

# WORKING PAPERS IN ECONOMICS

Do Movements in the Forward Discount on the  
Australian Dollar Predict Movements in  
Domestic Interest Rates? Evidence from  
a Time Series Analysis of Covered Interest  
Parity in Australia in the late 1980s

by

C.I. Karfakis and A.J. Phipps

No. 187

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ABSTRACT

This paper examines covered interest parity using cointegration techniques on a daily data set for Australian dollar/US dollar spot and forward exchange rates and Australian and US interest rates. While the forward premium and the interest rate differential cointegrate in both the 3 month and 6 month markets, the data reject the formal restrictions of covered interest parity. Well defined error-correction mechanisms suggest that the forward premium bears the burden of adjustment to the long run cointegrating relationship and that past changes in the forward premium predict changes in the interest rate differential.

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*Do Movements in the Forward Discount on the Australian Dollar Predict Movements in Domestic Interest Rates? Evidence from a Time Series Analysis of Covered Interest Parity in Australia in the late 1980s<sup>1</sup>*

## I. INTRODUCTION

One of the major issues concerning Australian monetary policy in the deregulated environment of the late 1980s was the extent to which the Reserve Bank varied interest rates in order to stabilize the Australian dollar. This paper presents evidence which is consistent with the view that Australian financial markets at least anticipated that the Reserve Bank would vary interest rates to smooth the currency. The evidence emerges from a vector autoregressive (VAR) analysis of changes in the forward premium on the Australian dollar and changes in the interest rate differential on Australian and US bank accepted bills in the late 1980s. The VAR analysis supports strongly a short run dynamic model in which changes in the forward premium predict changes in the interest rate differential. This suggests that when the Australian dollar was under speculative pressure, as indicated by an increase in the forward discount, financial markets reacted by raising interest rates on Australian commercial paper (relative to US interest rates) in anticipation of a tightening of monetary policy by the Reserve Bank.

The VAR results mentioned above come from a cointegration analysis of covered interest parity (CIP) in Australia and the depiction of the associated short run dynamics by error correction (EC) VAR models. The CIP theorem is, of course, of interest in its own right. The interest rate parity theorem provides a theoretical framework linking domestic and foreign money markets and the spot and forward exchange markets. CIP is among the best known of conditions for an absence of pure arbitrage opportunities and, as such, is taken as evidence for market efficiency. Given the strong empirical support for CIP provided by Taylor (1987) using a high quality, high frequency data set and given the innate plausibility of the theorem itself, it is rather surprising that there has been a dearth of supporting empirical

<sup>1</sup>The authors are extremely grateful to the Reserve Bank and the Commonwealth Bank for the provision of the data used in this study and in particular to Hugh Harley and Gary Shilson-Josting for their friendly help and advice. Needless to say, they are in no way responsible for the conclusions expressed in this paper. Jack Towe provided sterling assistance with data preparation and computing and Ashok Parikh and Jeff Sheen offered helpful comments on an earlier draft.

evidence for CIP in Australia. Turnovsky and Bell (1983) provide weak support for CIP over the period 1974-1983, while the analysis of Chong (1987) supports CIP since the floating of the dollar. Neither study allows for the possibility of unit roots in the data generating processes for the forward premium and the interest rate differential with the consequence that the standard statistical procedures used to test for CIP may be invalid. Clearly there is a need to examine further the long run relationship between the forward premium on the Australian dollar and an appropriate interest rate differential and to test properly the restrictions implied by the CIP theorem. However, we believe that rejection of the exact restrictions of CIP need not be interpreted as grounds for rejecting the hypothesis that the Australian dollar forward market is efficient. Failure of CIP to hold exactly may be explained by the existence of adjustment costs and/or variable risk premia.

The purpose of the present study is threefold. First, we test for cointegration between the forward premium on the Australian dollar (in terms of the US dollar) and the interest rate differential on Australian and US bank accepted bills for both the 3 month and 6 month markets, using techniques developed by Johansen and Juselius (1990).<sup>2</sup> Second, we examine the CIP theorem by testing the implied coefficient restrictions both directly in the cointegrating vector by standard likelihood ratio tests and indirectly by the method suggested by Liu and Maddala (1992). Finally, the paper examines the information content of the cointegrating relationship by examining the short run dynamics implied by the associated EC models. In particular, we investigate whether changes in the interest rate differential between Australia and the US help predict movements in the forward premium/discount, and vice versa. The major innovation of this study is that it is based on daily data. This moderately high frequency data set allows the estimation of EC models in which adjustment may take place quite rapidly. There appears to be a general consensus that adjustment in financial markets is very rapid.

The rest of the paper is organised as follows. Section II, outlines theoretical and methodological issues. Section III, presents and discusses the empirical results. Concluding remarks are provided in Section IV.

<sup>2</sup> This paper is a refinement and extension of Karfakis and Phipps (1991). The major improvements in this study as compared with the previous one are: (i) the size of the sample is extended by two years; (ii) the cointegration tests are done using the maximum likelihood techniques of Johansen and Juselius (1990) rather than the approach of Engle and Granger (1987); (iii) we test the restrictions implied by CIP directly and (iv) we pay far more attention to the information content of the EC models associated with CIP.

## II. THEORETICAL AND METHODOLOGICAL ISSUES

For two countries with a high degree of international capital mobility, the CIP hypothesis may be tested by examining restrictions in the following equation

$$(F_{t+1} - S_t) / S_t = \alpha + \beta(r_t - r_t^*) + u_t \quad (1)$$

where  $F_{t+1}$  denotes the forward exchange rate contracted at time  $t$  for payment at time  $t+1$  and  $S_t$  denotes the spot exchange rate;  $r$  and  $r^*$  are domestic and foreign interest rates and  $u_t$  is the error term. The hypothesis of CIP implies that  $\alpha=0$ ,  $\beta=1$  and that  $u_t$  is white noise.<sup>3</sup>

CIP may be thought of as a long run equilibrium relationship and, as such, may be examined from the viewpoint of cointegration. Before applying cointegration testing procedures, it is customary to test the hypothesis that the individual series contain a unit root, that is, that they are integrated of order one (I(1)) since standard statistical inference may be invalidated if the data are nonstationary (Fuller, 1985). To this end, the Augmented Dickey-Fuller (ADF) test (Dickey and Fuller, 1981) may be used. The existence of a long run cointegrating relationship, like CIP above, may be tested by examining the stationarity of the process  $u_t$  (Engle and Granger, 1987).

An alternative way of modelling and testing equilibrium constraints postulated by economic theory is in the context of an EC model (Banerjee *et al*, 1986). The Granger representation theorem, presented by Granger and Engle (1987), states that if a set of variables are cointegrated there always exists at least one associated EC mechanism, and vice versa. Thus, if we establish that the forward premium on the Australian dollar (FP) and an associated interest differential (IDIF) are both I(1) and are cointegrated there exists an EC VAR specification of the form

$$\Delta FP_t = a_1 RES_{t-1} + \sum_{i=1}^n b_{1i} \Delta FP_{t-i} + \sum_{i=1}^n c_{1i} \Delta IDIF_{t-i} + \varepsilon_{1t}; \quad a_1 < 0 \quad (2)$$

$$\Delta IDIF_t = a_2 RES_{t-1} + \sum_{i=1}^n b_{2i} \Delta FP_{t-i} + \sum_{i=1}^n c_{2i} \Delta IDIF_{t-i} + \varepsilon_{2t} \quad (3)$$

where RES are the residuals from the cointegrating vector. The error terms  $\varepsilon_{1t}$  and  $\varepsilon_{2t}$

<sup>3</sup> It is well known that equation (1) is only an approximation. The condition for an absence of interest arbitrage, indeed for an absence of speculator and trader arbitrage as well, is

$$(1+r) = (1/S)(1+r^*)F$$

which reduces to

$$(F-S)/S = (r-r^*)/(1+r^*)$$

All the cointegration analysis, including estimation of the EC models, presented in this paper has been repeated for the more exact relationship with no changes to the major findings reported here.

are assumed to be white noise processes.

A cointegrating vector may thus be thought of as a long run equilibrium relationship which drives the short run dynamic EC mechanisms. Johansen and Juselius (1990) show how the cointegrating vector(s), the error correction parameters and the parameters of lagged differences of endogenous variables may be estimated using full information maximum likelihood techniques. Tests for the existence of any number of cointegrating (eigen)vectors are then based upon the number of significant associated eigenvalues, with the appropriate tables for inference being given in their paper. When one has found a system that has acceptable statistical properties, restrictions on the cointegrating vector(s) and the error correction parameters may be tested using standard likelihood ratio tests. If the forward premium and the associated interest rate differential turn out to be cointegrated, the existence of one or two associated EC mechanisms implies that there must be Granger causality running in at least one direction between the two variables and that therefore one variable can be used to forecast the other (Granger, 1986).

## III. EMPIRICAL ANALYSIS

### III.1. The Data

Our analysis involves the following empirical counterparts to the variables introduced in Section II:

- (i) the 3 month forward premium on the A\$ (denoted FP3) was calculated by expressing the forward margin on the A\$ in US cents in the Interbank Market as a percentage of the spot rate;
- (ii) the 6 month forward premium on the A\$ (FP6) was calculated in an analogous manner;
- (iii) the 3 month interest rate differential (IDIF3) was approximated by the difference between the yield on US 90 day bank bills and the yield (quoted to buyers) on Australian 90 day bank-accepted bills;
- (iv) the 6 month interest rate differential (IDIF6) was approximated analogously by the difference between US and Australian 180 day bank-accepted bill yields.

The data were available for five years from the beginning of January 1984 to the end of December 1988. The data are illustrated in Charts 1 and 2. One might expect there to have been a learning period associated with the newly introduced floating exchange rate regime, particularly since it was coupled with *perceived* monetary instability leading up to the abandonment of monetary targetting

in January 1985.<sup>4</sup> For this reason, we confined our analysis to the period from the beginning of February 1985 to the end of December 1988. The failure of CIP to hold in the earlier sub-period and the strong relationships thereafter are well illustrated in Charts 1 and 2.

CHART 1: FP3 AND IDIF3 (1/1/84 to 31/12/88)

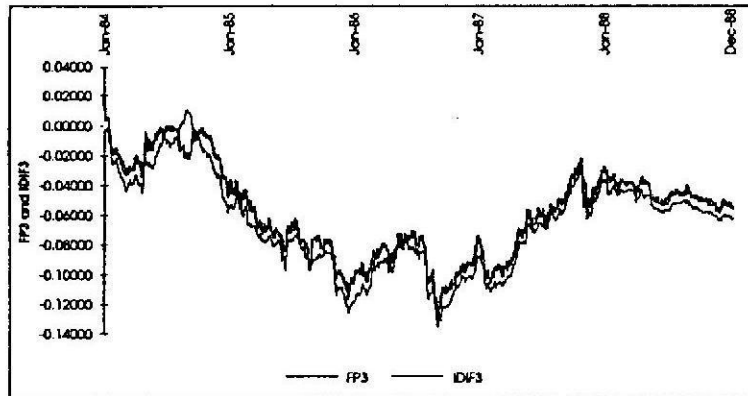
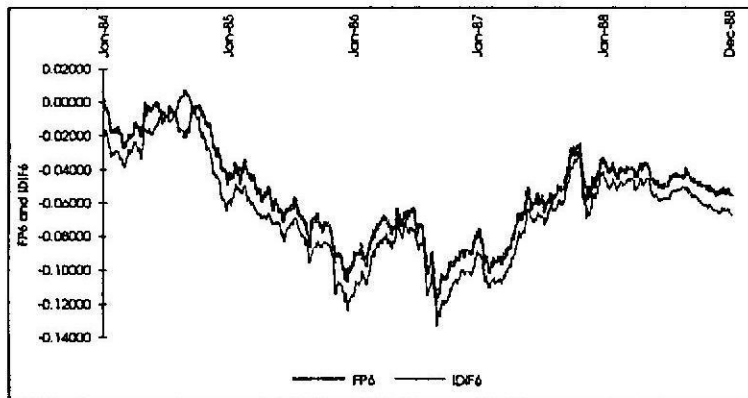


CHART 2: FP6 AND IDIF6 (1/1/84 to 31/12/88)



<sup>4</sup> From the beginning of January 1984 to the end of January 1985, FP3 and IDIF3 failed to cointegrate. Similar problems existed for the 6 month forward market. Indeed, from 20th August to 20th September 1984, the only time during the whole sample period that the interest rate differential favoured the US, the A\$ remained at a substantial forward discount.

Because the daily data excluded observations for each weekend and public holiday, we were faced with the choice of treating the weekend and holiday breaks either as identical to overnight breaks to produce a continuous data set or as substantially different and hence as missing observations. The main cost of treating breaks as missing observations is that incorporating longer lags into an estimated relationship requires a substantial reduction in the number of available observations and makes it difficult to compare estimates with different lags. For these reasons, we have opted in this study to work with a continuous data set.<sup>5</sup>

### III.2 Tests for Unit Roots and Cointegration

With respect to the univariate time series properties of the data, the results reported in Table 1 indicate that nonstationarity cannot be rejected for the levels of all the series at the 5% significance level.

TABLE 1: ADF TESTS FOR UNIT ROOTS

	<i>Variable</i>			
	FP3	IDIF3	FP6	IDIF6
	<i>Levels</i>			
<i>ADF Statistic</i>				
without trend	-2.4438	-1.8990	-2.1602	-1.6512
with trend	-2.1355	-1.6361	-1.8450	-1.4276
	<i>First Differences</i>			
without trend	-23.7378	-21.9540	-23.0583	-22.6670
with trend	-23.7837	-22.0048	-23.0843	-22.6908

Notes: (i) the ADF statistics were calculated with one lag of the dependent variable which was sufficient to ensure that the residuals were "white noise".

(ii) The number of observations for the ADF test on the levels was 1195 and for the test in first differences 1194. The critical values for ADF at the 95% level were -2.8644 (without trend) and -3.4160 (with trend).

In contrast, when the data are differenced, nonstationarity can be rejected in all cases. Hence, time-series modelling of CIP requires first-difference transformation of the variables to induce stationarity. Otherwise, by estimating the relationship in the levels one can obtain spurious outcomes (Granger and Newbold, 1974; Phillips, 1986).

<sup>5</sup> In our earlier study, Karfakis and Phipps (1991), we chose to treat the breaks as missing observations. Although the estimation method in that study was Engle and Granger rather than Johansen and Juselius, the estimated cointegrating vector was almost identical. The major differences between the two studies are, as one might expect, in the estimated short-run dynamics.

The results of the Johansen maximum likelihood procedures used to test for cointegration between the forward premium and the interest rate differential are set out in Table 2 (for the 3 month market) and in Table 3 (for the 6 month market). The number of lags in the VAR was set at 8 by reference to standard LR tests. We reject the null hypothesis of zero cointegrating vectors in favour of the alternative that there exists at most one such vector. This conclusion is supported by both the maximum eigenvalue and the trace tests described by Johansen and Juselius (1990).

**TABLE 2: JOHANSEN MAXIMUM LIKELIHOOD TESTS FOR COINTEGRATION;  
3 MONTH FORWARD PREMIUM AND INTEREST RATE DIFFERENTIAL**

List of variables included in the cointegrating vector: FP3      IDIF3      Intercept 945 observations from 253 to 1197. Maximum lag in VAR = 8.				
Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix				
Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	51.8012	15.6720	13.7520
r ≤ 1	r = 2	3.0782	9.2430	7.5250
Cointegration LR Test Based on Trace of the Stochastic Matrix				
Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r ≥ 1	54.8794	19.9640	17.8520
r ≤ 1	r = 2	3.0782	9.2430	7.5250
Estimated Cointegrated Vector(s) (Normalized in Brackets). Chosen r = 1.				
FP3	IDIF3	Intercept		
17.2666	-16.3579	-.065017		
(-1.0000)	(.94737)	(-.00397)		
Estimated Adjustment Matrix (Normalized in Brackets). Chosen r = 1.				
FP3	IDIF3			
-.011978	.003699			
(.20683)	(-.063869)			
Test of CIP Restrictions (Intercept = 0.00 Slope = 1.00) on Cointegrated Vector.				
LR Test of Restrictions	CHI-SQ( 2)=	48.9551	[.000]	

**TABLE 3: JOHANSEN MAXIMUM LIKELIHOOD TESTS FOR COINTEGRATION;  
6 MONTH FORWARD PREMIUM AND INTEREST RATE DIFFERENTIAL**

List of variables included in the cointegrating vector: FP6      IDIF6      Intercept 945 observations from 253 to 1197. Maximum lag in VAR = 8.				
Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix				
Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	33.6039	15.6720	13.7520
r ≤ 1	r = 2	3.7368	9.2430	7.5250
Cointegration LR Test Based on Trace of the Stochastic Matrix				
Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r ≥ 1	37.3406	19.9640	17.8520
r ≤ 1	r = 2	3.7368	9.2430	7.5250
Estimated Cointegrated Vector(s) (Normalized in Brackets). Chosen r = 1.				
FP6	IDIF6	Intercept		
16.7275	-15.4091	-.07977		
(-1.0000)	(.92118)	(.00477)		
Estimated Adjustment Matrix (Normalized in Brackets). Chosen r = 1.				
FP6	IDIF6			
-.00725	.004525			
(.12131)	(-.07569)			
Test of CIP Restrictions (Intercept = 0.00 Slope = 1.00) on Cointegrated Vector.				
LR Test of Restrictions	CHI-SQ( 2)=	32.5141	[.000]	

Normalising the cointegrating vectors on the 3 month and 6 month forward premiums yields two equations which reflect long run relationships suggestive of CIP:

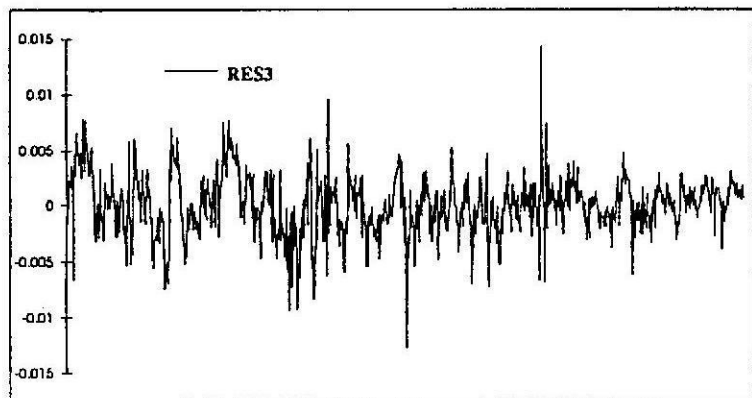
$$FP3 = 0.004 + 0.947*IDIF3 + RES3 \quad (4)$$

$$FP6 = 0.005 + 0.921*IDIF6 + RES6 \quad (5)$$

Graphs of the residuals from the cointegrating vector for the 3 month market are

presented in Chart 3.<sup>6</sup>

CHART 3: RESIDUALS FROM THE COINTEGRATING VECTOR (3 MONTH MARKET)



While the existence of a cointegrating relationship between the forward premium and the interest rate differential may be taken as evidence in favour of a long run relationship between the two such as CIP, it is important to test the exact restrictions implied by the model, namely that the slope coefficient is unity and that the intercept is zero. The likelihood ratio tests reported at the bottom of Tables 2 and 3 strongly reject such restrictions on the cointegrating vector.

An alternative way of testing the restrictions of CIP has been suggested by Liu and Maddala (1992). They propose a two part procedure involving an examination of the divergences from CIP ( $D = FP - IDIF$ ). A necessary condition for CIP to hold is that the divergences,  $D$ , should be  $I(0)$ . This condition is satisfied for both the 3 month and 6 month markets. The ADF(1) statistic for  $D3 (= FP3 - IDIF3)$  over our sample period is -21.97 which compares with a 95% critical value of -3.34 while the ADF(1) statistic for  $D6$  is -22.95. However, cointegration is not sufficient to establish the restriction that the slope coefficient is unity. It is also necessary that the divergences,  $D3$  and  $D6$ , be serially uncorrelated.<sup>7</sup> The data strongly reject the hypothesis that the  $D3$  divergences are serially uncorrelated. The LM statistic for 12

<sup>6</sup> A graph of the RES6 residuals, which is very similar to that of RES3, is presented in the Appendix.

<sup>7</sup> If  $D$  is  $I(0)$  and serially correlated, the unit slope coefficient and hence CIP are rejected. To see this rewrite the equation  $FP = \alpha + \beta IDIF + v$  as  $FP - IDIF = \alpha + (\beta - 1)IDIF + v$ . The presence of serial correlation in  $D (= FP - IDIF)$  implies it contains a permanent shock component i.e. it contains the same degree of persistence as  $IDIF$ , even though  $v$  is  $I(0)$ . That is, the presence of serial correlation in  $D$  implies  $\beta \neq 1$ .

lags in the autocorrelation function, which is distributed as  $\chi^2(12)$ , is 73.84 compared with a 95% critical value of 21.03. The corresponding LM statistic for  $D6$  is 155.12 which also leads to a rejection of the hypothesis of serially uncorrelated residuals. The Liu and Maddala tests reject the hypothesis that the slope coefficient is unity for both the 3 month and 6 month markets, confirming the results from the direct tests of the restrictions in the cointegrating vectors presented in Tables 2 and 3. In short, CIP appears not to hold exactly.

However, we believe that rejection of the exact restrictions of CIP need not be interpreted as grounds for rejecting the hypothesis that the Australian dollar forward market is efficient. Failure of CIP to hold exactly may be explained by the existence of adjustment costs and/or of variable risk premia. The existence of transactions costs implies that there is a band of values over which arbitrage conditions need not hold. Frenkel and Levich (1975) analysed the transactions cost band and concluded that, once such costs were taken into account, the empirical data were consistent with CIP.

An alternative explanation for the failure of CIP to hold exactly in our estimates may be found in the existence of variable risk premia. There appear to be substantial and variable risk premia on private paper<sup>8</sup> (Bank Accepted Bills in our study) relative to government paper (Treasury Notes in Australia and Treasury Bills in the US). The mean premium in proportionate terms on 3 month private paper in the US based on monthly data for our sample period was 0.108 (10.8%) while the mean premium in Australia was 0.066 (6.6%). In spite of this, the premium in absolute terms was higher for Australia than for the US because Australian interest rates were substantially higher for the whole of the sample period. As a result, the US-Australia Bank Bill interest rate differential was larger in absolute terms, on average and for most of the sample period, than the US-Australia Treasury Bill/Note interest rate differential. It may be argued that most interest arbitrage takes place in relatively risk-free Treasury Bills and Notes and hence that the appropriate interest rate differential for CIP is that between the rate on US Treasury Bills and the rate on Australian Treasury Notes. The interest rate differential on Bank Accepted Bills overstates the interest rate differential on Treasury Bills/Notes. This may well account for the findings reported in this paper.

### III.3 Error Correction Models

Having established that the forward premium and the interest rate differential cointegrate for both the 3 month and 6 month markets, it is appropriate to examine the associated EC mechanisms which describe the short run dynamics. Thus we estimate equations of the following form

<sup>8</sup> Evidence for such a risk premium in Australia is provided by McDonald, Kendall and Ridley (1993)



$$\Delta FP3_t = a_1 RES3_{t-1} + \sum_{i=1}^7 b_{1i} \Delta FP3_{t-i} + \sum_{i=1}^7 c_{1i} \Delta IDIF3_{t-i} + \varepsilon_{1t}; \quad a_1 < 0 \quad (6a)$$

$$\text{and } \Delta IDIF3_t = a_2 RES3_{t-1} + \sum_{i=1}^7 b_{2i} \Delta FP3_{t-i} + \sum_{i=1}^7 c_{2i} \Delta IDIF3_{t-i} + \varepsilon_{2t} \quad (7a)$$

where RES3 are the residuals from the cointegrating equation for the 3 month market. Corresponding equations (6b and 7b) may be estimated for  $\Delta FP6$  and  $\Delta IDIF6$ . The results are reported in Table 4.

TABLE 4: ESTIMATES OF THE EC MODELS AND GRANGER- CAUSALITY TESTS

Eq. No.	Dependent Variable	Coefficient Estimates			Summary Statistics		
		$a_1$	$\sum_{i=1}^7 b_{ki}$	$\sum_{i=1}^7 c_{ki}$	$R_{bar}^2$	SEE	Q(90) [MSL]
6a	$\Delta FP3$	-0.1987* (5.3976)	0.3951 (1.8595)	-0.0590 (0.2789)	0.053	0.002	75.89 [.855]
7a	$\Delta IDIF3$	0.0553* (1.8866)	0.9979* (5.9018)	-0.05844* (3.4714)	0.296	0.002	72.42 [.913]
6b	$\Delta FP6$	-0.1157* (3.5595)	0.1320 (0.6102)	0.1165 (0.5732)	0.051	0.002	85.26 [.622]
7b	$\Delta IDIF6$	0.0701* (2.5914)	1.1860* (6.5821)	-0.6480* (3.8275)	0.285	0.002	94.48 [.353]
Granger - Causality Tests							
Eq. No.	Dependent Variable	$H_0: b_{k1} = b_{k2} = \dots = b_{k7} = 0$		$H_0: c_{k1} = c_{k2} = \dots = c_{k7} = 0$			
		F-statistic	[MSL]	F-statistic	[MSL]		
6a	$\Delta FP3$	2.6901	[0.0092]	1.11733	[0.3153]		
7a	$\Delta IDIF3$	19.0654	[0.0000]	3.8326	[0.0004]		
6b	$\Delta FP6$	3.8276	[0.0004]	1.6656	[0.1139]		
7b	$\Delta IDIF6$	24.7476	[0.0000]	5.76750	[0.0000]		

Notes: (i) t-statistics are in rounded brackets below the estimated coefficients.

(ii) \* indicates coefficient is significant at 5% level.

(iii) The marginal significance levels [MSL] for the Ljung-Box Q-statistics (estimated with 90 degrees of freedom) and for the F-statistics in the Granger-causality tests are in square brackets.

(iv) The number of lags used was limited to 7 because LR tests rejected the inclusion of 8 lags in favour of the alternative of 7 lags.

The EC terms ( $a_{ki}$ ) are significant in the (3 and 6 month) forward premium equations and the (3 and 6 month) interest rate differential equations indicating the existence of forces in the (3 and 6 month) forward exchange and bank accepted bill markets that operate to restore long run equilibrium after a short run disturbance. The EC terms have the correct signs and indicate that deviations of the forward premium from its equilibrium value were corrected at a daily rate of 25% (19.87 + 5.53) for the 3 month contract, and 17% (11.57 + 7.01) for the 6 month contract. These estimated rates of adjustment are rather slower than we anticipated (but they are consistent with estimated adjustment matrices obtained by the Johansen and Juselius procedure and reported in Tables 2 and 3). Furthermore, the strong EC mechanisms are inconsistent with the hypothesis of market efficiency because information on changes in the interest rate differential was not incorporated immediately into the forward premium. This apparent market failure may again be accounted for by the existence of transaction costs. However, consistent with prior expectations, both the size and significance of the EC terms in the forward premium equations relative to their size and significance in the interest rate differential equations indicate that most of the burden of adjustment to long run equilibrium is borne by the 3 and 6 month forward premiums.

Of more interest are the results of the Granger-causality tests. Standard F tests allow us to reject strongly the null hypothesis that the coefficients on the lagged forward premium terms and lagged interest rate differential terms are zero in the 3 and 6 month interest rate differential equations, while we are unable to reject the hypothesis that the coefficients on those terms are zero in the 3 and 6 month forward premium equations. This indicates that, over our sample period, changes in the forward premium Granger-caused changes in the corresponding interest rate differential rather than the other way round. Hence, changes in the forward premium could at that time have been used to predict changes in the corresponding interest rate differential. The  $R_{bar}^2$  for both interest rate differential equations is about 30% indicating that short term interest rate movements are essentially predictable. We believe that the most likely explanation for these findings is that the money markets anticipated a well-defined Reserve Bank policy reaction to speculative attacks on the Australian dollar in this period. Thus, it seems likely that when the Australian dollar was under speculative pressure, as indicated by an increase in the forward discount, financial markets reacted by raising interest rates on Australian commercial paper (relative to US interest rates) in anticipation of a tightening of monetary policy by the Reserve Bank. The significant, negative coefficients on the lagged interest rate differential terms in the interest rate differential equations indicate that generally the market may have overreacted and corrected for this by reducing interest rates

subsequently. In general, these results are consistent with the view that Australian financial markets anticipated that the Reserve Bank would vary interest rates to smooth the currency. Of course, there may be other less obvious explanations for these results.

#### IV. CONCLUSIONS

This paper has examined the covered interest parity (CIP) hypothesis using cointegration techniques on a daily data set for Australian dollar/US dollar spot and forward exchange rates and Australian and US interest rates. Cointegration between the forward premium and the interest rate differential in both the 3 month and 6 month markets establishes CIP as a possible long run equilibrium relationship for the sample period. However, formal tests of the restrictions implied by CIP, both directly by likelihood ratio tests on the estimated parameters of the cointegrating vector and indirectly by the method suggested by Maddala and Liu (1992), lead us to reject CIP in its exact form. This and well-defined EC mechanisms suggest that the forward exchange market did not utilise all available information efficiently in the short run. However, we believe that rejection of the exact restrictions of CIP need not be interpreted as grounds for rejecting the hypothesis that the Australian dollar forward market is efficient. Failure of CIP to hold exactly may be explained by the existence of adjustment costs and/or variable risk premia.

The fact that changes in the forward premium Granger-caused changes in the corresponding interest rate differential rather than the other way round over our sample period implies that changes in the forward premium could, at that time, have been used to predict changes in the corresponding interest rate differential. We believe that the most likely explanation of these findings is that Australian financial markets anticipated that the Reserve Bank would vary interest rates to smooth the currency.

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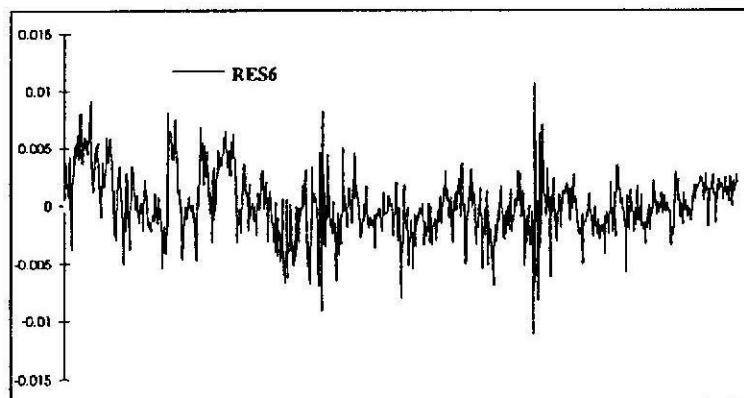
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## APPENDIX

CHART 4: RESIDUALS FROM THE COINTEGRATING VECTOR (6 MONTH MARKET)



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