

**MONTHLY MOVEMENTS IN THE AUSTRALIAN DOLLAR
AND REAL SHORT-TERM INTEREST DIFFERENTIALS:
AN APPLICATION OF THE KALMAN FILTER**

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ABSTRACT

This paper applies a rational expectations model of the real exchange rate to Australian data. Specifically, it decomposes monthly movements in Australia's real exchange rate into a transitory and a permanent component. The transitory component is identified with changes in the unobservable *ex ante* short-term real interest differential. The permanent component is denoted as changes in the unobservable long-run equilibrium real exchange rate. A state space model provides the framework for the treatment of these unobservable components and the traditional assumptions of the expectations hypothesis of the term structure of interest rates and no cross-currency risk premium are relaxed.

The *ex ante* real interest differential is found to explain very little of the month-to-month movement in the real exchange rate. However, given that the Australian data fails to unambiguously support the existence of a risk premium in the foreign exchange market, the model collapses to an uncovered interest parity relation which finds little empirical support in the literature. These results imply that the model's assumption of rational expectations and hence, an efficient market in foreign exchange, may be inappropriate for describing the monthly variation in Australia's real exchange rate.

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1. INTRODUCTION

Despite an impressive volume of research, consensus on the forces responsible for the **observed** behaviour of real exchange rates is yet to be reached. Since the float of our dollar in December 1983, Australia's financial liberalization appears to have accentuated the role of financial markets in determining our exchange rate (Miller and Weller (1991)).

In particular, the relationship between the real exchange rate and real interest rates has fallen under scrutiny since the late 1980s. Over this period in Australia, real interest rates have generally been high relative to our major trading partners. At the same time, the Australian dollar has appreciated significantly in real terms. It would seem clear that the high domestic real interest rates contributed to the dollar's appreciation by attracting foreign capital. However, this link has not always been supported by observation in the shorter run. For example, domestic unofficial cash rates have been reduced by more than seven per cent since December 1989, substantially narrowing both nominal and real interest differentials. To date,¹ the strength of the dollar has persisted.

This paper investigates the role of interest differentials in determining short-term activity in the dollar. In one scenario, where interest differentials matter, the foreign exchange market is efficient² and therefore influenced, even in the short-term, by the behaviour of fundamentals.³ It

¹ June 1991.

² An efficient market is one in which the asset price in question fully reflects all available information and the forces of competition ensure the optimal allocation of scarce resources among the various investment opportunities. Such a market requires no government interference and profitable market players are "rational".

³ That is, variables such as the terms of trade, productivity, etc., which indicate the underlying strength of Australia's economy relative to the rest of the world. These variables are generally believed to dominate the real exchange rate over the *long run*.

therefore seems reasonable and tractable to approach this question from the theory of uncovered interest parity (UIP). UIP postulates that, given an efficient market in foreign exchange, risk-neutral traders and negligible transactions costs, the expected change in the real exchange rate reflects the expected real interest rate differential. UIP in its purest form has been rejected by empirical evidence.⁴ It has been suggested that its failure is due to the existence of a risk premium in the foreign exchange market.⁵ If we accept this as a reasonable proposition, then the theory of UIP can be retained. However, its underlying assumption that traders are risk-neutral must be relaxed to allow for the existence of a time-varying risk premium. This methodology was adopted by Campbell and Clarida (1987) to examine what proportion of the monthly movement in the real US dollar could be explained by real short-term interest differentials relative to a number of their trading partners. This paper estimates their rational expectations model with Australian data. Monthly statistics for Australia against the United States, Japan, West Germany, United Kingdom and a constructed trade weighted index⁶ are examined for the post-float period, extending from December 1983 to August 1990.

Monthly movement in the real exchange rate is simplified into two components. The first component is identified with *ex ante*⁷ real short-term interest differentials; the second with all remaining economic fundamentals⁸ as embodied in the expected long-run real exchange rate. Both of these explanatory variables are unobservable. Earlier studies⁹ have commonly assumed the long-run exchange rate constant, and adopted proxy measures for the *ex ante* short-term real interest differential.

⁴ See Hodrick (1987) for international evidence with a range of currencies, and Smith and Gruen (1989) for the Australian/US exchange rate.

⁵ Researchers continue to debate the existence of such a risk premium. Refer to Smith and Gruen (1989) for a detailed discussion.

⁶ See Appendix 1 for a detailed description of data methods and sources.

⁷ The *ex ante* real interest differential is the one-period ahead expectation of the realized or *ex post* real interest differential. This *ex post* real differential is calculated as the nominal one-month interest differential between two countries at time $t-1$, adjusted for their actual inflation differential at time t . *Ex ante/ex post* analysis examines any divergence between expected and realized values of this differential.

⁸ These fundamentals include measurable variables such as the terms of trade and productivity shocks, and unmeasurable variables such as changing consumer preferences.

⁹ Shafer and Loopesko (1983); Sachs (1985).

Campbell and Clarida's (1987) model avoids such compromises by using a more flexible estimation technique able to accommodate models which include unobservable variables - the Kalman filtering prediction error technique. The Kalman filter is generally used to evaluate difficult likelihood functions. In this paper, however, it is used primarily to make variance decomposition statements about the unobservable components of the model.

The remainder of the paper is organised into five parts. Section 2 briefly discusses different schools of thought which have developed to explain the observed behaviour of the real exchange rate. Section 3 develops Campbell and Clarida's (1987) structural model from the principles of uncovered interest parity. Section 4 discusses the Kalman filter approach to estimation which requires this structural model to be cast in state space form. Variance decomposition techniques are primarily used to investigate the stochastic¹⁰ relationships between the real exchange rate and its unobservable components. Section 5 presents empirical results for the Australian data. We find that since the float of the Australian dollar in December 1983, *ex ante* real interest differentials have not accounted for the greater proportion of monthly variation in the real exchange rate. Therefore, given the specification of this model, movements in the dollar's real exchange rate have been dominated by unanticipated shifts in the expected long-run real exchange rate. Section 6 discusses the implications and limitations of Campbell and Clarida's (1987) methodology and identifies the direction in which we feel future research into the shorter-run dynamics of the Australian dollar should advance.

2. APPROACHES TO THE DETERMINATION OF SHORT-TERM MOVEMENTS IN THE REAL EXCHANGE RATE

If we think of foreign currency as an asset price, then an efficient foreign exchange market would ensure that the observed variation in the Australian dollar reflected the behaviour of fundamental economic variables. We would expect that the stance of Australia's monetary policy relative to that of our trading partners, as embodied in real short-term

¹⁰ As per Campbell and Clarida (1987), our approach assumes that the stochastic processes governing real interest rates are stable through time.

interest rate differentials, would be reflected in this price. This paper focuses on the role of the real short-term interest differential in determining the **monthly** variation in Australia's real exchange rate.

Uncovered interest parity theory (UIP) provides a convenient building block from which an investigation of the relationship between real interest and exchange rates may be developed. As mentioned, UIP is based on three theoretical assumptions. Firstly, the market in foreign exchange is efficient (such that market expectations are rational). Secondly, any transactions costs incurred by market participants are negligible, and finally, the actions of these participants are not affected by any risk considerations. While little empirical evidence can be found to support UIP, this may be the result of the inappropriateness of one or more of the theory's underlying assumptions:

If we accept that the first premise, that of an efficient market in foreign exchange, is appropriate, then the failure of UIP may be due to the fact that traders are not risk-neutral. Risk premia which vary over time may exist in and distort the market. Otherwise, transactions costs, although usually maintained to be too small, may effect exchange rate transactions. Baldwin (1990) proposes that such costs occasion "bands of inaction" in the price of foreign exchange.¹¹

Alternatively, market participants may hold an ongoing belief that an infrequently occurring event relevant to the determination of the exchange rate is imminent. The low probability of this event being captured in sample is the so-called "peso" problem. Market participants may incorrectly be labelled "irrational" if the sample period doesn't contain the infrequent event that is driving their behaviour.

Within this **efficient** market framework, the literature is divided as to the mechanism by which the observed real exchange rate is predominantly determined.¹² The empirical content of this paper concentrates on the

¹¹ That is, transactions costs accompanied with uncertainty, may result in interest differentials unmatched with any expected exchange rate change (allowing for the possibility of a risk premium). Within a small band then, interest differentials do not induce capital flows towards the country offering the highest return because the transactions costs make such movement of capital suboptimal.

¹² See Coughlin and Koedijk (1990) and Dornbusch (1989) for more comprehensive reviews of the different theoretical approaches to real exchange rate determination.

monetary approach which isolates international investment decisions channelling capital into those countries offering the highest real return on their assets (that is, the highest real rates of interest) as the dominant explainer of exchange rates in the shorter-term. A second "rational" school advocates the *real* approach, believing observed variation in the Australian dollar to reflect the market modifying its expectations of the equilibrium exchange rate to account for real factors (Coughlin and Koedijk (1990)). As the relative price of foreign goods in terms of domestic goods, the real exchange rate should be expected to reflect shocks to the supply (for example, changes in productivity, technology or the labour supply) and demand (for example, shifts in preferences) for each country's product (Ghosh (1990)). As a small open economy, Australia's terms-of-trade for example, have a substantial effect on our real exchange rate (refer Blundell-Wignall and Gregory (1989)). Meese and Rogoff (1988), investigating the case for the US, find little empirical evidence to support the existence of any stable relationship between real interest rates and real exchange rates. Instead, they propose that real disturbances may be a major source of exchange rate volatility. Such real variables have not typically been considered responsible for the observed **short-run** movements in real exchange rates.¹³ They are more commonly believed to explain the longer-term value of the dollar.

If, alternatively, we reject the first premise that the foreign exchange market is efficient, then two main theoretical explanations are offered for the generation of inefficiencies. One school of thought believes that the unexplained short-term changes in exchange rates can be accounted for by the existence of rational speculative bubbles in the foreign exchange market.¹⁴ These speculative bubbles occur when market expectations of the movement in the exchange rate are self-fulfilling. As a result, market

¹³ Explanations have more commonly depended on the Dornbusch (1976) overshooting model. However, Coughlin and Koedijk (1990) and Stockman (1987) mention the potential importance of random real shocks for moving the real exchange rate.

¹⁴ Refer, for example, to Flood (1987) for evidence of rational bubbles in the foreign exchange market. "Rational" bubbles occur when the market does indeed know the true model describing the exchange rate. That is, the market makes a "rational" response to irrational factors and whilst the path of the exchange rate may deviate from that consistent with fundamentals, rational arbitrage conditions are satisfied along this path (Miller and Weller (1991)).

expectations alone have a strong influence on the value of the dollar, causing inefficiency in the foreign exchange market, at least for a time. There is little explanation for the origin of such bubbles and their existence is underpinned by the belief that the market will ultimately return to fundamentals.

Another school proposes that the foreign exchange market is inefficient because expectations are not fully rational. That is, heterogeneous groups of market participants may drive the exchange rate away from its fundamental value either through irrational sentiments, investor misperception or "fads" (Miller and Weller (1991)). If the foreign exchange market is inefficient because market expectations are irrational, fundamental economic variables will not figure in the determination of short-run movements in the dollar. Therefore, interest differentials would not be expected to have any explanatory power. Theory in this area identifies two specific groups of "irrational" traders responsible for destabilizing exchange rate movements.

The first group, the *feedback* traders, base their demand for foreign exchange on historical returns rather than on any expectation of future fundamentals (Cutler et al. (1990)). Therefore, their forecasts over short time horizons (1 week to 3 months) employ extrapolative rules.¹⁵

The second group, the *noise* traders, irrationally perceive random fluctuations in the exchange rate as a viable source of information on which to make a profitable trade (Black (1986)). The very existence of these unpredictable traders introduces a risk in the price of foreign exchange that discourages rational arbitrageurs from the market. De Long et al. (1990) describe the process by which generally "bullish" noise traders can earn higher expected returns than sophisticated traders: despite the absence of any fundamental risk, the observed exchange rate diverges from its

¹⁵ Over long time horizons (1 or more years), feedback traders assume regressive expectations in order to make their forecasts. See Cutler et al. (1990) for a discussion of the role of feedback traders in increasing the variability of the exchange rate.

fundamental value as the risk generated by the presence of the noise traders depresses its price.¹⁶

Questions concerning the cost-benefit analysis of actively acquiring information about the equilibrium value of the exchange rate may also have implications for explaining inefficiency in this market. If participants perceive the foreign exchange market to be efficient (that is, they believe that the current real exchange rate is a price signal that reflects all the private and public information available in the market) then they may not believe that research into the equilibrium real exchange rate is worthwhile. These issues are beyond the scope of this paper and we recognise the technical difficulties associated with their empirical estimation.

This paper now undertakes the specific technical application of a rational expectations, efficient market model of the real Australian exchange rate as estimated by Campbell and Clarida for the United States. Based on the ambiguous results returned by this application, the alternative explanations for the short-run activity of the dollar, as canvassed in this section, must be seen as live possibilities.

¹⁶ Miller and Weller (1991) present a comprehensive survey of the literature in this area. They note that three assumptions are crucial to the results of the De Long et al. model:

- (i) the sophisticated investors, taking positions on the basis of fundamental indicators, have time horizons that are not "too long";
- (ii) the erroneous beliefs of the noise traders are positively correlated; and
- (iii) the expectations of noise traders are generally "bullish".

3. THE STRUCTURAL MODEL: SPECIFICATION

The model, as developed by Campbell and Clarida (1987), is based on uncovered interest parity:

$$q_t = E_t[q_{t+1}] + E_t[d_{t,t+1}] + \delta_t \quad [1]$$

where:

- q_t is the natural log of the real exchange rate quoted as the value in domestic currency of one unit of foreign currency. A real dollar appreciation corresponds to a fall in q_t .
- $E_t[q_{t+1}]$ is the expected value in period t , of the natural log of the real exchange rate in the next period, $t+1$.
- d_t is the *ex post* short-term real interest differential, the *ex post* one-period real interest rate realized on foreign assets in period t , less that realized on domestic assets held in period t .
- $E_t[d_{t,t+1}]$ ¹⁷ is the *ex ante* short-term real interest differential.

The traditional restriction of strict equality between the real short-term interest differential and the expected exchange rate change is relaxed, allowing for a time-varying risk premium, denoted δ_t .¹⁸ Two versions of the model are estimated. A linear risk premium model estimates the value of δ_t and a restricted form sets $\delta_t=0$. When δ_t is zero, equation [1] can be interpreted as a logarithmic approximation to uncovered interest parity.

In order to make inferences about the long-run relationships of these variables, equation [1] is solved forward. Iterative expectations¹⁹ (Samuelson (1965)) yields:

¹⁷ $E_t(d_{t,t+1})$ embodies future inflation expectations. That is, $d_{t,t+1}$ is the differential on assets held from t to $t+1$.

¹⁸ Other evidence (Frankel (1985), Frankel (1988), Smith and Gruen (1989)) finds the risk premium in the foreign exchange market is very small compared to *ex post* real interest differentials.

¹⁹ $E_t[E_{t+i}(x_{t+i})] = E_t(x_{t+i})$.

$$q_t = E_t[q_{t+j+1}] + \sum_{i=0}^j E_t[d_{t+i-1, t+i}] + \sum_{i=0}^j E_t[\delta_{t+i}] \quad [1']$$

$$= w_t + \sum_{i=0}^{\infty} E_t[d_{t+i-1, t+i}] + \sum_{i=0}^{\infty} E_t[\delta_{t+i}] \quad [2a]$$

where w_t is the expected long-run log real exchange rate calculated as $\lim_{i \rightarrow \infty} E_t[q_{t+i}]$.²⁰ We model w_t by a random walk. This specification is clearly less restrictive than earlier approaches which assumed w_t to be fixed. However, there are two features of w_t which must be noted:

- (i) If q_t is indistinguishable from a random walk (Meese and Rogoff (1983)) and $E_t(d_{t,t+1})$ is stationary, then the **long-run** movements in q_t will be explained by w_t *a priori*. However, this scenario does not impose predominant explanatory power on the long-run real exchange rate for **short-run** movements in q_t .
- (ii) w_t is clearly a "catch-all" for all other influences on q_t apart from the real interest differential.

One way of accessing real interest rate data is to use long-term (typically ten-year) real interest rates across countries.²¹ This involves assuming the expectations hypothesis of the term structure of interest rates. This theory, proposing no transactions costs and rational expectations, leads to the conclusion that the long rate is an average of current and expected short rates. Most empirical tests find no evidence supporting the expectations hypothesis (Fama and Bliss (1987), Shiller (1979)). Furthermore, the use of *ex post* long-term real interest rates truncates the sample period.²² Therefore, proxies for long-term inflation expectations have to be

²⁰ Campbell and Clarida note that assuming this limit to exist requires that *ex post* real interest differentials follow a stationary stochastic process with zero mean. For evidence supporting stationary short-term real interest differentials see Meese and Rogoff (1988).

²¹ Refer, for example, to Sachs (1985) and Shafer and Loopenko (1983).

²² The most recent *ex post* real ten-year interest rate that could be examined in this paper, would be a 1981 observation.

calculated. These proxies introduce measurement error and bias.²³ In order to avoid introducing unnecessary error, this model relaxes the traditional assumption of the expectations hypothesis and investigates the link between **short-term** real interest differentials and the real exchange rate.

The real exchange rate is thus expressed as a function of three unobservable variables: the expected long-run equilibrium exchange rate, the undiscounted sum of all current and expected future one-period interest differentials, and the undiscounted sum of current and expected future "risk premia".

Using the notation that the error terms $u_{i,t}$, $i = 1,2,3$ are white noise processes, seven time series properties are imposed on the model:

1. Expectations are assumed to be rational;

$$2. \quad E_t[d_{t,t+1}] = \rho E_t[d_{t-1,t}] + v_t \quad ; \quad \rho < 1 \quad [3]$$

The *ex ante* real interest differential is assumed to follow a stationary AR(1) process. This formulation, which Campbell and Clarida (1987) note is consistent with Dornbusch's (1976) overshooting model, is both simple and tractable.

$$3. \quad v_t = \lambda e_{t-1} + u_{1,t} \quad [4]$$

Innovations to this AR(1) process comprise two separate effects. A proportion of the error associated with forecasting inflation differentials in period $t-1$, is denoted by the term λe_{t-1} where $e_t = d_t - E_t[d_{t,t+1}]$. $u_{1,t}$ is an independent error term, uncorrelated with inflation surprise.

The second and third restrictions imply that d_t is an ARMA(1,1) process. An examination of the autocorrelations of d_t (refer Table 1 and discussion in

²³ Campbell and Clarida (1987) present survey evidence (Hoey (1986)) to show that conclusions drawn from studies based on an examination of proxied long-term real interest rates are very sensitive to the author's method of calculating inflationary expectations.

Section 5) do not exclude this as a possibility, except in the case of the United Kingdom.²⁴

$$4. \quad e_t = u_{2,t} \quad [5]$$

Dictated by the assumption of rational expectations, the inflation surprise is white noise.

$$5. \quad w_t = w_{t-1} + u_{3,t} \quad [6]$$

Restriction 5 imposes unforecastability upon changes in the expected long-run real exchange rate.

$$6. \quad E[u_{i,t}u_{j,t-k}] = 0 ; \quad i,j = 1,2,3 ; \quad k \geq 1 \quad [7]$$

All error terms are white noise. That is, inflation forecast errors ($u_{2,t}$) are assumed to be uncorrelated with past innovations in the expected long-run real exchange rate ($u_{3,t-k}$) and uncorrelated with innovations to $E_t(d_{t,t+1})$ unrelated to inflation surprise, ($u_{1,t-k}$). This assumption follows directly from the first assumption since rational expectations imply unforecastable inflation forecast errors. Therefore, all past information, including past innovations in the expected long-run real exchange rate and in real interest differentials, contains no information about the future value of these errors.

Past inflation forecast errors are assumed to be uncorrelated with innovations in the real interest differential and with the expected long-run real exchange rate. This restriction is commonly used in rational expectations models of real interest rates (see Hamilton (1985)), although Fama and Gibbons (1982) find a negative correlation between expected inflation and real interest rates. Mishkin (1987) uses this evidence to infer some negative correlation between past positive inflation forecast errors and future innovations in real interest differentials. In general, an innovation in the expected long-run real exchange rate (w_t) would be influenced by past inflation forecast errors (e_{t-1}), as are innovations in the

²⁴ This probably explains the poor performance of the model for the United Kingdom.

ex ante differential (v_t). However, (as per restriction 5), this correlation is excluded for simplicity.²⁵

The correlation between innovations in the expected long-run real exchange rate and innovations in the real interest differentials is left unrestricted.

$$7. \quad \delta_t = (\beta-1)E_t[d_{t,t+1}] \quad [8a]$$

δ_t is assumed to be proportional to the *ex ante* real interest differential.²⁶ A value for β less than one corresponds to a conventional risk premium. Both variables are unobservable and endogenous.

$$\delta_t = 0 \Leftrightarrow \beta = 1 \quad [8b]$$

The real interest differential is exactly equal to the expected exchange rate change when $\beta = 1$. This form of the model is the pure uncovered interest parity specification.

Having thus defined the constant of proportionality in [8a], equation [1] becomes:

$$q_t = E_t[q_{t+1}] + \beta\{E_t[d_{t,t+1}]\} \quad [9]$$

Equation [9] can then be solved forward to yield the long-run solution for the linear risk premium model:

$$q_t = \sum_{j=0}^{\infty} \beta E_t[d_{t+j-1, t+j}] + w_t \quad [2b]$$

Applying iterative expectations, equation [3] is solved forward in time. The assumption that $\rho < 1$ (i.e. $E_t(d_{t,t+1})$ is stationary) ensures that the infinite

²⁵ Estimation of the linear risk premium model for the trade-weighted index with this correlation left unrestricted yielded estimates very similar to those reported in Tables 2 and 4, and the increase in the log likelihood function was insignificant.

²⁶ While Campbell and Clarida (1987) insist on theoretical purity in the rest of their model, this assumption is quite arbitrary. However, as discussed in Section 5, the choice of the assumption about whether the risk premium is zero or proportional to the real interest differential does not affect the conclusions of the paper.

sum of the resulting geometric progression converges. The resulting equation is substituted into [2b] to yield the following expression for q_t :

$$q_t = \left[\frac{\beta}{1-\rho} \right] E_t[d_{t,t+1}] + w_t \quad [10]$$

Together with the time series properties (1) - (7), equation [10] defines the linear risk premium model. When $\beta = 1$, equation [10] defines the uncovered interest parity model.

Huizinga (1987) uses a univariate time series model to decompose shocks to the real exchange rate into permanent and transitory disturbances (via the Beveridge-Nelson (1981) method²⁷). Campbell and Clarida (1987) go a step further by identifying the transitory component with an endogenous economic variable: namely, the *ex ante* real interest differential.²⁸

Within this framework, two parameters determine the extent to which *ex ante* real interest differentials could be expected to explain the greater proportion of monthly fluctuations in the real exchange rate:

- (i) ρ is large if real interest differentials are highly persistent. This means that foreign investors perceive the return on Australian assets to be stable and relatively secure. According to the monetary approach, this encourages capital into the country and appreciates our dollar. This mechanism is illustrated by equations [3] and [10]. As the coefficient, ρ ,

²⁷ The Beveridge-Nelson (1981) decomposition can be explained as:

$$q_t = a_t + b_t$$

where a_t , interpreted as the permanent component, follows a random walk and b_t , interpreted as the transitory component, is a stationary process.

Like Huizinga (1987) and Campbell and Clarida (1987), we take advantage of the fact that a_t can conveniently be interpreted as the expected long-run real exchange rate. Assuming the logarithm of the long-run real exchange rate to be a random walk implies that actual changes in this variable are permanent changes.

Campbell and Clarida (1987) identify b_t with short-term real interest differentials (see Meese and Rogoff (1988) for evidence of stationarity in short-term real interest differentials).

²⁸ This abstracts from reality to the extent to which other transitory components (e.g. temporary terms of trade shocks) influence the exchange rate. Certainly this is more of an issue in Australia than in the US.

in equation [3] approaches 1, the AR(1) process describing the *ex ante* real interest differential becomes increasingly persistent. The coefficient on the *ex ante* real interest differential in [10] approaches ∞ , implying a larger real exchange rate change over the long run given any particular real interest differential.

- (ii) The second parameter, β , is unrestricted in the linear risk premium model (equation [10]). If investors are risk averse (i.e. $\beta < 1$) then the effect on the exchange rate of an increase in, say, the return on Australian assets *vis-a-vis* the rest of the world, will be muted.

4. ECONOMETRIC METHODOLOGY

A state space representation of the structural model is derived to avoid the computational problems arising from the inclusion of unobserved variables. The resulting representation can then be estimated using the Kalman filter. Harvey (1988)²⁹ describes this method of model specification and estimation as being "... in many ways preferable to the more conventional approach based on ARIMA processes". This section briefly reviews Kalman filter estimation and its application in our paper (see Harvey (1981, 1988, and 1989) for a detailed discussion).

4.1 Motivation for Choice of Estimation Technique

The Kalman filter was derived for applications in control engineering (Kalman (1960)). In this paper, its importance for maximum likelihood estimation is exploited. Depending on the form of the function to be optimised, maximum likelihood problems can be solved either analytically or numerically. The analytical solution involves equating the vector of first derivatives to zero and solving for the parameters. More complex functions are solved numerically. Two approaches exist for numerical optimisation:

²⁹ Harvey (1988), Chapter 8, p.285.

- (i) Direct evaluation is used for functions which can be calculated but not differentiated.
- (ii) Kalman filter evaluation is used either because the likelihood function cannot be evaluated directly, or because the system contains unobservable variables.

Our paper uses the Kalman filter because it permits a more flexible and rigorous treatment of the unobservable variables, the *ex ante* real interest differential, $E_t(d_{t,t+1})$, the long-run real exchange rate, w_t , and the forecast error, e_t .³⁰ The alternative approach, selecting proxy variables before estimation, is subject to measurement error and may not be optimal in any well-defined sense.

The more standard method of estimating models with unobservable components is that of "instrumental variables" (IV). This technique relies on choosing a proxy (or "instrument") which, while highly correlated with its unobservable equivalent, is not correlated with the model's disturbance terms. The parameter estimates obtained in this way are sensitive to the choice of instrument and may be biased. The IV technique would also restrict the dynamics of Campbell and Clarida's (1987) structural model:

- The long-run real equilibrium exchange rate and the risk premium are allowed to vary over time. Determining an instrument for either of these variables is problematic. With respect to w_t , a separate model could be estimated, but this exercise lies beyond the scope of our paper. The temptation would be to assume w_t constant as in Shafer and Loopesko (1983). Moreover, these assumptions reduce the explanatory power of the model.

³⁰ Clearly, the specification of the unobservable components in the state space model must depend, to some extent, on *a priori* considerations. However, estimates of the unobservable variables generated by the Kalman filter are optimal in the minimum MSE sense.

- We suggest that d_{t-1} would be an appropriate instrument for $E_t(d_{t,t+1})$.³¹ Using IV, equation [10] could be estimated as:

$$q_t = \left[\frac{\beta}{1-\rho} \right] d_{t-1} + w + \omega_t \quad [10']$$

where:

- w is a constant term representing the long-run real exchange rate;
- ω_t is a composite error term incorporating $u_{1,t}$, $u_{2,t}$ and $u_{3,t}$.

Estimation of [10'] reveals strong serial correlation. This can be corrected by using an appropriate estimation technique (for example simple Cochrane-Orcutt AR(1) modification). Alternatively, the model can be estimated in first differences. Although this approach would yield coefficient estimates, it would eliminate evaluation of the proportion of the variation in q_t explained by $E_t(d_{t,t+1})$. Rather, it would enable us to evaluate the proportion of the variation in the change in q_t explained by the change in the proxy for $E_t(d_{t,t+1})$.

- As well, the Kalman filter allows solution of the model regardless of stationary/non-stationary data considerations.

³¹ Since the second time-series assumption describes $E_t(d_{t,t+1})$ as a stationary AR(1) process, d_t can be shown to follow an ARMA(1,1) process of the form:

$$d_t = \rho d_{t-1} + u_{1,t} + u_{2,t} + (\lambda - \rho)u_{2,t-1}$$

(See Fama and Gibbons (1982) for other working to show that if *ex ante* real interest rates follow a univariate AR(1), then *ex post* real rates must follow a univariate ARMA (1,1)). The complex error structure of this equation highlights a limitation to the IV approach (given that the error structure proposed by Campbell and Clarida (1987) is a plausible one). Restricting $u_{1,t}$ to zero transforms this equation into a standard ARMA(1,1) that may be estimated by IV. The resulting estimate of ρ using data for our trade-weighted index is not significantly different from one. In contrast, estimates of ρ generated by the Kalman filter are 0.59 in the linear risk premium model and 0.62 in the uncovered interest parity model. These estimates accord with those of Gruen and Wilkinson (1991) who, using the augmented Dickey-Fuller test, estimate equivalent coefficients at 0.66.

Nevertheless, given the IV approach to estimation, $\rho = 1$, and d_{t-1} becomes a suitable instrument for $E_t(d_{t,t+1})$.

4.2 Application of the Kalman Filter

The filter can only be applied once the structural model is cast in a framework which enables the treatment of its unobservable components. A linear dynamic system can be written in the so-called *state space* form.³² A key feature of this representation is the presence of an unobservable vector, α_t , called the state vector. In the context of this paper, α_t is $[E_t(d_{t,t+1}) \ e_t \ w_t]'$.

The state space form consists of two equations: the measurement equation and the transition equation. The measurement equation relates the vector of unobservable, explanatory variables, α_t , to a vector of observable variables, y_t , such that:

$$y_t = Z \alpha_t \quad [11]$$

(2x1) (2x3)(3x1)

For our model this takes the form:

$$\begin{bmatrix} q_t \\ d_t \end{bmatrix} = \begin{bmatrix} \beta/(1-\rho) & 0 & 1 \\ 1 & 1 & 0 \end{bmatrix} \begin{bmatrix} E_t[d_{t,t+1}] \\ e_t \\ w_t \end{bmatrix} \quad [12]$$

$\beta \neq 1$: Linear risk premium model;

$\beta = 1$: Uncovered interest parity model;

β and ρ are assumed constant for these two models and, in keeping with our theory-based relationships, an error term - generally added onto the right-hand-side of equation [11] - is not included.

³² Such models, driven by innovations of some macroeconomic time series, are sometimes referred to as "innovation models" (see Aoki and Havenner (1986)).

The second equation in the state space formulation is the transition equation. This describes the evolution of the state vector over time such that:

$$\alpha_t = T \alpha_{t-1} + \eta_t \quad ; \quad \eta_t \sim \text{NID}(0, Q_t) \quad [13]$$

(3x1) (3x3) (3x1) (3x1)

For our model:

$$\begin{bmatrix} E_t[d_{t,t+1}] \\ e_t \\ w_t \end{bmatrix} = \begin{bmatrix} \rho & \lambda & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} E_{t-1}[d_{t-2,t-1}] \\ e_{t-1} \\ w_{t-1} \end{bmatrix} + \begin{bmatrix} u_{1t} \\ u_{2t} \\ u_{3t} \end{bmatrix} \quad [14]$$

u_{1t} is the innovation, unrelated to inflation surprise, associated with the *ex ante* real interest differential. u_{2t} is the inflation forecast error; rational expectations dictates that this is uncorrelated contemporaneously with either u_{1t} or u_{3t} . u_{3t} is the innovation to the long-run real exchange rate. All errors are serially uncorrelated and have standard deviations σ_1 , σ_2 and σ_3 respectively. σ_{13} , the contemporaneous covariance between u_{1t} and u_{3t} , is not restricted.

Initialised with a set of priors, the Kalman filter estimates (that is, it "predicts" and "updates") the vector of unobservable variables, α_t , period by period. "Predicting" means predicting α_t given the information set of period $t-1$. "Updating" means using a one-step-ahead forecast error of y_t (necessarily calculated by the filter at each execution) to update the predicted value of α_t . This two stage procedure, carried out at each new time period, defines the Kalman filter. The linear estimate of α_t , denoted $\hat{\alpha}_t$, is optimal in the minimum mean square error sense. The filter also provides the error covariance matrix of $\hat{\alpha}_t$, denoted P_t . Starting values for $\hat{\alpha}_t$ and P_0 must be provided. At time period t , observations $y_1, y_2 \dots y_t$ are available.

Apart from its ability to generate optimal estimates (as opposed to proxies) for the vector of unobservable variables, α_t , the Kalman filter may be used in maximum likelihood estimation. Its execution generates a time series for the one-step-ahead prediction errors made in estimating the vector of

observable, dependent variables, y_t . Harvey (1985) notes that the likelihood can be calculated from these normally distributed errors and their variances. Estimation of our model yields a bivariate normal log likelihood function which can be expressed as:

$$\log L(y, \theta) = -\ln 2\pi - 0.5 \ln |F_t| - 0.5 v_t' F_t^{-1} v_t \quad [15]$$

where:

- v_t is the set of one-step-ahead prediction errors made in estimating the vector of observable, explanatory variables, y_t (the real exchange rate, q_t , and the *ex post* real interest differential, d_t , in our model). That is, $v_t = y_t - E_{t-1}[y_t]$. The one-step-ahead prediction of the observation y_t is calculated as: $E_{t-1}[y_t] = Z E_{t-1}[\hat{\alpha}_t]$ where $E_{t-1}[\hat{\alpha}_t]$ is the optimal predictor of the state vector, α_t , given observations up to $t-1$ and Z is the matrix of coefficients on the state vector.
- F_t is the estimated covariance matrix of these one-step-ahead prediction errors.

Hence, the Kalman filter can be run in conjunction with an optimisation routine³³ - calculating the likelihood where required by the routine. This is what it means to "estimate" a model using the Kalman filter. For each run through the filter, the coefficients on the model are fixed. However, the likelihood function [15] can be calculated for each run.³⁴

Equations [12] and [14] are estimated with the Kalman filter. The model contains seven unknown parameters, $\beta, \rho, \lambda, \sigma_1, \sigma_2, \sigma_3$ and σ_{13} . Where equality between the expected real exchange rate change and the real interest differential is imposed, β is set equal to one. We estimate models for Australia's real exchange rate with the United States, Japan, West

³³ Since it is explicitly designed for optimisation of likelihood functions, the Berndt-Hall-Hausman (1974) algorithm (BHHH) is used on the final iteration for all models, with the exception of the United Kingdom. BHHH uses a modified method of scoring (see Berndt et al. (1974)) to optimise the function. On the final iteration of the United Kingdom's uncovered interest parity model, a more general purpose algorithm is applied.

³⁴ See Appendix 2 for a flow chart of our estimation procedure.

Germany and the United Kingdom, as well as a trade-weighted index. In each case, models are estimated with β unrestricted and with β restricted to one.

4.3 Variance Decomposition

Innovation relationships were investigated by expanding the one-period ahead variance of the actual dependent variable (namely, the observable real exchange rate) in the following manner.

Equation [10] implies:

$$\sigma_{q^2} = \left[\frac{\beta^2}{(1-\rho)^2} \right] \sigma_{E(d)}^2 + \sigma_w^2 + 2\sigma_{E(d),w} \quad [16]$$

where the subscripts of σ carry their obvious meaning (refer equation [10]).

The variance T periods ahead (i.e., $E_t(q_{t+T} - E_t(q_{t+T}))^2$) is

$$\begin{aligned} \sigma_{q^2}^2(T) = & \frac{\beta^2}{(1-\rho)^2} [\lambda^2 \sigma_2^2 + \sigma_1^2] (1 + \rho^2 + \rho^4 + \dots + \rho^{2(T-1)}) + T\sigma_3^2 \\ & + \frac{2\beta\sigma_{13}}{1-\rho} (1 + \rho + \rho^2 + \dots + \rho^{T-1}) \end{aligned} \quad [17]$$

as $T \rightarrow \infty$, the first and third terms of the RHS converge, i.e.:

$$\lim_{T \rightarrow \infty} \sigma_{q^2}^2(T) = \frac{\beta^2 [\lambda^2 \sigma_2^2 + \sigma_1^2]}{(1-\rho)^3(1+\rho)} + \lim_{T \rightarrow \infty} T\sigma_3^2 + \frac{2\beta\sigma_{13}}{(1-\rho)^2}$$

Clearly, a variance decomposition calculated over a long time horizon will always attribute a very large proportion of the variation in q_T to w_t *a priori*. This is a direct result of the assumption that w_t is a random walk and that $E_t(d_t)$ is stationary. In the limit, w_t explains 100 per cent of the

variation in q_T . However, what is not imposed is how much of the variability in q_T is explained by w_t in the short run. In principle, ρ could be very close to one, in which case $E_t(d_{t,t+1})$ could explain a large proportion of the variability of q_t for a small value of T . Using the parameter estimates obtained from the Kalman filter and the optimisation algorithm, the appropriate expression for the variance (equation [17] with $T=1$) is used to calculate estimated variance decompositions. That is, each of the three terms on the right hand side of equation [17] are expressed as a percentage of $\sigma_{q(1)}^2$ and reported in Tables 4 and 5.

5. ESTIMATION RESULTS

Summary statistics for all the data are presented in Table 1 (refer to Appendix 1 for a further discussion of data methods and sources). The first two rows display the sample means and standard deviations of real exchange rate changes over the post-float period. The sample means are positive, corresponding to real Australian dollar depreciation, for all countries except the United States and the trade-weighted index. The standard deviations for the United States and in the case of the trade-weighted index are around 45 per cent per month **at an annualized rate**, while other countries' exceed 50 per cent.

The next two rows of the table detail the same statistics for *ex post* real interest differentials. Sample means are negative across the board, indicating that real interest rates were higher on average in Australia than overseas. Sample standard deviations range from 3.5 per cent to around 6.5 per cent, highlighting the much greater volatility of real exchange rates compared with *ex post* real interest differentials over the post-float period.

The rational expectations framework implies that the *ex post* real interest differential comprises its *ex ante* equivalent and an unforecastable error. The variance of the *ex post* differential therefore forms the upper bound on the variance of the *ex ante*. Hence, an interpretation of the data consistent with rational expectations implies that the *ex ante* real interest differential in Australia over the post-float period has a low variance. A comparison

Table 1: Summary Statistics, 1983:12 - 1990:09

| Statistic [†] | Country | | | | |
|---|---------|--------|---------|--------|--------|
| | U.S. | Japan | Germany | U.K. | TWI* |
| Mean change in real exchange rate | -1.740 | 2.970 | 4.190 | 4.280 | -1.390 |
| Standard deviation of change in real exchange rate | 46.200 | 57.100 | 56.400 | 53.500 | 45.200 |
| Mean real interest differential | -3.360 | -3.810 | -3.770 | -2.110 | -3.640 |
| Standard deviation of <i>ex post</i> real interest differential | 4.030 | 6.550 | 3.420 | 6.240 | 3.870 |
| Correlation of change in real exchange rate and <i>ex post</i> real interest differential | -0.062 | -0.006 | -0.039 | -0.171 | -0.019 |
| Autocorrelations of <i>ex post</i> real interest differential: | | | | | |
| 1 | 0.710 | 0.348 | 0.750 | 0.241 | 0.656 |
| 2 | 0.515 | 0.084 | 0.595 | 0.086 | 0.424 |
| 3 | 0.443 | 0.161 | 0.550 | -0.136 | 0.484 |
| 4 | 0.427 | 0.240 | 0.491 | 0.029 | 0.510 |
| 5 | 0.453 | 0.320 | 0.390 | 0.156 | 0.473 |
| 6 | 0.419 | 0.402 | 0.325 | 0.182 | 0.470 |

* United States 42.65%, Japan 35.63%, West Germany 21.74%.

[†] All results (except autocorrelations) are annualized.

with the variance of the real exchange rate could lead us to believe that real *ex ante* interest differentials are not variable enough to be capable of explaining the high variance of the real exchange rate. The econometric results of our paper provide more rigorous evidence of this.

The first six autocorrelations of the *ex post* real interest differential, calculated with the assumption of a zero true mean for the real interest differential, are reported at the bottom of Table 1. Alternatively, assuming the sample mean differential to be the true mean, similar but smaller autocorrelation values are obtained. Data for the United States, Germany and the trade weighted index reveal the largest autocorrelations, and suggest the possibility that the *ex post* real interest differentials follow some ARMA(p,q) process. Remaining countries' autocorrelations are smaller, beginning at 0.24 for the United Kingdom and 0.35 for Japan, and exhibit a much more irregular evolution.

Tables 2 to 5 provide estimation results for each of the bilateral currencies and for the trade-weighted index. Results are given for the two versions of the model. All coefficients, β , ρ and λ are reported with asymptotic standard errors, together with the standard deviations of the error terms, σ_1 , σ_2 and σ_3 and the correlation $\sigma_{13}/\sigma_1\sigma_3$. Variance decompositions reveal some of the implications of these estimates.

Parameter estimates are detailed in Tables 2 and 3 (following). For each model, log likelihood ratio tests fail to reject the hypothesis that $\beta = 1$ at a 95 per cent significance level. With $\beta = 1$, the risk premium is equal to zero. The low power of these tests, however, means that the restricted models (detailed in Table 3) could not be rejected with β equivalent to values less than 1. Remaining coefficient, variance and covariance estimates are robust across alternative restrictions over β .

Therefore, the value ascribed the time-varying risk premium in this model does not affect any of the conclusions. This result is disappointing since without conclusive evidence to either accept or reject the existence of a risk premium, it is not clear whether Campbell and Clarida's (1987) model collapses to a simple UIP specification for which little empirical support can be found.

Table 2: Parameter Estimates - Linear Risk Premium Model

| | Trade-Weighted Index | United States | West Germany | United Kingdom | Japan |
|--------------------------------|-------------------------|---------------|--------------|----------------|-------------|
| <i>Log Likelihood:</i> | 494.54 | 500.29 | 488.82 | 442.87 | 427.53 |
| Parameters | s.e. | s.e. | s.e. | s.e. | s.e. |
| β | 1.86 (1.50) | 0.88 (1.55) | 0.85 (1.26) | 1.57 (5.27) | 0.33 (2.63) |
| ρ | 0.59 (0.15) | 0.85 (0.11) | 0.76 (0.13) | 0.33 (0.69) | 0.19 (0.34) |
| λ | 0.88 (0.33) | 0.43 (0.48) | 6.87 (--) | -0.54 (17.78) | 5.41 (--) |
| σ_1 | 2.51 | 2.63 | 3.36 | 5.72 | 0.38 |
| σ_2 | 3.00 | 2.28 | 0.01 | 2.14 | 1.36 |
| σ_3 | 44.53 | 45.93 | 56.25 | 54.88 | 56.83 |
| $\sigma_{13}/\sigma_1\sigma_3$ | -0.30 | -0.20 | -0.12 | -0.30 | -1.00 |

Table 3: Parameter Estimates - Uncovered Interest Parity Model

| | Trade-Weighted Index | United States | West Germany | United Kingdom | Japan |
|--------------------------------|-------------------------|---------------|--------------|----------------|--------------|
| <i>Log Likelihood:</i> | 494.37 | 500.29 | 490.19 | 442.76 | 427.50 |
| Parameters | s.e. | s.e. | s.e. | s.e. | s.e. |
| β | 1.00 (fixed) | 1.00 (fixed) | 1.00 (fixed) | 1.00 (fixed) | 1.00 (fixed) |
| ρ | 0.62 (0.15) | 0.84 (0.11) | 0.84 (0.11) | 0.37 (--) | 0.11 (0.33) |
| λ | 0.79 (0.21) | 0.46 (0.44) | 0.52 (0.37) | -2.31 (--) | 0.76 (1.95) |
| σ_1 | 1.61 | 2.65 | 2.24 | 4.87 | 4.93 |
| σ_2 | 3.60 | 2.27 | 2.41 | 1.46 | 4.57 |
| σ_3 | 44.06 | 46.17 | 54.35 | 53.90 | 56.93 |
| $\sigma_{13}/\sigma_1\sigma_3$ | -0.08 | -0.23 | -0.09 | -0.27 | -0.13 |

Remaining parameter estimates are interpreted below. For explanatory convenience, we concentrate on the results obtained from the uncovered interest parity model of Australia's real exchange rate with a constructed trade-weighted index³⁵ (refer Table 3):

- ρ is estimated at 0.62, revealing moderate persistence in the *ex ante* real short-term interest differential between Australia and this trade-weighted index.
- λ is estimated at 0.79, suggesting that innovations to the AR(1) process involve errors associated with forecasting inflation in period $t-1$.

From the parameter estimates, we can calculate the coefficient on the *ex ante* real interest differential in equation [10] described in section 3. For the trade weighted index case, this coefficient assumes a value of 2.6 in the uncovered interest parity model and 4.5 in the linear risk premium model.³⁶ This implies, for the former model, that a 1 percentage point increase in the annual *ex ante* real interest differential, *ceteris paribus*, will induce a 0.2 percent change in the real exchange rate.

Estimates of the variances σ_1 , σ_2 , σ_3 and the covariance σ_{13} are used for variance decompositions reported in Tables 4 and 5.³⁷

For the trade-weighted index case, the proportion of the variance of the real exchange rate (q_t) accounted for by the innovation variance of the *ex ante* real interest differential (v_t) is approximately 3.79 per cent for the uncovered interest parity model and around 14.28 per cent for the linear risk premium model. Robust across all models and countries, the innovation variance of the expected long-run real exchange rate must explain the greatest proportion of the variance of the actual real exchange rate because *ex ante* real interest differentials do not.

³⁵ United States 42.65%; Japan 35.63%; and West Germany 21.74%.

³⁶ The spread of these coefficient values is a reflection of the lack of certainty associated with point estimates.

³⁷ Structural interpretation of the variance decompositions is not unambiguous; they do not imply unidirectional causality.

**Table 4: Variance Decompositions and Correlations -
Linear Risk Premium Model**

| Country | Variance Decompositions (%) | | | | |
|----------------|-----------------------------|-------|------------------|--------|------------|
| | d_t | | q_t | | |
| | $E_t(d_{t,t+1})$ | e_t | $E_t(d_{t,t+1})$ | w_t | Covariance |
| TWI | 69.54 | 30.46 | 14.28 | 101.64 | -15.92 |
| United States | 84.39 | 15.61 | 12.67 | 100.75 | -13.42 |
| West Germany | 100.00 | 0.00 | 4.49 | 100.73 | -5.21 |
| United Kingdom | 89.33 | 10.67 | 6.93 | 109.46 | -16.39 |
| Japan | 96.82 | 3.18 | 0.27 | 100.27 | -0.54 |

**Table 5: Variance Decompositions and Correlations -
Uncovered Interest Parity Model**

| Country | Variance Decompositions (%) | | | | |
|----------------|-----------------------------|-------|------------------|--------|------------|
| | d_t | | q_t | | |
| | $E_t(d_{t,t+1})$ | e_t | $E_t(d_{t,t+1})$ | w_t | Covariance |
| TWI | 57.32 | 42.68 | 3.79 | 97.80 | -1.59 |
| United States | 84.29 | 15.71 | 15.07 | 101.85 | -16.92 |
| West Germany | 79.77 | 20.23 | 8.74 | 95.66 | -4.40 |
| United Kingdom | 95.04 | 4.96 | 3.21 | 104.96 | -8.17 |
| Japan | 63.80 | 36.20 | 1.45 | 101.18 | -2.63 |

It is important to note that when the proportion of the variance of the real exchange rate accounted for by the innovation variance of the expected long-run real exchange rate (u_{3t}) exceeds 100 per cent, this does not imply zero explanatory power for the *ex ante* real interest differential. The existence of covariance between the two explanatory variables makes direct interpretation of the variance decompositions difficult. However, the overwhelming result remains clear: the movement in the real Australian dollar exchange rate over the period of this study is predominantly explained by the innovation variance of the long-run real exchange rate.

Ex ante real interest differentials are incapable of explaining the variability of the Australian dollar real exchange rate because the models' estimates of their persistence (that is, the ρ values), although quite high in some cases (estimates for ρ exceed 0.6 in half of the models estimated) are nevertheless insufficient to compensate their minor standard deviation³⁸ relative to that of changes in the real exchange rate.³⁹

In summary, the results implied by direct application of Campbell and Clarida's (1987) methodology to Australian data do not contradict the idea that over a short horizon, a change in the level of Australia's real interest rates relative to the rest of the world *ceteris paribus*, exerts some small influence on the real exchange rate. They do indicate, however, that given the influence of all remaining factors, this effect is not substantial enough to be responsible for the monthly variability observed in the real exchange rate.

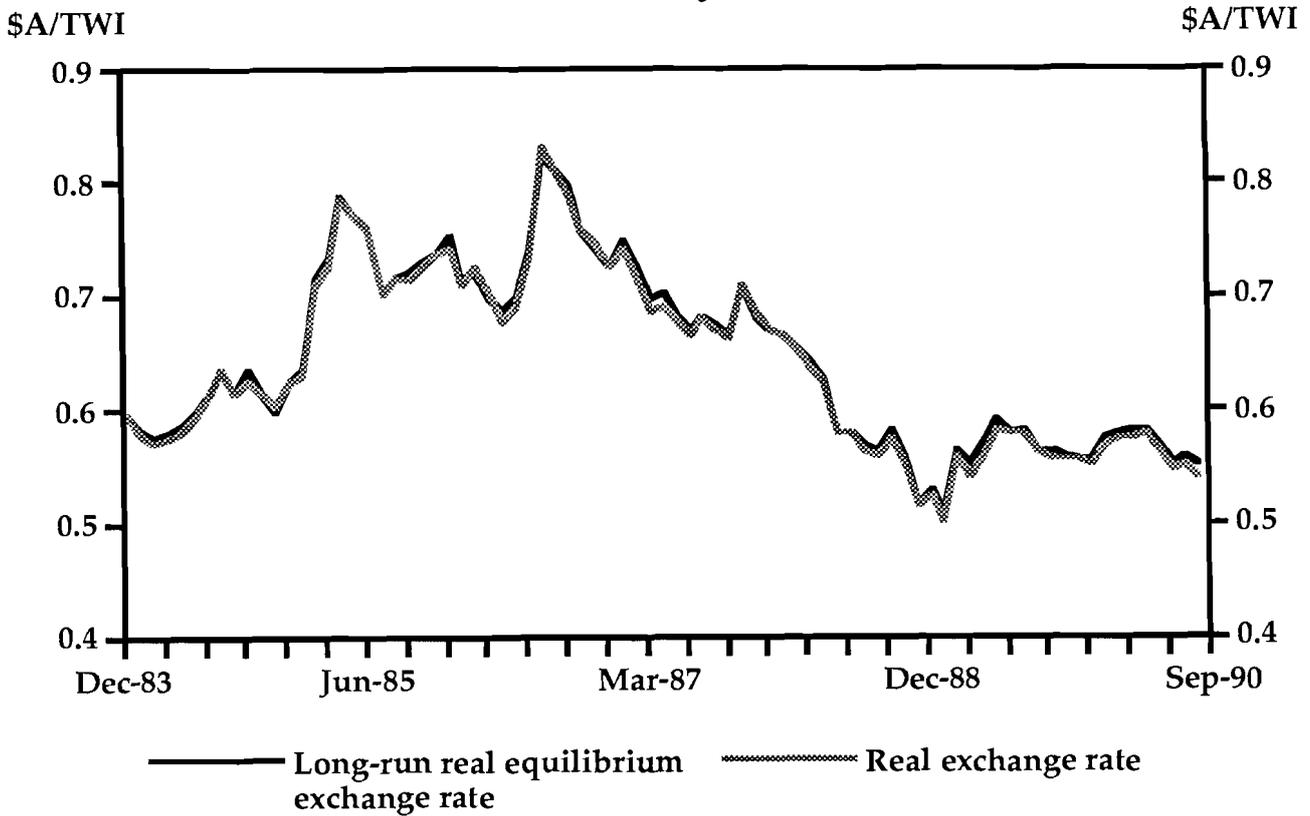
³⁸ The standard deviation of *ex post* real interest differentials, calculated in Table 1, is very small in comparison to that calculated for the mean change in the real exchange rate. Our rational expectations framework implies that the variance of the *ex post* differential forms the upper bound on the variance of the *ex ante* differential.

³⁹ Campbell and Clarida (1987) find that for the US, higher estimates of ρ are largely accompanied by a smaller share of the *ex ante* differential in the variance of the *ex post* differential. This is not always the case in the Australian data.

Also in contrast with the findings of Campbell and Clarida (1987), we find that *ex ante* real interest differentials, as opposed to inflation innovations, account for the greater proportion of the variance of *ex post* differentials. This is not surprising, since inflation exhibits very little variability over a short time horizon.

This conclusion can be illustrated graphically:

**Graph 1: The Real Exchange Rate and its Long-Run Real Equilibrium Rate
(Monthly)**

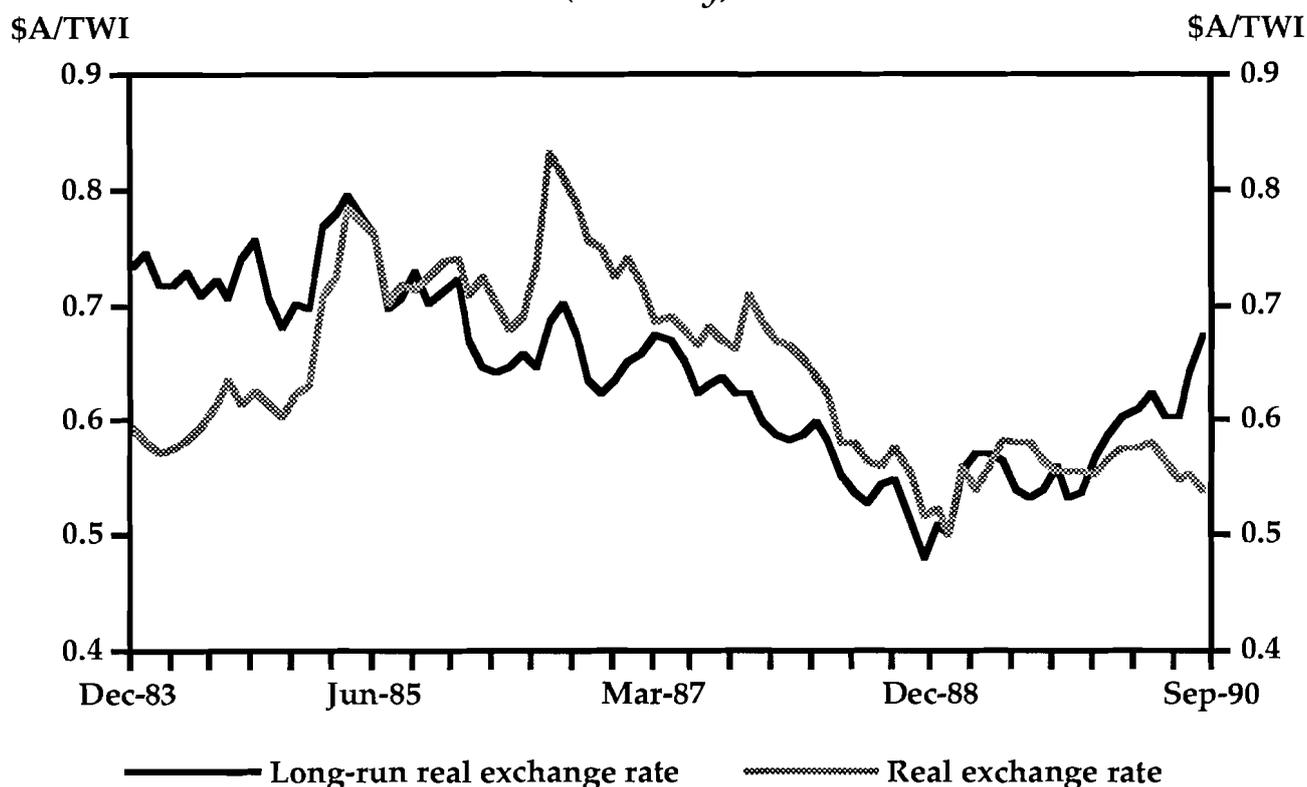


Graph 1 plots the real exchange rate against its observable long-run real equilibrium rate (w_t as estimated by our model) for Australia and a constructed trade-weighted index. They track so closely that the reader may be forgiven for not seeing two different sets of data. This

demonstrates that the $\left[\frac{\beta}{1-\rho} \right] E_t[d_{t,t+1}]$ term in equation [10] is so small as to be insignificant.

Graph 2 is offered as a counterpoint. It illustrates what the expected long-run real exchange rate would have looked like if the *ex ante* real interest differential between Australia and the trade-weighted index had been almost perfectly persistent. That is, we synthesize a series for the *ex post* real interest differential such that the model then estimates a value for ρ (the coefficient on the AR(1) process characterising the *ex ante* real interest

Graph 2: The Real Exchange Rate and a Synthetic Long-Run Real Equilibrium Rate
(Monthly)



differential) very close to 1. This significantly reduces the explanatory power of the expected long-run real exchange rate and verifies that w_t in Graph 1 is not just an artifact of the model. The lack of explanatory power found in short-term real interest differentials is not a function of the model's assumption that the expected long-run real exchange rate (like the actual real exchange rate) follows a random walk.⁴⁰

⁴⁰ It should also be noted that our sample period encompasses a regime shift in 1987. Between early 1985 and late 1987, monetary policy was directed towards both internal and external objectives. Short-term interest rates were used to directly target the exchange rate and the market expected central bank intervention to dampen any speculative activity. Post-1987, monetary policy became directed primarily towards an internal objective (Macfarlane and Tease (1989)). While it is possible that expectations for Australian dollar movements relative to interest rate changes may be different across these two periods, our results are not unique in the literature. We suspect that estimation of the model over a shorter sample period would produce negligible changes.

6. ASSESSMENTS AND CONCLUSIONS

Direct technical application of Campbell and Clarida's (1987) rational expectations model to Australian post-float data has eliminated real *ex ante* short-term interest differentials as a source of monthly variation in the real exchange rate. Our empirical results with Australian statistics are consistent with Campbell and Clarida's (1987) original findings for the United States.

The overwhelming conclusion is that, over the post-float period in Australia, the permanent component of the model, namely the expected long-run equilibrium real exchange rate, accounts for the greatest proportion of the month-to-month movements observed in the actual real exchange rate. This is because the transitory component, identified with *ex ante* short-term real interest differentials, does not. This result does not deny the existence of a weak short-run relationship between Australia's real exchange rate and real interest rate relative to that of the rest of the world. It does, however, indicate that given the influence of all remaining factors, this effect is not substantial enough to be responsible for the observed short-run dynamics of the real exchange rate.

We could interpret this finding as evidence that the *monetary* approach to real exchange rate determination, which focuses on the role of international capital flows, plays a secondary role in describing the path of the observed real Australian exchange rate. If rational expectations is a valid assumption, then this model suggests that random real shocks to variables like for example, commodity prices, determine the real value of the Australian dollar over the shorter run.

While the solution of Campbell and Clarida's (1987) model is complex, its underlying specification is both transparent and flexible. Its treatment of the unobservable explanatory variables represents an advance on earlier rational expectations models of the real exchange rate. It allows the expected long-run exchange rate to vary over time; provides for the existence of a time-varying risk premium; relaxes any dependence on the expectations hypothesis of the term-structure of interest rates; and allows the vector of explanatory variables to evolve slowly over time, as is appropriate for time-series data.

However, the success of Campbell and Clarida's methodology (although more flexible than earlier rational expectations models) relies on their assumptions about the nature of the foreign exchange market, namely that it is efficient (i.e. that expectations are rational). An efficient market in foreign exchange is one in which the majority of market participants have access to all available information which is reflected in the price; only *new* information moves the exchange rate (that is, the exchange rate is only affected to the extent that a particular outcome differs from the value expected by the market, even if that outcome is poor⁴¹); and all new information is a *shock* (it is completely random and thus, unpredictable). Allan et al. (1990, p. 97) point out that *consistent* profits in such a market are impossible to incur since "... no forecasting procedure can give more accurate forecasts than tossing a coin can".

While it has been both common and tractable to assume market efficiency for empirical work on the short-run determinants of exchange rates, anecdotal evidence from market participants suggests that it is inappropriate.⁴² We feel that the application of Campbell and Clarida's (1987) methodology to Australian data is compromised by the ambiguity surrounding the existence or otherwise of a risk premium. In the absence of a significant risk premium effect, their model would collapse to an uncovered interest parity relation. Such a relation has been discarded empirically. Gruen and Menzies (1991) propose that if the costs of sluggish portfolio adjustment are insignificant, then the failure of uncovered interest parity may result from the survival of near-rational agents in the foreign exchange market. This argument introduces the idea that the essential reason for the failure of uncovered interest parity may be that the assumption of rational expectations is inadequate in the foreign exchange market (Cutler, Poterba and Summers (1990)).

We believe the direction for future research into the **short-run** dynamics of the real exchange rate should allow for alternative expectations

⁴¹ Allan et al. (1990) provide a good example by using market reaction to a monthly current account deficit figure: "... if the market predicts that this figure will be AUD1 billion, this information will be reflected in the current exchange rate. [If the actual outcome is] AUD800 million, the dollar is likely to appreciate, even ... [though] this is a poor result."

⁴² They argue that consistent profits have been earned from taking positions in the foreign exchange market.

formation.⁴³ In so far as Campbell and Clarida's approach found little explanatory power for interest differentials, we feel that this should not necessarily be interpreted as evidence that real shocks determine **shorter-run** exchange rate movements. Rather, the result may be indicative of the inappropriateness of rational expectations.

In conclusion, this paper has replicated Campbell and Clarida's (1987) rational expectations model of the US foreign exchange market for Australian data. As in their study, little role is found for real interest differentials in real exchange rate determination over the short run. Within the confines of their assumptions, this leaves the long-run equilibrium real exchange rate to do all the explaining of the real exchange rate.

However, given the growing body of literature which now opposes both rational expectations and efficient foreign exchange markets, alternative explanations are possible. As canvassed earlier in the paper, a number of options, some of which entail the inefficient formation of expectations, could be the driving influence behind the monthly movements in the Australian dollar. In so far as monthly movements in the exchange rate motivate resource allocation decisions, less than rational expectations imply that these decisions are suboptimal.

⁴³ Miller and Weller (1991) provide a concise review of the most recent literature in this area including, for example, the investigation by De Long et al. (1990) into the survival of noise traders in financial markets.

Appendix 1: Data Methods and Sources

All data is sampled as at the last trading day of each month. This study covers the post-float period in Australia, extending from December 1983 to September 1990, inclusive.

All data is quoted according to the United States convention. All bilateral exchange rates (4 p.m., Sydney) are obtained from the International Department of the Reserve Bank of Australia, and are quoted directly, such that the value of one unit of foreign currency is expressed in terms of Australian dollars.

An index of trade-weighted exchange rates is also constructed. 1990 trade weights for all of Australia's trading partners are apportioned to one of three categories bearing the name of the largest component country (United States 42.65%, Japan 35.63% and West Germany 21.74%).

The three-month Eurocurrency rates for the United States, Japan, West Germany and the United Kingdom are obtained from the International Department of the Reserve Bank of Australia. The USD, JPY, and DEM rates are London rates. The first of these refers to the close while the latter two refer to midrates. The GBP rate is a Paris midrate. The 90-day bank bill rate is chosen as the representative short-term rate for Australia, and is obtained from the Domestic Markets Department of the Reserve Bank of Australia. A trade-weighted interest rate is constructed using the weights described above for the trade-weighted exchange rate index.

With the exception of Australia, consumer price indices are provided by the Overseas Economies Section of the Economic Analysis Department of the Reserve Bank of Australia. Quarterly Australian data, obtained from the ABS, are Medicare adjusted. Simple linear interpolation⁴⁴ provides monthly statistics.

⁴⁴ Linear rather than geometric interpolation is used since the rate of change of Australian CPI data is very small. We note that calculating monthly CPI data in this manner gives investors the benefit of information they do not yet possess.

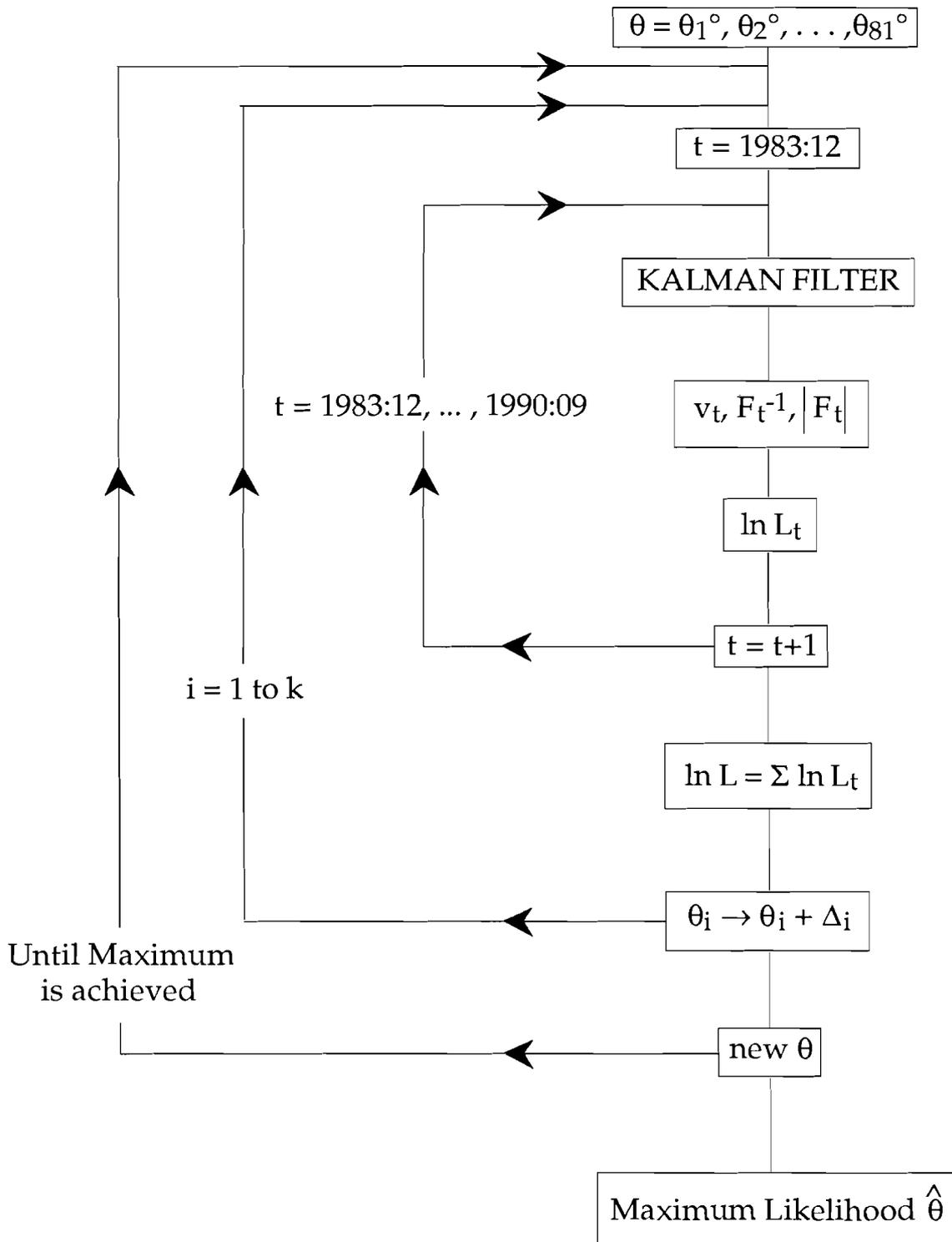
Real exchange rates are constructed as:

$$q_t = \ln[s_t(p_{f,t}/p_{d,t})] \quad [24]$$

where s_t is the nominal exchange rate and $p_{f,t}/p_{d,t}$ is the ratio of foreign to domestic CPIs.

Ex post real interest differentials are constructed by subtracting the actual inflation differential (as measured by one-period logarithmic differences in the ratio of foreign to domestic CPIs) at time t , from the nominal one-month interest differential at time $t-1$.

All results are reported in units of per cent per month at an annualized rate.

Appendix 2: Our Estimation Procedure - A Flow Chart⁴⁵

⁴⁵ This appendix benefits from discussion with Professor Howard Doran of the University of New England, and this diagrammatic representation of the Kalman Filter is an adaptation of that found in Doran (1990).

Our estimation process is activated with an initial choice of θ . The Kalman filter recursion begins at our first data observation, December 1983, and uses this first θ to compute the Kalman predictors via:

$$(i) \quad E_{t-1}(\hat{\alpha}_t) = T \hat{\alpha}_{t-1}$$

$$(ii) \quad E_{t-1}(P_t) = T P_{t-1} T' + Q_t$$

where:

- $E_{t-1}(\hat{\alpha}_t)$ is the minimum MS linear estimator of the state vector, α_t ;
- T is the matrix of coefficients on α_{t-1} in the transition equation of our state space model;
- P_t is the error covariance matrix of $\hat{\alpha}_t$;
- Q_t is the covariance matrix for η_t , the vector of disturbance terms in the transition equation.

The following "updating" equations are then applied:

$$(i) \quad \hat{\alpha}_t = E_{t-1}(\hat{\alpha}_t) + G_t v_t$$

where: $G_t = E_{t-1}(P_t) Z' F_t^{-1}$ is known as the Kalman Gain.

This "correction" term modifies the estimator, $\hat{\alpha}_{t|t-1}$;
and $v_t = y_t - Z E_{t-1}(\hat{\alpha}_t)$.

$$(ii) \quad P_t = E_{t-1}(P_t) - E_{t-1}(P_t) Z' F_t^{-1} Z E_{t-1}(P_t)$$

$$(iii) \quad F_t = Z E_{t-1}(P_t) Z'$$

where:

- Z is the vector of coefficients on α_t in the measurement equation;

- F_t is the estimated covariance matrix of the one-step-ahead prediction errors made in estimating the vector of observable explanatory variables and denoted v_t , such that:

$$v_t = y_t - E_{t-1}(y_t).$$

Therefore, three different errors play a role in the Kalman Filter:

- (i) $\alpha_t - \hat{\alpha}_t$, the error in the optimal predictor.
This error has covariance matrix P_t .
- (ii) $\theta_t = \alpha_t - E_{t-1}(\hat{\alpha}_t)$, where $E_{t-1}(\hat{\alpha}_t)$ is the optimal predictor of α_t given observations up to $t-1$.
 $E_{t-1}(P_t)$ is its covariance matrix.
- (iii) $v_t = y_t - E_{t-1}(y_t)$, where $E_{t-1}(y_t)$ is the one-step-ahead prediction of the observation y_t :

$$E_{t-1}(y_t) = Z E_{t-1}(\hat{\alpha}_t)$$

F_t is the covariance matrix of v_t .

Given $\hat{\alpha}_{t-1}$ and P_{t-1} , the workings of the filter can be described by a simple diagram:

| Predicting | Updating |
|--|--|
| $P_{t-1} \rightarrow E_{t-1}(P_t)$ | $\rightarrow F_t \rightarrow P_t$ |
| $\hat{\alpha}_{t-1} \rightarrow E_{t-1}(\hat{\alpha}_t)$ | $\rightarrow v_t \rightarrow \hat{\alpha}_t$ |

The normally distributed prediction errors and their variances generated in this way by the filter are used to calculate the likelihood function. This process is continued for each period in our data set.

The likelihood is then evaluated from the first time period, December 1983. Our search routine, the BHHH algorithm, finds that value of q which maximises $\log L$.

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