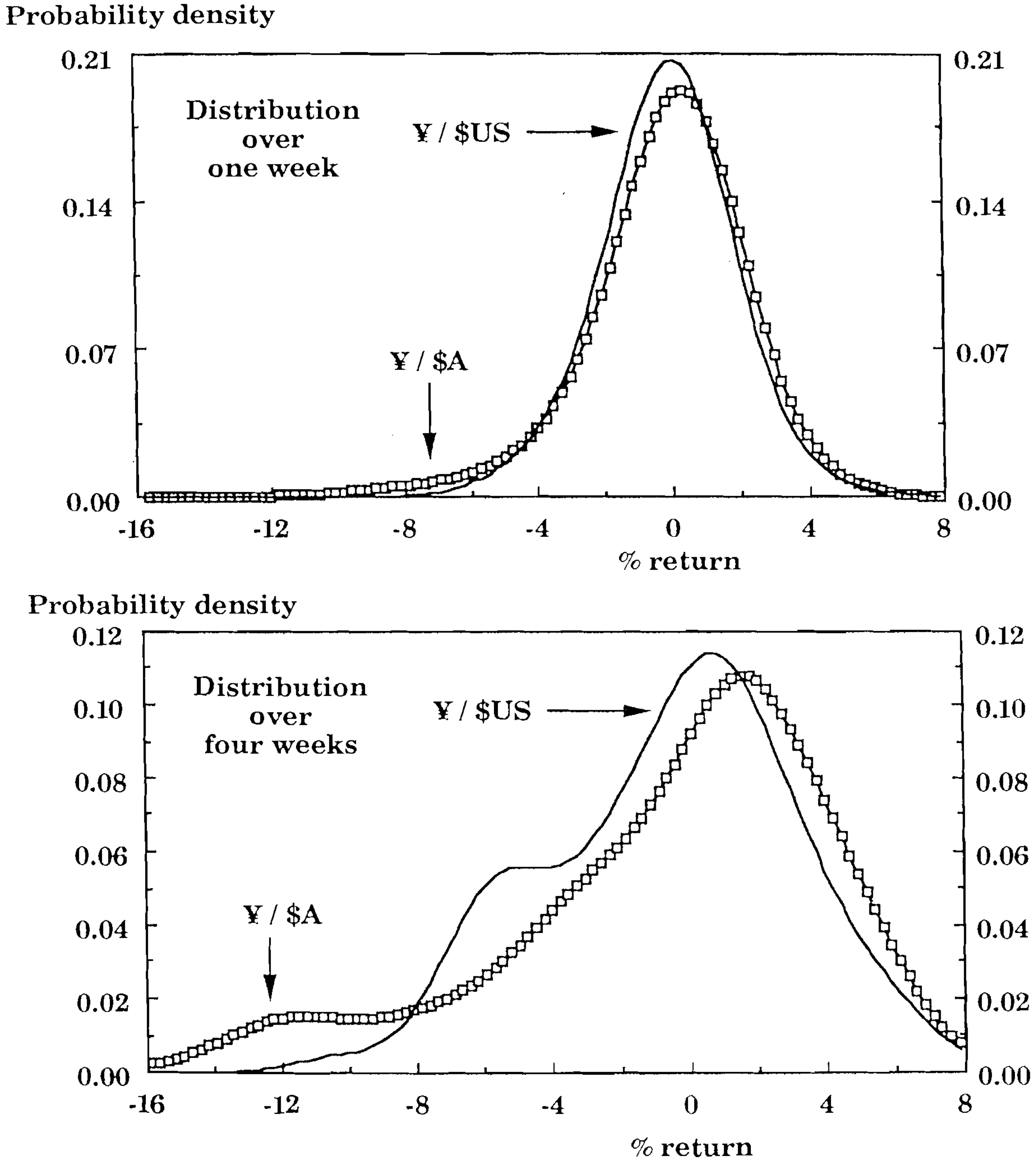




In both cases, the normal distribution has the same mean and variance as the actual distribution. Results are derived from dataset B: daily rates Jan 86 – Apr 89.

Figure 9

Distribution of weekly and four weekly returns: ¥ / $A and ¥ / $US



The ¥ / $US distributions have been shifted by + 0.08% and + 0.38% respectively, so that their means coincide with the means of the ¥ / $A distributions. Results are derived from dataset B: daily rates Jan 86 – Apr 89.

TABLE 10

Analysis of the ten largest weekly falls (Friday to Friday) of $A against the TWI, Jan 86 to Apr 89 using Dataset B.

Week	% fall	Reason – judged by reports during the week in the *Australian Financial Review.*
31 Jan – 7 Feb 86	3.2	No obvious event.
30 May – 6 June 86	4.6	No obvious event. "US investors are beginning to get nervous about the magnitude of the economic problems facing the Australian Government." Lead article, 5 June. This article is almost exclusively about adjustment to the external trade imbalance.
20 – 27 June 86	3.6	On 24 June, *TheWall Street Journal* editorial page is quoted asking: "Will Australia become the next Banana Republic?"
27 June – 4 July 86	5.0	Rumours that Mr. Keating had resigned as Treasurer. Removal of exemptions for withholding tax. Push by unions for wider superannuation coverage. Waterside Workers nationwide strike (resolved on 3 July).
18 – 25 July 86	4.5	Unexpectedly large June quarter CPI figure announced (CPI up 1.7%).
15 – 22 Aug 86	3.1	Unexpectedly large July current account deficit announced ($1.35bn).
9 – 16 Jan 87	4.2	No obvious event. "Sudden change of sentiment in foreign exchange market." Front page, 14 Jan. Perhaps related to the EMS realignment occurring at the time.
23 – 30 Oct 87	6.7	Delayed reaction to stock market crash (on 19, 20 Oct). "The world stockmarket crash has ...[left] Australia exposed because of its high foreign debt burden and dependence on commodity exports." Front page, 30 Oct.
24 June – 1 July 88	2.9	Global strength of $US
10 – 17 Feb 89	6.8	Unexpectedly large Jan current account deficit announced ($1.54bn). "The dollar is now diminishing our competitiveness. When demand conditions moderate I expect, and indeed hope, that the dollar will fall. And certainly, the day that starts we will not be standing in the way of stopping it." Mr. Keating, 16 Feb.

The data consists of 170 weekly changes. The median change is 0.19%. Of the ten largest deviations from this median (values of y_j from Appendix D), all ten are falls. They form the basis for this Table.

In this section, we have focussed our attention particularly on the period since Jan 86. Since then, most of the exogenous "news" from Australia's terms of trade has been favourable. The terms of trade had already fallen from 100.6 in Mar Q 85 to 91.9 in Dec Q 85. In 86, the fall continued at a slower rate to a low of 87.0 in Mar Q 87, then climbed rapidly to 113.4 by Mar Q 89 (from RBA Bulletin, Table K.3, Dec 88 and Jun 89). So, it is hard to sustain the argument that since Jan 86 Australia's external sector has suffered a series of unfavourable exogenous shocks which have required a succession of falls in the $A.[25]

Table 10 [26] provides evidence on the possible causes of the ten largest weekly falls of the $A from Jan 86 to Apr 89.[27] The large fall which followed the 1987 stockmarket crash suggests that the exchange rate is (sometimes) very sensitive to external shocks. Most of the other events which were identified **seem related to the unexpectedly slow progress being made by the real economy onto a sustainable net external debt/ GDP path.**[28] This comment applies most obviously to the second and

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[25] Of course, one could claim that the outcome was unfavourable compared to what was expected. But for the outcome to have appeared unfavourable, the expectations must have been for a massive improvement in the terms of trade. We are not aware of any such optimistic expectations.

[26] We use an Australian financial newspaper for Table 10 rather than a foreign one because Wong (1988) finds that most of the movements in the $A/$US exchange rate occur while the foreign exchange markets are open in Australia.

[27] As Frank Milne pointed out to us, the Table suffers from an important weakness. Rather than examining all events of a particular kind (e.g. all current account announcements) to test whether, on average, unexpected outcomes have an impact on the exchange rate, we search for the 'causes' of the biggest falls, *ex post*. We did use Dataset B to examine the percentage change in the $A against the TWI from the day before monthly current account announcements to four working days after ($s_{t+4} - s_{t-1}$, where t is the day of the release). We find a correlation coefficient between ($s_{t+4} - s_{t-1}$) and the nominal $A value of the announced current account deficit, of – 0.39. Surprisingly, we find an average value of ($s_{t+4} - s_{t-1}$) of 0.27% over the 39 announcements in the sample, compared to a mean weekly change for the whole sample of 0.006%. So, on average, during a week which included a current account announcement, the $A appreciated substantially more than during an arbitrarily chosen week.

[28] The debt/GDP ratio will increase (potentially without bound) if net exports are in deficit (as they are in Australia) **and** if the nominal interest rate exceeds the

third events as well as to the two current account announcements. We return to this theme in the final section of the paper.

VI. DISCUSSION

Despite continuous repetition of the claim that the efficient markets hypothesis implies that exchange rates should move as random walks with no drift, the claim is false. In fact, the joint hypothesis that (i) market-participants are risk-neutral, (ii) transactions costs are small enough to be ignored[29] and (iii) the market is efficient (or equivalently, market-participants form and use rational expectations) implies that the exchange rate must undergo a random walk around the forward rate.[30] As we have seen, there is now very strong international evidence – supported by our analysis of the \$US/\$A market – that exchange rates do not do this. There are three possible interpretations of this evidence. Firstly, there could be a time-varying risk premium which investors demand to hold, say, \$A nominal assets. Secondly, the market could be inefficient. Thirdly, there could be a 'peso problem'. We examine each of these interpretations in turn. For convenience, we repeat the arbitrage conditions introduced at the beginning of the paper which should be satisfied by Australian interest rates:

$$i^A = i^{US} - E(\Delta S / S) + rp \tag{1}$$

or

$$r^A = r^{US} - E(\Delta S^R / S^R) + rp^R. \tag{2}$$

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nominal growth rate of GDP (which, at least for \$A-denominated debt, it does). Therefore, **at present**, the real economy is on an unsustainable debt/GDP path.

[29] This should be true for large enough transactions. Goodhart and Taylor (1987) provide detailed estimates of transactions costs in the London and Chicago futures markets.

[30] To be precise, the three conditions imply that the forward rate must be an unbiased predictor of the future spot rate or, equivalently, that $s_{t+k} - f_{t,k}$ must be a martingale.

A time-varying risk premium

Both our evidence, and the more extensive survey evidence of Thorpe et. al. (1988) suggest that there was a statistically significant gap between the forward discount and market expectations of depreciation – and hence a risk premium – from late 84 to late 85. Since then, the survey data leads us both to accept the null hypothesis of zero risk premium. Section III of our paper presents two *a priori* calculations of the average magnitude of the risk premium necessary to induce a US consumer-investor to hold a small part of her wealth in Australian nominal assets. Within the chosen framework the calculations are as realistic as we could make them, but in some respects they are pretty naive.[31] Be that as it may, these theoretically-based estimates of the average risk premium are so small as to be negligible (compared, for example, with the average real interest differentials between Australia and the US derived in section IV). These calculations are consistent with the fairly well established inability of current models of time-varying risk premia to account for interest rate differences between countries with essentially no impediments to the movement of capital (e.g., Hodrick, 1987; Cumby, 1988).

An inefficient foreign exchange market

Consider the Dornbusch (1976) model of a small open economy with perfect capital mobility, rational market participants, sticky goods prices and no risk premia.[32] In this model, an unanticipated increase in domestic nominal and real interest rates (i.e., a tightening of monetary policy in the small economy) leads to a market anticipation that in the long-run, domestic inflation will be lower than previously expected. As a direct consequence, in the long-run, the domestic nominal

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[31] Investors are assumed to be homogeneous with a constant relative risk-aversion and to do their intertemporal optimization one period at a time.

[32] This seminal model forms the basis of most modern open-economy macro-models including, in the Australian context, the Murphy (1988) model and the MSG2 (McKibbin and Elliott, 1989 and McKibbin and Sachs, 1989) model.

exchange rate will be higher than previously expected, while the long-run real exchange rate will be unchanged. Immediately the tightening is recognised by the market, the domestic exchange rate jumps up – overshooting its long-run nominal appreciation. This jump is necessary so that, **during adjustment to the long-run, the exchange rate depreciates** (in both real and nominal terms) at exactly the rate necessary to equate the return on domestic and world short-term nominal assets. If there are repeated shocks, they cause repeated jumps in the nominal exchange rate, but during periods in which there is no relevant new information, the return on short-term domestic nominal assets is the same as it is on foreign ones.

Unfortunately, the world does not seem to work like this model. Apparently unanticipated changes in domestic interest rates do not lead to jumps in the exchange rate (Goodhart, 1988). Tight domestic monetary policy does not lead to an adjustment path for the exchange rate like that described in the previous paragraph.[33] It is an oft repeated claim – and it seems to be true – that the domestic real (and nominal) exchange rate of an open economy tends to be held up **during** periods when the domestic real interest rate is higher than the rest of the world.[34] What is rarely remarked is that this claim strongly suggests that the foreign exchange market is inefficient. This follows because, provided risk premia are small (which, at least since late 85, all our evidence suggests they are), the

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[33] In discussing the fact that the long-term real interest differential between the US and its trading partners increased by about 5 percentage points from 1980 to mid-1984, and the real appreciation of the $US from 1980 to 1985, Dornbusch and Frankel (1987) comment: [the overshooting model implies that] "the entire increase in ... the value of the dollar should have occurred in one (or two or three) big jumps, for example when it was discovered that monetary policy was going to be tighter than previously expected, or fiscal policy looser. Yet the appreciation in fact took place month-by-month, over four years (with investor expectations, as reflected in the forward discount, interest differential or survey data, all forecasting a depreciation)."

[34] As an example of this it is clear from Table 4 that there has been a substantial return from holding short-term $A nominal assets during the extended period of high real interest rates on these assets since late 84 (or late 85).

expectation that the exchange rate will be held up combined with high domestic real interest rates should lead to a flood of foreign demands for domestic interest-bearing assets – setting off the adjustment process described in the previous paragraph. In the real world, instead of a flood, there is a trickle, which, on average, slowly appreciates the domestic currency.[35] **As a consequence, for an extended period (many months) there is the opportunity for substantial gain (with what seems an associated small risk) but this opportunity is seized by comparatively few.**

These observations can be rationalized by the evidence of Froot and Frankel (1989) that survey expectations follow the forward rate. Because of covered interest parity, when domestic nominal interest rates are higher than world nominal interest rates, the forward rate predicts depreciation of the domestic currency (at a rate which, if it were realised, would equalise the yield from domestic and foreign assets). For market participants who use the forward rate as an 'anchor' for their expectations of the future spot rate,[36] the gain from higher domestic nominal interest rates is completely offset by an expected depreciation. This may explain why there is no massive foreign demand for domestic nominal assets when domestic interest rates are high. **But if that is so, what remains is the substantial**

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[35] Given covered interest parity, regressions of equation (3) support this statement. Both the regressions in Table 1, as well as most of those quoted by Goodhart (1988), Obstfeld (1988), Thorpe et. al. (1988) and many others, find negative estimates of β – although they are usually insignificantly different from zero. These regressions suggest that a higher domestic nominal interest rate leads, on average over the next month or three months, to a higher (or perhaps unchanged) domestic nominal exchange rate than would otherwise have been the case. See also Meese and Rogoff (1988).

[36] "In many [uncertain] situations, people make estimates by starting from an initial value [the anchor] that is adjusted to yield the final answer. The initial value ... may be suggested by the formulation of the problem, or it may be the result of a partial computation. In either case, adjustments are typically insufficient. That is, different starting points yield different estimates, which are biased toward the initial values." Tversky and Kahneman (1974). Of course, expectations formed by such a process are not rational in an economist's sense.

puzzle that the actual behaviour of exchange rates does not alter market expectations.

A peso problem

In regions renowned for earthquakes, many people invest in earthquake insurance. Consider econometricians studying this behaviour during a time in which there have been no earthquakes. They may falsely conclude that the behaviour is irrational, because, over their sample, there has been no return from the investment. A similar difficulty exists in the foreign exchange market. In Mexico in the 1970's, the peso was permanently at a forward discount compared to the $US despite a fixed exchange rate between Mexico and the US which had been in place for years (Krasker, 1980). The market continually expected a devaluation of the peso, and while it did not occur a 'peso problem' was said to exist. This term has now become standard to describe this small-sample problem (Hodrick, 1987). A peso problem is often invoked in defence of the hypothesis of rational market expectations, but our comments in this sub-section also apply if the market's expectations are not rational.

We have some evidence that the $A does suffer from a peso problem. Firstly, our survey expectations data from Mar 85 (Nov 85) to Sept 87 imply that, over this period, market participants expected a real depreciation of the $A at an average rate of 2.4% p.a. (4% p.a.) while the actual rate appreciated. Secondly, since late 84 (late 85), Australian ex post short-term real interest rates have been at an average premium of 2.4% p.a. (2.8% p.a.) compared to the US.

By themselves, these observations do not constitute overwhelming evidence that the $A suffers from a peso problem. They could be explained by the argument presented at the end of the last sub-section.[37] What makes a peso problem for Australia a distinct possibility are two further observations. Firstly, our skewness analysis demonstrates that, unlike all the other currencies we examined, over one week and four weeks the $A is subject to infrequent, unpredictable and large depreciations. So it seems reasonable that the market should have such events built into their expectations. Secondly, there is an obvious candidate for the cause of a peso problem – that in the longer run the real economy must adjust to put us on a sustainable net external debt/ GDP path – with a lower real (and nominal) exchange rate during the adjustment process. Examining what appear to have been the causes of the ten largest weekly depreciations of the $A since Jan 86 (Table 10), seven appear interpretable in terms of events in the Australian economy (all but the first, seventh and ninth falls in the Table). Of these seven, five (the second, third, sixth, eighth and tenth) appear related to the need for the real economy to adjust to put us on a sustainable net external debt/ GDP path. This evidence supports the view that if the $A suffers from a peso problem, its cause is a market perception of the need for a lower real exchange rate to put the economy on a sustainable net external debt/ GDP path.[38]

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[37] With reference to that argument, the average magnitude of depreciation predicted by the Australian market participants is close to that predicted by the forward rate (see Figure 2 and associated discussion). Note however that we find that the inflation differential dominates the forward rate as an explanator of market expectations (equation 6).

[38] An alternative candidate for the cause of a peso problem is a market perception that there is a non-negligible chance that Australian inflation will dramatically accelerate in the future and depreciate the nominal exchange rate. A belief in the possibility of either accelerated monetary expansion or the collapse of the Prices and Incomes Accord could be the source of this expectation. With the benefit of hindsight, although Australian inflation over the past four years has been substantially higher than comparable countries (see Figure 6) it has shown few signs of acceleration. If inflationary expectations were the source of a peso problem, agents should have been continuously surprised that the event(s) they were anticipating did not eventuate, and presumably have revised their expectations. Only one of the events in Table 10 seems related to inflation (the CPI announcement).

We briefly deal with two questions suggested by this analysis. Firstly: if there is a widespread and long-held expectation that the real value of the $A will fall, why doesn't it? The answer is probably that high short-term real interest rates have held it above the level it would otherwise have been.[39]

Secondly: if a peso problem exists, what will solve it? From the end of May to the end of July 86, the $A fell 15% against the $US and 19% against the TWI. The survey of market participants (Figure 2) for the ten Fridays immediately following this depreciation show that, on average, they expected a depreciation of the $A of 1.03% over the next four weeks; equivalent to depreciation at an annual rate of 12.6%. This number is again substantially higher than the depreciation predicted by the inflation differential, 7.2% p.a.[40], but closer to the annualized depreciation predicted by the 1 month forward rate over these ten weeks (10.8% p.a.).[41] This suggests that even a reasonably large depreciation may not be enough to influence expectations sufficiently to eliminate a peso problem. In principle, there should be a depreciation large enough to turn expectations around – but it is hard to know how large is sufficient. **An alternative possibility is that the peso problem will not be eliminated until the real economy is clearly seen to be moving onto a sustainable path for net external debt / GDP.**

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[39] Nevertheless, we remain impressed by the conclusions of Meese and Rogoff (1983, 1988) that macroeconomic fundamentals (and interest rate differentials in particular) have little significant capacity to explain movements in the nominal or real exchange rate, even over periods as long as a year.

[40] Derived using annual inflation rates which had been published when the expectations were formed – as for Figure 2.

[41] Interestingly, market participants' reaction to the previous rapid fall in the $A had been quite different. From the beginning of Feb to 8 Mar 85, the $A fell 16% against the $US and 13% against the TWI. Over the next ten Fridays, they predict an average **appreciation** of the $A at an annual rate of 8.8%. Perhaps the subsequent behaviour of the $A (and the terms of trade and current account deficit) changed their attitude.

Consequences for external imbalance

Finally, we comment on the relevance of these observations for the current debate on Australia's external imbalance. For an extended period since late 85, market participants have expected the real value of the \$A to fall against the \$US at roughly 4% p.a. This may explain why a substantial real interest premium on short-term \$A denominated assets can persist without setting off an adjustment of the Dornbusch (1976) type. If the market perceives significant real depreciation is necessary to put Australia onto a sustainable debt path, these consequences follow. Firstly, while the monetary authorities keep short-term interest rates high, the Australian economy pays a real interest premium on the substantial proportion of short-term \$A-denominated external debt. Secondly, when the monetary authorities reduce Australian short-term nominal interest rates, at some unpredictable time there may be a big fall in demand for Australian nominal assets and a large rapid depreciation. Presumably after a sufficiently large depreciation the market will cease to expect further real depreciation – but the depreciation required to change market expectations may be very large indeed. The hope is that such an abrupt exchange rate adjustment does not have serious adverse consequences for the wider economy.

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DATA APPENDIX

Exchange rate data

Dataset A: Weekly (Friday) exchange rates, 5 Jan 79 to 21 Apr 89.

The spot rate, one month forward rate, and three month forward rate for the \$A/\$US market and the Australian trade-weighted index (TWI) are from I. P. Sharp Associates 'Australian Financial Markets Data Base'. The first three of these series are, in turn, from the Commonwealth Bank of Australia, and are the average of buy and sell rates at the close of trade on each Friday. The TWI is from the Reserve Bank of Australia (RBA) at 4 p.m. The spot rates \$US/¥, \$US/£, \$US/DM, \$US/\$C are from I. P. Sharp Associates 'Currency Exchange Rates Data Base', and are the noon buying price in \$US in New York. These data are from the Federal Reserve System, N.Y. Bank.

Dataset B: Daily exchange rates, 1 Jan 86 to 11 Apr 89.

Rates are the daily representative rates from the RBA for \$US/\$A, ¥/\$A, £/\$A, DM/\$A, and the TWI.

For both datasets, cross-rates are derived by dividing the appropriate rates (e.g., for dataset B, ¥/£ is derived as [¥/\$A] / [£/\$A]).

Survey data on exchange rate expectations

"An attempt was made to contact the same individual each week, however if the usual respondent was unavailable then [foreign exchange] expectations would be elicited from an alternative forecaster. This method of survey guaranteed a high response rate." (Hunt, 1987). The sixteen companies in the survey were: A.N.Z. Bank, B.N.P., The Australian Bank, Barclays, B.A., B.T., Citicorp, Commonwealth Bank, Elders, Lloyds, Macquarie Bank, National Australia Bank, Rural and Industry Bank, Schroders, State Bank of N.S.W. and Westpac. Each Friday, we have the lowest, the highest, and the arithmetic mean expectation of the sixteen participants. The expectations data runs from Friday 8 Mar 85 to Friday 18 Sept 87, with five missing weeks: leaving 128 weeks of data. The regressions and figures in this section require spot and forward rates for the

\$US/\$A. We use the Friday close spot rate in the wholesale market, and the Friday one month forward average of buy and sell rates (quoted in the *Australian Financial Review* on the following Monday).

The data for section IV comes from these sources: 3 month Treasury bill interest rates and 10 year bond rates for the US and Australia as well as Australian CPI, net external debt and terms of trade are from RBA Bulletins (various issues). Other short-term interest rates and foreign CPI are from IMF International Financial Statistics (IFS). To evaluate ex post 3 month real interest rates, we used the average yield on Australian 3 month Treasury bills for the last tender in each quarter (from RBA Bulletin), and the quoted yield on 3 month US Treasury bills on the last trading day of the quarter (from the New York Times) along with realised CPI inflation over the next 3 months.

APPENDIX A

Here, technical details concerning the estimation of equation (3) are discussed. The results in Table 1 with k = four weeks use the one month forward rate to generate the variable $fd_{t,k}$, while the results with k = thirteen weeks use the three month forward rate. The fact that the forward rates are defined for a slightly different time length than the change in the spot rate makes minimal difference for our purposes. For example, for the first regression in Table 1, it amounts to ignoring the difference between s_{t+28} and s_{t+30} (with t measured in days). Viewed from time t, $s_{t+30} - s_{t+28}$ is very closely modelled as a random variable with zero mean, and so may be included in the error term in equation (3). This timing issue is, however, critical for alternative tests of the efficiency of the foreign exchange market (see Tease, 1988).

Recursive least squares regression of equation (3) with k = four weeks using data beginning on 6 Jan 84 shows that the point estimate of the coefficient β is strongly positive (as large as 17) and unstable up to the beginning of 1985, after which it becomes negative and fairly stable[42] to the end of the sample (21 Apr 89). Therefore in Table 1, we report regressions starting in Jan 84 and in Feb 85. The latter date is chosen to correspond as closely as possible to Tease (1988).

Using data on the exchange rates of five countries against the US and assuming H_0: $\alpha = 0$, $\beta = 1$, Cumby and Obstfeld (1984) strongly reject the assumption of conditional homoscedasticity for the errors in equation (3) for four of the exchange rates. Applying their test to our problems leads to test statistics of 2.02 and 2.93 for k = four weeks and 1.73 and 0.897 for k = thirteen weeks. The test statistic is asymptotically distributed $\chi^2(2)$ which has a critical value of 5.99 at the 5% level. Thus, with this test, we cannot reject the null hypothesis of conditional homoscedasticity in all cases.

APPENDIX B

This appendix examines the statistical properties of the exchange rates in Datasets A and B, including the conditional variance, covariance and skewness of the exchange rates. We provide a summary of our findings – further details are given in Smith (1989b).

For all the exchange rates in the two datasets we establish the following results. Using Perron and Phillips (1987) tests, we accept the null hypothesis that the log of the spot exchange rate at a one week interval (Dataset A) and one day interval (Dataset B) has no significant time trend but requires, at least, one unit difference to be stationary. Using Augmented Dickey-Fuller (1979) tests, we reject the hypothesis that the log exchange rate requires a second unit difference to be

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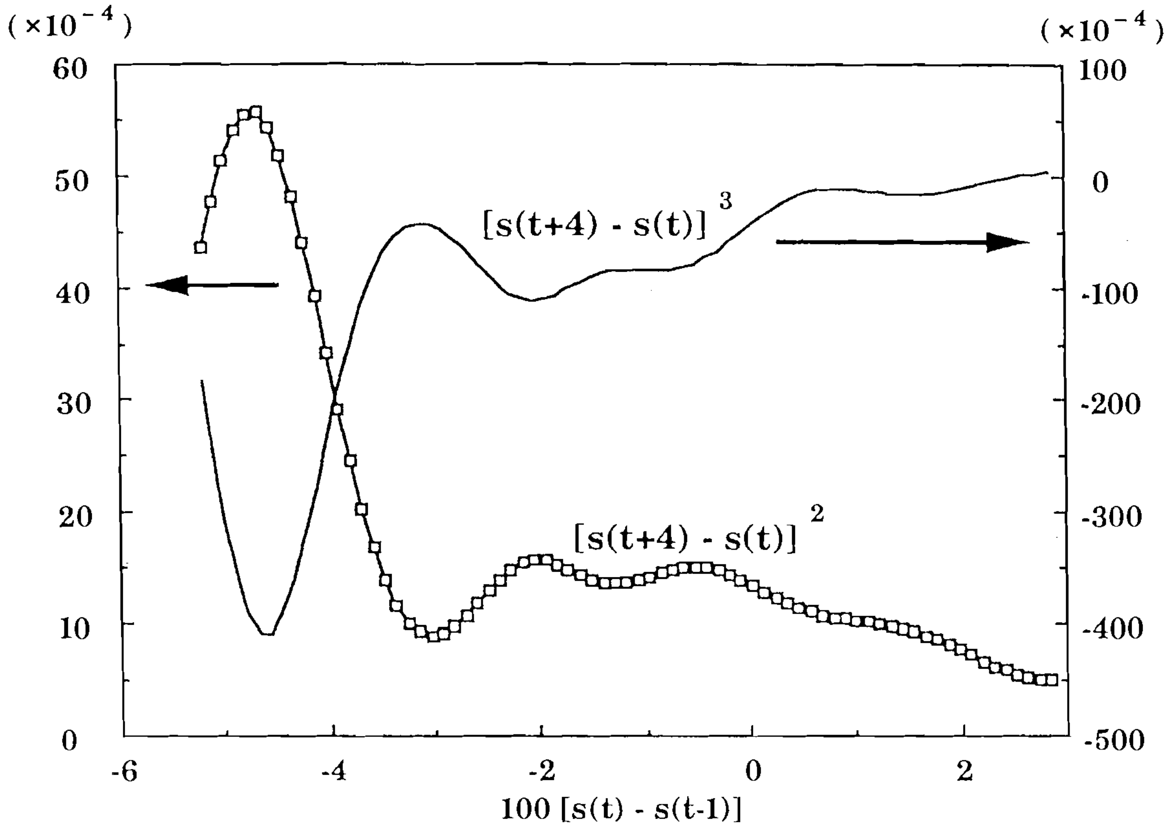
[42] Despite this fair degree of stability, estimates of equation (3) for sub-periods can produce rather different results – see equation (3) in Table 2.

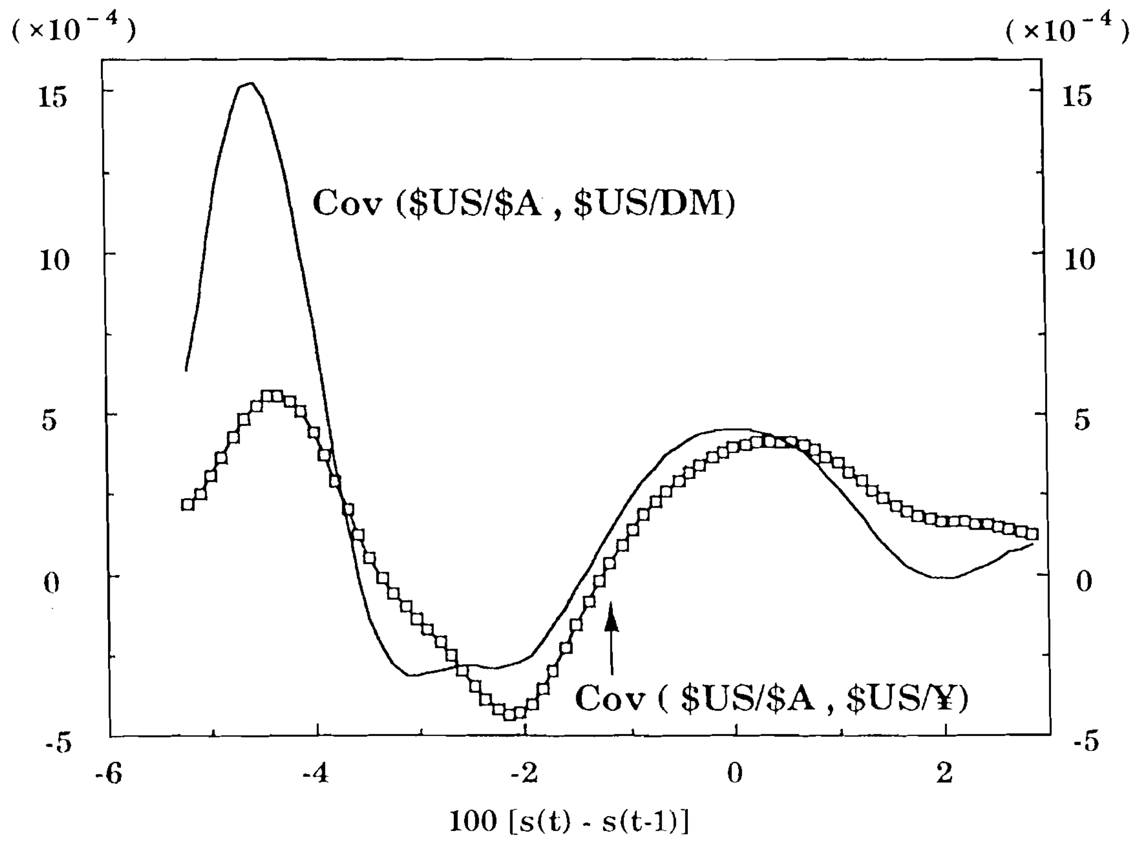
stationary. Thus, over the sample periods, these tests do not reject the hypothesis that each exchange rate follows a random walk with no drift.

The conditional variance of log changes in the spot exchange rate is successively modelled as autoregressive conditional heteroscedasticity [ARCH] (Engle, 1982), generalised ARCH [GARCH] (Bollerslev, 1986) and exponential GARCH [EGARCH] (Nelson, 1988). An ARCH(5) model is fitted to both datasets, and evidence for ARCH is found – although the explanatory power of the models is low. A GARCH(1,1) model is then fitted to both datasets, and the likelihood function shows this model to be far superior to the ARCH model. These two models impose a symmetrical distribution for the estimated conditional variance, while the EGARCH model does not. Given the significant skewness reported in section V, we expected to find evidence of EGARCH. In almost all cases, the parameter point estimates in our EGARCH model imply that there is greater volatility in the immediate aftermath of a fall in the $A than in the immediate aftermath of a rise. Unfortunately, the standard errors of the estimates are so large that this asymmetry is not statistically significant.

Rather than quote all the parameter estimates for each of the models, Figure 10 shows non-parametric estimates of the variance, skewness and two covariances for four-weekly changes in s[$US/$A], conditional on the change in s[$US/$A] over the previous week. We established that conditioning on the change in s[$US/$A] over the previous week provided more variability in the estimates than conditioning on the change in s[$US/$A] over the previous four weeks. Define $s_t - s_{t-1} = g(t)$, and imagine a variable G(t+4,t) defined in terms of exchange rates at times t and t+4 (e.g., $G(t+4,t) = [s_{t+4} - s_t]^2$ or $G(t+4,t) = [s_{t+4} - s_t]^3$). With a sample s_t, t = 1,...,N, the non-parametric estimator of $G(\tau+4,\tau)$ conditional on $g(\tau) = g^*$ is

Figure 10
Conditional second and third moments and covariances of four-week changes in s[$US/$A] using Dataset A, Jan 84 to Apr 89





$$G(\tau+4,\tau \mid g(\tau)=g^*) = \sum_{t=2}^{N-4} G(t+4,t)\,.\,\omega_{t\,g^*}$$

where
$$\omega_{t\,g^*} = \frac{\exp\left(-\frac{1}{2}\left[\frac{g(t)-g^*}{h}\right]^2\right)}{\sum_{t=2}^{N-4}\exp\left(-\frac{1}{2}\left[\frac{g(t)-g^*}{h}\right]^2\right)},$$

and $h = \sigma_g.\ (N-5)^{-1/5}$ and σ_g is the standard error of the observations $g(t)$, $t = 2,\ldots,N-4$. This non-parametric estimator is very similar to one suggested by Pagan and Schwert (1989).[43] Figure 10 shows conditional estimates of $(s_{t+4}[\$US/\$A] - s_t[\$US/\$A])^2$, $(s_{t+4}[\$US/\$A] - s_t[\$US/\$A])^3$, $(s_{t+4}[\$US/\$A] - s_t[\$US/\$A]).(s_{t+4}[\$US/¥] - s_t[\$US/¥])$ and $(s_{t+4}[\$US/\$A] - s_t[\$US/\$A]).(s_{t+4}[\$US/DM] - s_t[\$US/DM])$ based on Dataset A from Jan, 84 to Apr, 89. The third and fourth of these estimators correspond to covariances provided the expected change in the exchange rate over four weeks is zero, which is a good approximation. The unconditional sample estimates of the four variables above are, respectively, 12.63×10^{-4}, -47.91×10^{-4}, 2.30×10^{-4} and 2.46×10^{-4}.

Within the sample, Figure 10 shows that the behaviour of the exchange rate over the next four weeks depends to a considerable extent on its movement in the previous week.[44] Based on this figure, the risk premium required by a US investor after a fall of the \$A of between 4% and 5% in the previous week should be substantial (using equation (10) and assuming a portfolio share in Australia of

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[43] To deal with the problem of outliers in the $g(\tau)$ data, Pagan and Schwert (1989) suggest a slightly different estimate. We deal with this problem in a different way: after ordering the $g(\tau)$ sample from the most negative to the most positive, we only estimate $G(\tau+4,\tau)$ for g^* values between the fifth smallest value of $g(\tau)$ and the fifth largest value of $g(\tau)$.

[44] However, the results of Pagan and Schwert (1989) suggest that, out of sample, the predictive capacity implied by the Figure is probably substantially overstated.

0.02 as well as 0.10 in other foreign countries in the proportions used in the text, the risk premium is 0.5% p.a.). But, both this analysis and the analysis summarized above based on parametric approaches (ARCH, GARCH and EGARCH) suggest that the conditional moments of the change in the exchange rate over the next four weeks only depend on the exchange rate over the previous few weeks (probably no more than three weeks). As a consequence, these results do not undermine the conclusions reached in section III of the text. On average over extended periods (several weeks or several months), the risk premium which a utility maximizing US consumer-investor should demand for holding short-term nominal Australian assets seems to be negligible – compared, for example, to the average short-term real interest differential between Australia and the US from late 84.

APPENDIX C

This appendix describes how the estimates in Table 3 were derived. We measure time in weeks and use exchange rate Dataset A from 6 Jan 84 to 21 Apr 89. We do not have interest-rate data for each country, so we use the approximation $(1 + i^j) \approx 1$, for all j. Then,

$$z^j \equiv \rho^j - \rho^{us} \equiv \frac{P_t^{us}}{P_{t+4}^{us}}[(1+i^j)(1+\frac{\Delta S_t^j}{S_t^j}) - (1+i^{us})] \approx \frac{P_t^{us}}{P_{t+4}^{us}} \cdot \frac{\Delta S_t^j}{S_t^j},$$

where $\Delta S^j{}_t \equiv S_{t+4}[j\,/\,\$US] - S_t[j\,/\,\$US]$. Since the period of analysis is four weeks, the approximation, $(1 + i^j) \approx 1$, introduces an average error of the order of (or less than) 1%. Further, as noted previously, almost all the variation in z^j arises from exchange rate variation.

Price data comes from monthly US CPI data from OECD Main Economic Indicators (various issues). All Fridays in any given month are assigned the price index for that month. To evaluate cov (z^A , ρ^{us}), we require $E_t[P^{us}{}_t / P^{us}{}_{t+4}]$. We assume a simple form of adaptive expectations: $E_t [P^{us}_t / P^{us}_{t+4}] = P^{us}_{t-16} / P^{us}_{t-12}$. The sixteen week lag is used to ensure that only published price indices are used in forming the expectation. A more sophisticated expectation formation assumption[45] should presumably reduce $P^{us}_t / P^{us}_{t+4} - E_t [P^{us}_t / P^{us}_{t+4}]$, and hence reduce our estimate of cov (z^A , ρ^{us}). It suits our purposes if our estimate is an overestimate. To evaluate the other covariances and the variance, we assume

$$E_t [(P^{us}_t / P^{us}_{t+4}).(\Delta S^j_t / S^j_t)] = 0, \qquad (C.1)$$

because, over four weeks, exchange rates changes for the currencies we consider are well approximated as an unpredictable random variable with zero mean. We have established that using the one month forward discount in equation (C.1) as the expected depreciation of the \$A against the \$US makes only a small change to our estimate of var(z^A) – it increases the estimate from 11.9×10^{-4} to 12.5×10^{-4}.

APPENDIX D

We describe here a non-parametric statistical test for the skewness of the distribution D_i [a/b]. Assume we have a sample with an odd number (2n + 1) of independent[46] observations from D_i [a/b].[47] Order the sample from the most negative to the most positive and define d_j as the j^{th} observation (so $d_{j-1} \le d_j \le d_{j+1}$ for $1 < j < 2n + 1$). d_{n+1} is the median of the sample. Define $y_j = d_j - d_{n+1}$, $j = 1,\ldots, 2n+1$, and form the random variables Y_k , $k = 1,\ldots,2n$, defined by

$Y_k = -1$ when y_j is the k^{th} largest of the y_j 's **in absolute value** and y_j is negative;

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[45] Given the dramatic change in the income velocity of money in the 1980s (see, for example, Friedman, 1988 – especially Figure 2), *ex ante* it might have been quite difficult to have had more accurate inflation expectations than the backward looking ones used here.

[46] The assumption of independence makes the analysis exact. We examine the removal of this assumption at the end of this appendix.

[47] If we have an even number (2n + 2) of independent observations, we define d_j as described, but now define $y_j = d_j - (d_{n+1} + d_{n+2}) / 2$, $j = 1,\ldots, 2n+2$. The random variables Y_k , $k = 1,\ldots,2n$, are defined as described and equation (D.1) is again the basis of our non-parametric test for the skewness of D_i [a/b].

$Y_k = 1$ when y_j is the k^{th} largest of the y_j 's **in absolute value** and y_j is non-negative. Finally, define the random walk Z_k, by

$$Z_0 = 0, \qquad Z_k = \sum_{j=1}^{k} Y_j , \quad k = 1,\dots,2n.$$

Provided that $d_j \neq d_{n+1}$ for all $j \neq n+1$,[48] there are exactly n '−1' values and n '+1' values taken by the random variables Y_k, $k = 1,\dots,2n$, and hence the random walk, Z_k, walks from $Z_0 = 0$ to $Z_{2n} = 0$. Crucially, under the null hypothesis that the distribution D_i [a/b] is symmetric, all distributions of the n '−1' values and n '+1' values among the random variables Y_k, $k = 1,\dots,2n$, are equally likely and all walks Z_k from $Z_0 = 0$ to $Z_{2n} = 0$ are also equally likely. By contrast, if D_i [a/b] is negatively (positively) skewed, Z_k, $k = 1,\dots,2n$ will be more likely to walk to large negative (positive) values before returning to zero when $k = 2n$. No specific assumption about the distribution of D_i [a/b] is necessary – the null hypothesis is simply that D_i [a/b] is symmetric.

Define the random variable W_M as the number of the random variables Y_k, $k = 1,\dots,M$, which take the value '−1'. Under H_0, $Pr(W_M = w)$, is

$$\Pr (W_M = w) = \frac{\binom{M}{w} \cdot \binom{2n-M}{n-w}}{\binom{2n}{n}} . \tag{D.1}$$

Our one-sided test for negative [positive] skewness involves evaluating the probability, $Pr(W_M \geq w)$ [$Pr(W_M \leq w)$]. For the results in Table 9, $n = 84$, and $M = 10$ was chosen. Evaluation of (D.1) gives: $Pr(W_{10} = 0) = Pr(W_{10} = 10) = 0.0007$, $Pr(W_{10} \leq 1) = Pr(W_{10} \geq 9) = 0.0090$, $Pr(W_{10} \leq 2) = Pr(W_{10} \geq 8) = 0.0494$. Thus, sample values of W_{10} of 9 or 10 (0 or 1) imply rejection of H_0 at the 1% level of significance

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[48] With the exception of the TWI data (which is quoted to three figures), all our exchange rate data is quoted to (at least) four significant figures, so it is unlikely that any two values of d_j would be the same.

against the alternative of negative (positive) skewness, while a value of 8 (2) implies rejection at the 5% level. Sample values $3 \leq W_{10} \leq 7$ are insignificant.[49]

As discussed in Appendix B, in general the distribution of D_i [a/b] at time t (D^t_i [a/b]) depends on observations of $s_{\tau+i} - s_\tau$, $\tau < t$. Clearly, this invalidates our assumption of the independence of the observations, and our test of skewness must be modified. The null hypothesis is now that each of the D^t_i [a/b] distributions is symmetric with a common mean, μ. Under this null, the distribution of W_{10} depends on how different are the distributions D^t_i [a/b], t = 1,..., 2n+1. At one extreme is the case already examined when all the distributions are identical, and each Y_k, k = 1,...,2n has an equal chance of coming from any of the D^t_i [a/b], t = 1,...,2n+1. At another extreme, assume that under the null there are only two distinct (symmetrical) distributions: for ten particular times, t(j), j = 1,...,10, the distributions $D^{t(j)}_i$ [a/b] $\equiv D^+$, and at all other times, τ, $\tau \neq$ t(j), j = 1,...,10, D^τ_i [a/b] $\equiv D^*$. D^* is assumed to have all its probability weight "near" μ, while D^+ is assumed to have all its probability weight in two tails "far from" μ, so that D^* and D^+ have no overlap. In this contrived case, we can be sure that for j = 1,...,10, Y_j must come from D^+ and hence from the ten particular times, t(j), j = 1,...,10. Then under the null hypothesis, $\Pr(W_{10} = w)$ is simply

$$\Pr(W_{10} = w) = \binom{10}{w} / 2^{10}. \quad \text{(D.2)}$$

Equation (D.2) gives: $\Pr(W_{10} = 0) = \Pr(W_{10} = 10) = 0.00098$, $\Pr(W_{10} \leq 1) = \Pr(W_{10} \geq 9) = 0.011$, $\Pr(W_{10} \leq 2) = \Pr(W_{10} \geq 8) = 0.055$. Thus, even in this extreme case, the critical values of W_{10} are only changed slightly.

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[49] An alternative test based on the distribution of the maximum (or minimum) value taken by the walk, Z_k , k = 1,...,2n, was also examined but found to have little power. One of us (J. S.) examined the skewness of the data assuming that D_5 [a/b] has a distribution of the stable Paretian form. The results are similar to those reported here.