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Financial Barriers in the Pacific Basin: 1982-1992

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Abstract

Interest rate links strengthened among some Pacific Rim countries over the period 1982-1992, even though substantial country barriers remain. The covered interest differential narrowed for Australia and New Zealand, as their programs of financial liberalization admitted them to the club whose members already included Canada, Hong Kong, Japan and Singapore. The covered interest differential actually increased slightly over certain subperiods in Japan and Canada, but the economic magnitudes were quite small, and are probably due to other factors besides the presence of capital barriers.

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I. Introduction

This paper seeks to investigate the extent to which Pacific Rim financial markets are becoming more tightly linked, by analyzing interest rate co-movements adjusted by forward discounts. Complete financial integration as defined by the covered interest parity (CIP) condition would be an important phenomenon for many reasons.[1] It might mean, for example, that monetary authorities in small countries had lost the ability to affect the exchange rate, independent of monetary policies (except perhaps through portfolio effects). Earlier studies of capital mobility in the Pacific context, including Glick [1987], Glick and Hutchison [1990], and the papers in Cheng [1988], have usually not directed their attention to the CIP criterion since forward markets have not been well developed in the much of the region. Here we focus on a basic question: As the result of financial liberalization and innovation, particularly the removal of capital controls and other barriers to international capital mobility, are interest rates in Pacific countries becoming increasingly linked to the core world financial market?[2]

Figure 1 illustrates one finding of this paper: that capital mobility, has indeed increased *in the aggregate* over the 1982-92 period for the countries examined: Australia, Canada, Hong Kong, Japan, Malaysia, New Zealand and Singapore. The mean covered interest differential has moved from minus one percent to roughly zero (on an annualized basis), while the standard deviation of interest differentials shrank to near zero.[3]

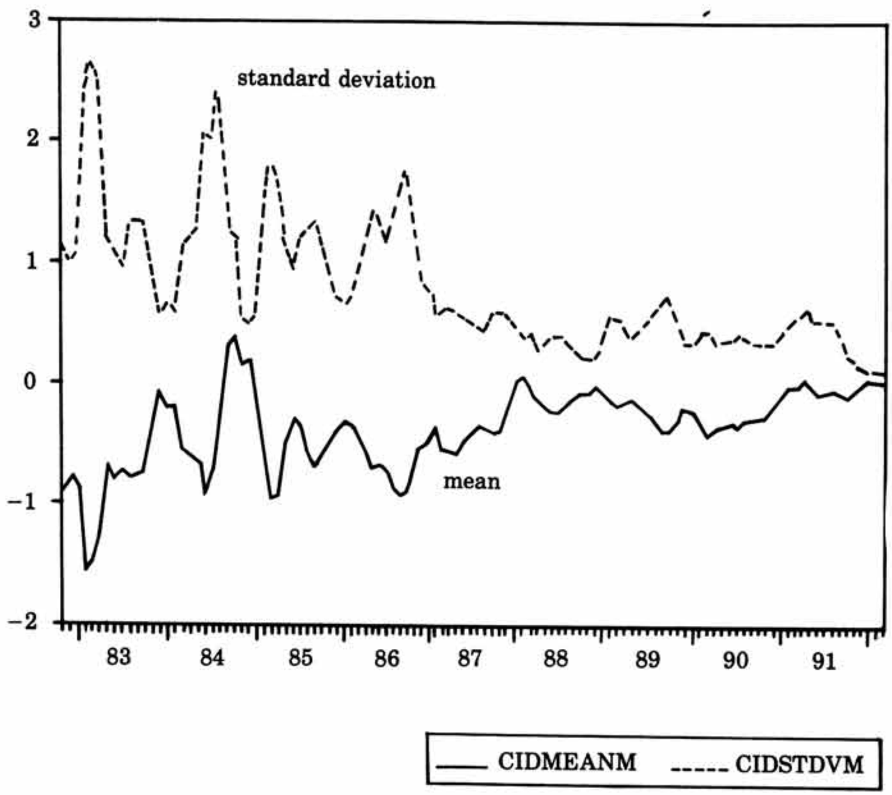
This graph is merely suggestive, because summary statistics can be misleading. Consequently, this paper examines the linkage in a variety of ways. In Section II, we estimate the trends in the absolute values of the covered interest differential. In Section III we run covered interest parity regres-

1. Frankel [1991]refers to the condition of CIP as "perfect capital mobility", to differentiate it from uncovered interest parity (UIP), or "perfect capital substitutability." When CIP holds, then capital barriers, or the threat thereof, are absent.

2. In another paper ⟨Chinn and Frankel [forthcoming]⟩, we examine both this issue (what we call "financial links"), as well as the role of "currency links" (the absence of an exchange risk premium, and the stability of the exchange rate). In the event that both types of links are strong, interest rates will be equated internationally.

3. A 3 month moving average series. Malaysia drops out of the sample at 1990:04, which tends to bias downwards both the mean and standard deviation.

Figure 1
Mean and Std. Dev. of the Covered Interest Differential



sions, taking into account issues of unit roots and cointegration. We also allow for time variation in the parameters, as a proxy for changes in the degree of capital mobility. We do this using both recursive and rolling regression techniques. Finally, using the cointegration results obtained in Section III, we adopt a vector autoregression technique (Section IV) to determine how much foreign market factors impact upon local rates.

Our key findings are as follows: Some of the interest rates appear to be increasingly influenced by covered interest rates, indicating increasing financial integration for some countries (Australia and New Zealand, most obviously). For certain other countries, the level of financial integration was high at both the beginning and end of the sample period, so that no trends were discernible (*e.g.*, Hong Kong and Singapore). Finally, for Canada and

Japan the observed covered interest differential is increasing, but likely for reasons besides capital barriers.

II. Absolute Covered Interest Rate Differentials

Frankel [1991] found that covered interest differentials, *vis-a-vis* the Eurodollar rate during the period September 1982 to January 1988, on average were as small for Hong Kong, Singapore, Japan and Canada as for the financially most open European countries, and that they were bigger and more variable for Malaysia, Australia and New Zealand. These results pertained to the 1982-88 period.

We now turn to the question how the covered interest rate behavior has changed over the period 1982:09-1992:03. Table 1 shows the results of regressing the absolute covered interest differential against a constant and a linear time trend, over the ten-year period.[4] As in the earlier study, there are only seven currencies covered, since forward markets exist only for these.

Australia and New Zealand show diminishing covered interest differentials, as one would expect from their programs of financial liberalization in the 1980s. The trends for the other countries are not statistically significant.[5] The Australian results appear to be dominated by a spike in early 1983 (due to a large forward discount of the Australian dollar). We examine the covered interest differentials more closely below.

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4. All data are monthly. The spot and 3 month forward rate data are dated contemporaneously with the interest rates, and are drawn from DRI's Financial and Credit Statistics database, and the IMF's *International Financial Statistics* database. Details are provided in the Data Appendix.

5. Except for the statistically significant positive trend for the case of Japan, an unexpected result in light of past findings by Otani [1983], Frankel [1984] and others that the large differentials of the 1970s fell to zero in the early 1980s. The results have been cross-checked using the Japanese EuroYen rate. This series seems to obey covered interest parity better, although there is still a rise in the differential to about 0.60% in mid-1989 even in this series. One possible explanation is that the Gensaki was a better "representative rate" earlier in the 1980s, and as it has decreased in importance ⟨Feldman [1986, p. 42], Bonser-Neal and Roley [1993]⟩, certificates of deposit (CDs) have become more representative.

Table 1

Regressions of Absolute Covered Differential on Trend

Currency	Constant	Trend	$\bar{R}^2$	DW	N	Sample
Australia	1.572* (0.243) [0.759]	–0.190* (0.046) [0.080]	.12	1.306	115	
Canada	0.121 (0.038) [0.113]	0.020 (0.007) [0.012]	.06	1.021	115	
Hong Kong	0.100 (0.058) [0.174]	0.028 (0.011) [0.018]	.05	1.331	115	
Japan	0.142 (0.034) [0.103]	0.027* (0.006) [0.011]	.14	0.761	115	
Malaysia	1.460* (0.240) [0.720]	–0.040 (0.056) [0.096]	–.01	0.621	87	82:09-90:03
New Zealand	3.165** (0.272) [0.816]	–0.388** (0.050) [0.086]	.35	1.920	115	
Singapore	0.269 (0.055) [0.164]	0.028 (0.010) [0.018]	.05	1.082	112	82:09-91:12

Notes: All parameter estimates in percent terms. Trend terms are "annualized". Figures in parentheses (.) are asymptotic standard errors; figures in brackets [.] are standard errors assuming N/3 independent observations.
*(**) indicates significance at 5% (1%) level using the adjusted standard errors.

III. Covered Interest Parity Regressions

We consider the following regression:[6]

6. An alternative regression, with $(i + fd)$ as a regressand is defensible on the grounds that investor perceptions of possible exchange rate changes (captured by fd) are endogeneous. This regression specification suffers from the drawback that the regressand is not clearly the variable of interest. When the covered local rate is the regressand, the point estimates for Canada, Hong Kong, Japan and Singapore are

$$i_t = \beta_0 + \beta_1(i_t^{US} - fd_t) + u_t \tag{1}$$

where:

$fd_t \equiv f_t - s_t$; the forward discount, in US$ / foreign currency unit.
u_t; a regression residual

Before the regressions are implemented, it is necessary to determine whether the data generating processes are such that OLS yields proper estimates of the parameters. Hence we test for stationarity and cointegration, and specify the regressions in light of the results, as suggested by Chinn [1991]. All the raw data series appear to be integrated of order one, using the standard Augmented Dickey-Fuller tests (results not reported). In Table 2, standard Engle-Granger cointegration test results are reported; according to the results in the first two columns, neither the interest differentials, nor the forward discount appear to be stationary.[7] This latter result is in itself interesting, as it is at variance with a large body of literature which finds cointegration between spot and forward rates.[8] As Brenner and Kroner [1992] point out, there is no reason to believe that spot and forward rates are cointegrated unless interest rates are. Moreover, there is no evidence in column 3 that all four variables are cointegrated. However, since it is well known that tests of cointegration have notoriously low power, we use a more powerful test of cointegration based on the discussion in Kremers, Ericsson and Dolado [1992]. They show that standard asymptotic theory can be used when posing the test in the context of an error correction model; specifically, the *t*-statistic on the error correction term coefficient (γ_7) has the usual distribution.

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unchanged. The point estimates for Australia and Malaysia are higher, but so too are the standard errors (by a factor of 4 in the latter case). This result arises because $(i + fd)$ is much more variable than i^{US} in these two cases. Consequently, we focus on the regression results reported below.

7. While these terms do not exactly conform to the regression variables, they are of interest because they have ready economic interpretations, and have been the object of study elsewhere. The cointegration results on i versus $i^{US} - fd$ are not appreciably different (*i.e.*, they both appear to contain unit roots).

8. See Baillie and Bollerslev [1989], Copeland [1991], and Hakkio and Rush [1989]. These studies pertain mostly to European countries *vis a vis* the US.

Table 2
Tests of Cointegration

Test Stat Coint.Vec	EG t-stat (i, i^{US})	EG t-stat (f, s)	EG t-stat (i, i^*, f, s)	ECM t-stat (i, i^*, f, s)
AU	–2.959	–2.402	–2.387	–1.272[a]
AU[c]	–2.604	–2.460	–2.324	–11.462***[b]
CN	–2.275	–2.696	–3.175	–1.647*[a, b]
HK	–1.959	–2.502	–3.575	–9.336***
JP	–2.780	–1.958	–3.316	–2.435**
MA	–1.691	–1.483	–1.669	–2.585**[a]
NZ	–1.468	–1.282	–1.720	–4.440***[a]
SI	–1.500	–1.923	–1.737	–3.662***[a]

Notes: The figures are t-statistics. EG denotes statistics from Engle-Granger cointegrating regressions, while ECM denotes t-statistics on the error correction based test. The critical t-values for the EG t-statistics are (for a two variable system) 3.086(10%); 3.399(5%); 4.010(1%); (for a four variable system) 3.893(10%); 4.207(5%); 4.824(1%) (for 12 lags and a constant).

*(**) [***] denotes statistical significance at the 10%(5%) [1%] level.

a. Heteroskedasticity at the 10% level or greater; t-statistics calculated using White robust standard errors.

b. A single lag of the dependent variable is included.

c. Sample period is 1983:03-1992:03.

$$\Delta i_t = \gamma_0 + \gamma_1 \Delta i_t^{US} + \gamma_2 \Delta i_{t-1}^{US} + \gamma_3 \Delta f_t + \gamma_4 \Delta f_{t-1} + \gamma_5 \Delta s_t + \gamma_6 \Delta s_{t-1} + \gamma_7 ECT_{t-1} \quad (2)$$

In this model, the candidate error correction term (ECT) is quite obvious – the covered interest differential. The statistics in column 4 are the relevant t-statistics; they indicate that the null hypothesis of no cointegration can be rejected in all cases except the Australian. This result appears to be dominated by the spike in February 1983, associated with an expected devaluation. Starting the sample in 1983:03 yields evidence of cointegration. Hence, estimation can proceed, although one might wish to consider estimates for Australia in first differences.

Table 3 reports the results of simple OLS regressions in levels for seven currencies: Australian, Canadian, Hong Kong, Malaysian, New Zealand, Sin-

Table 3
Covered Interest Parity Regressions

$$i = \beta_0 + \beta_1 (i^{US} - fd) + u$$

Currency	$\hat{\beta}_0$	$\hat{\beta}_1$[a]	$\bar{R}^2$	DW	Q-stat. ARMA Spec	Breaks
Australia	3.396* (0.499) [1.497]	0.707 (0.036) [0.062]	.77	0.825	119.56** AR1	1985
Canada	0.399 (0.137) [0.411]	0.943* (0.014) [0.024]	.98	0.851	223.70** AR2	1987, 1990
Hong Kong	0.376 (0.065) [0.195]	0.968* (0.008) [0.014]	.99	1.957	14.87	1983:08-11
Japan	0.360 (0.132) [0.396]	0.919* (0.023) [0.039]	.93	0.394	694.17** ARMA 1, 1	1986:02-3, 1991:02
Malay.[b]	0.860 (0.258) [0.774]	0.742** (0.029) [0.050]	.88	0.862	71**/14.8 AR1/—	1984:09, 1985:04,
New Zealand	2.997* (0.477) [1.431]	0.741** (0.029) [0.050]	.85	1.269	67.49** AR3	1983:07, 1984:08
Singapore	0.023 (0.117) [0.351]	0.935 (0.019) [0.033]	.96	0.911	186.63** AR3	1987:02, 1990:04

Notes: All parameter estimates in percent terms. Figures in parentheses (.) are asymptotic standard errors; figures in brackets [.] are standard errors assuming N/3 independent observations. Q-statistic indicates the Ljung-Box Q-statistic for lag order 13. "ARMA Spec." indicates the apparent ARMA specification for the residuals. "Break(s)" indicates likely breaks as indicated by a 1-step ahead recursive residuals test. *(**) indicates significance at 5% (1%) level using the adjusted standard errors.

a. The significance levels for the slope coefficient is for H_0: $\beta = 1$. Note that all the slope parameter estimates are statistically different from zero.

b. There are missing data for the period 1987:07-87:10. The figures to the left (right) of the / are for the first (second) subperiods.

gapore dollars and the Japanese Yen. The $\hat{\beta}_1$ coefficients are all of the correct positive sign and at a high degree of statistical significance. The CIP proposition is often posed as the joint null hypothesis H_0: $\beta_0 = 0$ and $\beta_1 = 1$. However, when the variables involved are integrated, then the appropriate null is $\beta_1 = 1$.[9] The four that have coefficients close to one are, as one would expect, the four that are known to have had open financial markets throughout the sample period: Canada, Hong Kong, Japan, and Singapore.[10]

We chose to examine each currency using time-varying parameter techniques, specifically recursive and rolling regressions of equation 1.[11] In Table 4, the slope coefficients at the end of 1984:08,[12] and at the end of the sample, 1992:03, are reported. The pattern of changes substantiates one's priors. In every case the recursive coefficient is higher at the end of the period than 1984:08. Australia and New Zealand, countries that undertook extensive programs of financial liberalization in the 1980s, are the two where the increase in substantial.[13] In the cases of Canada, Hong Kong, and Singapore the coefficients were already high at the beginning of the sample

9. When the processes are characterized as diffusion processes, then the constant is a function of the forecast horizon and the parameters of the Brownian motion processes – the mean and instantaneous standard deviations ⟨Brenner and Kroner [1992]⟩.

10. Even the coefficients on these are statistically significantly less than one. (The standard errors are very small.) A small amount of fluctuation in the covered interest differential within a narrow range dictated by the bid-ask spread or other transactions costs could give this result.

11. Recursive regressions increment the sample size recursively. Rolling regressions hold the sample size constant. Other, more complicated time-varying parameter approaches are available. A Kalman-filter approach as in Wolff [1987] is feasible, but is restrictive in forcing the evolution of the coefficients to follow a random walk. The stationary coefficients approach of Schinasi and Swamy [1990] is a possible alternative, and has some positive features, but is observationally equivalent to a regression equation with a time-varying intercept and heteroskedasticity.

12. The recursive coefficient series start in 1982:11, but it takes a while for the series to settle down to reasonable estimates with some degrees of freedom. At 1984:08, the degree of freedom is still only 22.

13. And statistically significant; the final estimate is outside the initial (1984:08) 2 standard error band. This could almost be said for Japan, but the initial standard error band is large. Moreover, these are unadjusted standard errors; presumably adjusting for serial correlation (which is very high in the Japan case) would make the increase even less apparently significant.

Table 4
Time-Varying Slope Coefficients from CIP Regressions

Currency	Procedure	$\hat{\beta}$ (1984:08)	$\hat{\beta}$ (1992:03)
Australia	recursive rolling	0.287 0.287	0.707 0.965
Canada	recursive rolling	0.916 0.916	0.943 0.887
Hong Kong	recursive rolling	0.974 0.974	0.968 1.011
Japan	recursive rolling	0.572 0.572	0.919 0.867
Malaysia	recursive rolling	0.707 0.707	0.742[a] 0.550[a]
New Zealand	recursive rolling	0.305 0.305	0.741 0.939
Singapore	recursive rolling	0.730 0.730	0.935[b] 0.865[b]

Notes: Rolling regressions use window size of 24 months.
a. Missing data for 1987:07-10, sample ends 1990:03.
b. Sample ends 1991:12.

period, so it is not surprising that the increase is small. Malaysia exhibits lack of integration with international markets. However, the sample ends somewhat earlier than for the others.

Several tests can use recursive techniques to assess structural breaks. We used the 1-step-ahead forecast *F*-test to find possible break points. The 1-step-ahead *F*-test is sometimes illuminating in itself. The break in 1984 for New Zealand fits nicely with the devaluation of the New Zealand dollar, and the breaks in 1985, with the abolition of exchange controls in Dec. 1984, and floating in March 1985 ⟨Coats [1988, p. 91]⟩. Hong Kong's end-of-1983 break matches the beginning of the pegged exchange rate for the Hong Kong dollar.

Rolling regressions might be deemed more important if one wishes to discount older information. These estimates are not substantially different, although they tend to move around more. An exception is Malaysia, where it appears that integration has lessened.

IV. Variance Decompositions from VARs

Another measure of how much international interest rates (appropriately covered) influence domestic interest rates can be obtained by segmenting the sample into early (1982:09-87:06) and late (1987:07-92:03) subsamples, and comparing the relationships implied by the results of vector autoregression (VAR) techniques. Specifically, the resultant decompositions of the variance of innovations into that attributable to own-variable shocks, and that to shocks of the other variable(s), can be then used as measures of influence. The process is implemented in two forms: the simplest bivariate VAR and a more complex multivariate VAR.

A. Multivariate VAR

We first adopt an agnostic approach, by allowing all four integrated variables to enter into the VAR separately. Given the results on cointegration in the previous section, we know that this is a valid procedure. Hence, the estimated system can be written compactly:

$$y_t = \sum_{i=4}^{4} B_i y_{t-i} + \varepsilon_t \tag{3}$$

where:

y is a $(n \times 1)$ vector of endogenous variables (i^{US}, i, s, f),
B is a $(n \times n)$ matrix of parameters,
ε is a $(n \times 1)$ vector of random error terms distributed $(0, \Omega)$.

This system is estimated for the six countries for which data are available. Before discussing the results, some technical complications must be addressed. The most critical of these is the ordering of the VARs. All the variables are treated as endogenous, but in calculating the impulse response functions and the variance decompositions, one must attribute some proportion of the variance of errors in the two equations to either one or the other equations (assuming the errors are correlated). This attribution is usually accomplished by allocating that variance to the first variable in the VAR ordering.[14]

14. This procedure is the standard Choleski decomposition of the covariance matrix, such that the innovations covariance matrix is lower triangular. *In general*, this is a quite arbitrary manner in which to do the attribution, and is consistent with assum-

In this analysis, we adopt the following ordering: US Eurorate, onshore rate, the spot rate, the forward rate. Since all of these economies, save the Japanese are small relative to the US, the location of the i^{US} variable is uncontroversial.[15] By placing the local interest rate in the second position, we weight the case towards local influences dominating, as one can argue that the other two variables are equally endogenous.

Estimation of the individual VAR equations for these countries yields fairly good results. Four lags of the endogenous variables (given that these are financial variables) appear to soak up any residual serial correlation. The impulse-response functions are reasonable for most of the systems.[16] The results for the first three months of the variance decompositions for the respective local rates are provided in Table 5. More periods can easily be calculated, but since we are presumably interested in rapidly adjusting financial variables, one quarter should be sufficient to make inferences.

The numbers reported are the proportions of total innovation in the local rate which can be accounted for by all other variables (i^{US}, s, f). One finds that for Australia and New Zealand, this proportion has risen. For Hong Kong and Singapore the proportion has remained fairly stable, while Canada shows a decline. This might result from the increased political risk associated with the unity question in Canada 1990. In fact, a deviation from CIP did begin with the breakdown of the Meech Lake Accords process in June 1990.

B. Bivariate VAR

Vector autoregression techniques are known to be sensitive to particular specifications, along several dimensions. Hence, we impose at this point priors that the covered US interest rate and the local rate are cointegrated. Then the estimated system is expressed most simply as:

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ing a recursive system where the first variable is exogenous with respect to the second.

15. Except for the case of Japan, where the issue of which interest rate was more exogenous is very difficult. For that reason, the Japan case is omitted from this portion of the paper.

16. An exception is Malaysia, where only 26 observations are available in the later subsample and the resultant impulse-response functions are nonsensical.

Table 5
Proportion of Innovation Variance in Local Rates Attributable to Non-Local Variables (Four-variable System)

	Per.	Early		Late	
		Proportion.	Std. Error	Proportion	Std. Error
AU	1	1.69%	(0.78)	8.11%	(0.35)
	2	7.98	(1.21)	25.02	(0.55)
	3	13.18	(1.37)	31.29	(0.75)
CN	1	28.02	(0.46)	26.06	(0.28)
	2	43.07	(0.58)	29.77	(0.45)
	3	54.56	(0.66)	30.51	(0.59)
HK	1	0.10	(1.11)	5.42	(0.58)
	2	10.85	(1.27)	10.53	(0.87)
	3	34.96	(1.49)	9.46	(0.99)
MA	1	22.85	(0.97)	na	
	2	46.28	(1.25)	na	
	3	51.28	(1.37)	na	
NZ	1	17.66	(1.69)	4.80	(0.44)
	2	20.50	(2.18)	25.16	(0.59)
	3	11.99	(2.31)	19.58	(0.71)
SI	1	23.38	(0.40)	3.04	(0.37)
	2	41.91	(0.56)	14.94	(0.55)
	3	49.26	(0.64)	43.27	(0.73)

Notes: The number under proportion is 100% minus the percent of total innovation variance in the local interest rate attributable to itself, assuming the following ordering (i^{US}, i, s, f). Std. error is the asymptotic standard error.

$$i_t = \gamma_0 + \sum_{i=1}^{4} \gamma_i i_{t-i} + \sum_{j=1}^{4} \delta_j (i^{US} - fd)_{t-j} + v_t \quad (4a)$$

$$(i^{US} - fd)_t = \theta_0 + \sum_{i=1}^{4} \theta_i \, i_{t-i} + \sum_{j=1}^{4} \phi_j (i^{US} - fd)_{t-j} + w_t \quad (4b)$$

In this analysis, we assume that the covered US interest rate is exogenous with respect to the local rate, and enter it first into the VAR, under the logic that the US capital market is large relative to those to any of the local mar-

Table 6

Proportion or Innovation Variance in Local Rates Attributable to the Covered US Rate (Bivariate System)

	Per.	Early		Late	
		Proportion.	Std. Error	Proportion	Std. Error
AU	1	58.03	(0.93)	85.47%	(0.42)
	2	70.28	(1.45)	88.07	(0.62)
	3	69.52	(1.63)	83.71	(0.87)
CN	1	93.71	(0.54)	65.42	(0.29)
	2	92.20	(0.75)	72.04	(0.47)
	3	85.95	(0.88)	71.70	(0.63)
HK	1	97.52	(1.40)	92.78	(0.65)
	2	98.10	(1.74)	96.28	(1.01)
	3	96.51	(1.87)	94.48	(1.21)
MA	1	69.01	(1.35)	na	
	2	78.60	(1.62)	na	
	3	83.17	(1.85)	na	
NZ	1	48.55	(1.96)	86.90	(0.56)
	2	48.44	(2.72)	84.21	(0.68)
	3	49.67	(3.11)	75.93	(0.81)
SI[a]	1	79.46	(0.48)	11.07	(0.41)
	2	87.97	(0.66)	24.76	(0.60)
	3	90.00	(0.76)	51.06	(0.76)

Notes: The number under proportion is the percent of total innovation variance in the local interest rate attributable to the covered US rate, assuming the latter is exogenous with respect to the former. The std. error is the asymptotic standard error.

a. The VAR system in the late subsample is of lag order 6.

kets. It is also consistent with the assumptions implicit in running the CIP regressions with the covered US rate on the right-hand-side.

The results in Table 6 are largely in line with earlier results. Australia and New Zealand show substantial increases in the influence of the US interest rate after 1987, while Hong Kong shows relatively constant effects. One surprise is Singapore, which in contrast to the previous results, now shows some decline in foreign influence.

V. Conclusions

A series of techniques have been implemented to evaluate the extent of financial linkages among the Pacific Rim countries for which we have forward market data, and how the links have evolved over the period 1982-1992. Briefly our findings can be summarized as follows.

US and local onshore interest rates *do* appear to be increasingly correlated as time passes. For countries with forward markets, there is substantial evidence of declining covered interest differentials, indicating a greater degree of financial integration, excepting Malaysia. This finding is not in itself surprising. What is surprising is that both Canada and Japan exhibit increasing covered interest differentials. The Canadian result can perhaps be explained by reference to political events regarding the unity question. The Japanese results are more paradoxical, since the recursive CIP regression coefficients indicate *increasing* integration, even though the absolute covered interest differential has widened over time. Inspection of the data suggests that the results are due to changes in the domestic financial markets, where the Gensaki is no longer a representative rate. Hence the co-movement between the local and the covered Eurodollar rate is increasing, even as the gap between them widens, probably due to the characteristics of this particular debt instrument. It is also important to recall that the magnitudes of the deviations from covered interest parity are very small (*maximum* of –0.75), especially when contrasted with the average deviations reported pre-liberalization of –1.84.[17]

Data Appendix

A. Interest Rates

Eurocurrency deposit rates: The US 3 month Eurocurrency deposit rates were the arithmetic average of the bid and offer rates in London at close of market, as reported by Bank of America up to October 6, 1986, and Reuter's

17. Frankel [1984, p. 23] reports mean weekly deviations between the Euroyen and Gensaki rates from Jan. 1975 to April 1979. After liberalization (May 1979 to Nov. 1983) the differential is +0.26 percentage points.

Information Service thereafter, and recorded by DRI in the DRIFACS database.

Local Market Rates: Where both WFM and DRI are indicated under "Source," WFM is the source until 1989:10, at which time DRIFACS becomes the source.

Country	Source	DRI Code	Description
US	DRI	USDO3	3 month Eurodollar rate
Australia	WFM, DRI	ADBBL90Q	90 day bank bill, quote
Canada	WFM, DRI	CACP90B, A	3 month prime finance company paper
Hong Kong	WFM, DRI	HKMO3B, A	3 month interbank dep. rate
Japan	WFM, DRI	JABGDS90Y	3 month Gensaki bond rate
Malaysia	WFMr		3 month interbank dep. rate
New Zealand	WFM, WFMr		3 month commercial bills to Dec. 1987.; 3 month bank bills thereafter.
Singapore	WFMr		3 month banker's acceptances to Aug. 87; 3 month commercial bills thereafter

Notes: DRI indicates DRIFACS; WFM indicates *World Financial Markets*; WFMr indicates Morgan Guaranty's database, as provided by Carlton Strong.

B. Exchange Rates

All exchange rates (except those indicated below) are London 3PM, arithmetic average of bid and offer rates as reported by Barclay's until end of March 1990, at which time the series is no longer recorded by DRIFACS. Thereafter, the London close rate is used, as reported by Reuter's Information Services. A consistent series is not used (*i.e.*, the London close all the way) because these series only begin in 1986.

C. Actual Regression Specifications

The covered interest rate regressions were run as:

$$\ln(1+i) = \beta_0 + \beta_1 [\ln(1+i^{US}) - (\ln F - \ln S)] \quad \text{(A1)}$$

where:

F is the forward exchange rate US$/foreign currency unit,

S is the spot exchange rate US$/foreign currency unit.

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