INTERNATIONAL REAL INTEREST RATE PARITY WITH ERROR CORRECTION MODELS

Jong-Cook Byun
Son-Nan Chen

I. INTRODUCTION

Under costless financial and commodity arbitrage, incorporation of expectational (relative) purchasing power parity (PPP) into uncovered interest parity results in the equality of real interest rates between two nations, which is referred to as International Fisher-open condition. Expectational PPP implies that exchange rates will move to offset changes in price differentials, whereas uncovered interest parity indicates that arbitrage between world financial markets in the form of international capital flows should ensure that the interest differential between any two countries is the unbiased predictor of the future change in the spot exchange rate. As a result, nominal interest rate differential must be equal to expected inflation differential between two countries. This is to say that real interest rates must be identical between two nations, called the Fisher-open condition. Testing the hypothesis that real interest rates for comparable securities are equal across countries is pivotal on the ground that it provides an important implication for exchange rate determination. The results of previous empirical studies, however, are controversial. Howard and Johnson (1983) demonstrated that PPP, the Fisher open condition, and interest rate parity could not simultaneously hold in a world of taxation. Ben-Zion and Weinblatt (1984) and McClure (1988) also provided the same conclusion. Furthermore, Mishkin (1984a) documented empirical evidence rejecting the hypothesis of the equality of real Eurorates across countries. The joint hypothesis of uncovered interest parity and ex-ante relative PPP were also strongly rejected. However, his test for the independency of the Fisher-open condition, the unbiasedness of forward rate forecasts, and ex-ante relative PPP showed few rejections and high marginal significance levels. As a result, his evidence did not rule out that there was a tendency for real rates to be equal across countries over time.

On the other hand, Hodrick (1980) analyzed the equality of expected real rates across countries and found that the expected real interest rates for some countries were not statistically different. Cumby and Mishkin (1986) analyzed the international linkage of real interest rates between the European countries and the U.S. Their analysis indicated that real rates had climbed dramatically from the 1970s to the 1980s in both the European countries and the United States and that there was a positive association between movements in the U.S. real rates and those in Europe. However, European real rates typically did not move in a one-to-one manner with U.S. real rates, thereby leaving open the possibility that European monetary policy could influence U.S. domestic economic activity.

The rejection of the equality of real interest rates can also be accounted for by the existence of transaction costs, according to Frenkel and Levich (1975). However, McCormick (1979) re-estimated transaction costs by adopting Frenkel and Levich's method with higher quality data collected from Reuters news service. He found that the transaction costs in the foreign exchange market are considerably lower than those provided by Frenkel and Levich and the they cannot fully account for the discrepancies in the real rates. Other explanations for the rejection of the hypothesis are given in terms of differential price changes (Dornbusch, 1976; Frenkel, 1981b; Mussa, 1982), political risk (Aliber, 1974), capital market imperfection (Prachowny, 1970; Frenkel, 1975),[1] and differential inflation (McClure, 1988).

The failure of international interest rate parity found in some previous empirical studies is attributable to the use of the improper statistical test procedure. Previous empirical tests for the equality of real interest rates employed two-stage least squares to test coefficient restrictions. This test procedure is improper. As pointed out by Mishkin (1984a), the null hypothesis of real rate equality was jointly tested with a zero regression coefficient on an explanatory variable.[2] This test procedure becomes invalid if the discrepancy in the real rates between two countries and the explanatory variables exhibit random-walk behavior. This is to say that if the deviations from real interest rate parity exhibit random-walk behavior and hence drift apart in the short run and continue to be further apart in the long run, the tests will show a rejection of the hypothesis that the real interest rates across countries are identical. This rejection is due to the nonstationarity of real interest rate deviations and the explanatory variable that will result in estimated standard errors in regression analysis to be inconsistent, hence leading to misleading t-statistics. Consequently, in our study we will implement an alternative sound procedure for the test of international real interest rate parity. This sound test procedure, called the theory of cointegration, is developed by Engle and Granger (1987). The theory of cointegration posits that if an equilibrium relationship such as international interest rate parity is true, the deviations from the equilibrium should be stationary.[3] This means that if the equality of real interest rates holds, inter-country financial and commodity arbitrage ensure that deviations from a linear combination of real interest rates in any two countries should be stationary. Therefore, the hypothesis of the equality of real interest rates across countries can be tested using the theory of cointegration. With this sound test procedure, our study shows that international real interest rate parity holds in the long run.[4] In addition, the error correction models show that deviations from real interest rate parity take, on average, about two years to return to the long-run equilibrium of real interest rates.[5]

The rest of this paper is divided as follows: In Section II and III, we describe, respectively, the data and the methodology of cointegration. A discussion of empirical results follows, and the last section provides the concluding remarks.

II. DATA

The choice of data becomes critically important since the hypothesis of the equality of real interest rates should be applied to the comparable securities. The rates of interest should satisfy the comparability criterion, which means that the pair of securities should be at least identical in terms of maturity, risk class, political risk, etc.

The data consist of the time series of the monthly rates of average daily rates on 90-day T-bills[6] of the Organization of Economic Corporation and Development (OECD) countries. The choice of 90-day T-bill rates of the OECD countries as short-term nominal interest rates is due not only to its forecastability of future expected inflation but also to its similarity of political risk for the industrial countries in the OECD. Ten countries (Canada, the U.S., Japan, Belgium, Germany, the Netherlands, Switzerland, U.K., France, and Sweden) are chosen from the OECD. The average monthly interest rates of daily rates on 90-day T-bills have been collected for the period from February 1960 to July 1991 from the Main Economic Indicator issued by the OECD (Paris, France), the International Financial Statistics, and the International Monetary Market Year Book.[7]

But the 90-day T-bills of Japan and the remaining countries are replaced, respectively, by the "Gensaki"[8] rates and the average monthly interest rates of daily rates on three-month call money. Short-term government bills in Japan have existed only since 1986, which is too short for this empirical research. As an alternative, the Gensaki rate has nowadays often been used as the short-term interest rate in many empirical studies of the Japanese economy. The Gensaki rate is the interest rate applied to bond repurchase agreements. Unlike the call money rate in Japan, the participants in the Gensaki market are no longer limited only to financial institutions but also include corporations, government pension funds, and non-residents.

The 90-day T-bill in Germany has been almost a fixed rate. The Germany 90-day T-bill rate cannot reflect well to changes in German economy so that it may be a poor proxy for the short-term interest rate available to general investors. Thus, as an alternative, the Frankfurt call money rate has often been used as the short-term interest rate. The Frankfurt call money rate is the mid-point between extremes. Furthermore, 90-day T-bills do not exist in Switzerland and France. The average monthly interest rate of daily rates on call money is used as an alternative. The call money rate in France and Switzerland is the short-term interest rate on the collateral of public bill and is guaranteed by the government. The substitution of the Gensaki and the call money rates for 90-day T-bills may violate the assumption of the comparability criterion but will not significantly affect the test results due to the government guarantee and regulation. In other words, all securities in this sample are likely to have a similar risk exposure in terms of political and intrinsic sense.

Monthly inflation rates for individual countries from February 1960 to July 1991 are calculated as the percent changes in the Consumer Price Indices (CPI) provided by the International Financial Statistics (IMF) issued by the International Monetary Fund. The monthly rates of the CPI are set to have a base of 100 as of July 1985.

III. METHODOLOGY

When all returns, inflation, and interest rates are assumed to be continuously compounded, the Fisher condition asserts that the nominal interest rate can be expressed as the ex-ante real rate plus the expected inflation rate.[9] The Fisher Hypothesis relates the ex-ante real rate not only to expected inflation but also to market efficiency. Thus, we will test the joint hypothesis that the market is efficient and the ex-ante real interest rate is independent of the expected inflation.

Let $R^n{}_{jt}$ and $E(\pi_{jt})$ be the nominal interest rate and the expected inflation rate in country j at time t. The ex-ante real return, $R^r{}_{jt}$, on the 90-day T-bill of country j can be determined from the Fisher equation as follows:

$$R^r{}_{jt} = R^n{}_{jt} - E(\pi_{jt}) \tag{1}$$

Since Fisher equation can be applied to every individual country in the sample, the ex-ante real interest rate on the 90-day T-bill in country i ($i \neq j$) in period t can be determined similarly:

$$R^r{}_{it} = R^n{}_{it} - E(\pi_{it}) \tag{2}$$

Subtracting equation (2) from (1) yields

$$R^r{}_{jt} - R^r{}_{it} = R^n{}_{jt} - E(\pi_{jt}) - (R^n{}_{it} - E(\pi_{it})) \tag{3}$$

From the Fisher-open condition and (3), we have the following testable hypothesis

H$_o$: $R^r{}_{jt} = R^r{}_{it}$ or the ex-ante real interest rates are identical across countries

H$_a$: $R^r{}_{jt} \neq R^r{}_{it}$ or the ex-ante real interest rates vary country to country

Under the null hypothesis that the ex-ante real interest rates on the 90-day T-bills of both countries i and j are equal, equation (3) can be rewritten as

$$DR^n{}_{ijt} = DEI_{ijt} \tag{4}$$

where

$$DR^{n}_{ijt} = R^{n}_{jt} - R^{n}_{it}$$

$$DEI_{ijt} = E(\pi_{jt}) - E(\pi_{it})$$

Thus, the cointegration regression form can be written as

$$DR^{n}_{ijt} - \alpha DEI_{ijt} = U_{ijt} \tag{5}$$

where U_{ijt} is a stochastic disturbance representing a deviation from the equality of real interest rates. Then the null hypothesis of the equality of real interest rates between two countries can be rewritten as

H$_O$: the difference in the nominal interest rates between country i and j is identical to the difference in the expected inflation rates between two countries. i.e., $\alpha = 1$ in (5).
H$_A$: the difference in the nominal interest rates is not equal to the difference in the expected inflation rates. i.e., $\alpha \neq 1$ in (5)

Before applying the cointegration technique to test the hypothesis implied by (5), the definition of cointegration is introduced first. Two time series, DR^{n}_{ijt} and DEI_{ijt} in (5), are said to be cointegrated, denoted by CI (d,b), b>0, if

1. They are nonstationary.
2. They become stationary after differencing d times.
3. A linear combination of the two time series is stationary.

According to Engle and Granger (1987), this definition of cointegration implies that a linear combination of two time series such as (5) represents an equilibrium relation. Any deviation from the equilibrium tends to be corrected by economic forces which push toward a new equilibrium point. To test cointegration, DR^{n}_{t} and DEI_t (subscripts i and j are omitted for simplicity) must be determined to be stationary after differencing d times. Nonstatioanarity of the time series will be first examined by using the following Dickey-Fuller test for a unit root;

$$(1\text{-}\rho L)DR^{n}_{t} = v_t \tag{6}$$

$$(1\text{-}\rho^{*}L)DRN_t = \alpha^{*}_{o} + v^{*}_{t} \tag{7}$$

$$(1\text{-}\alpha L)DR^{N}_{t} = \alpha_o + \alpha_1(t\text{-}N/2) + v_t \tag{8}$$

where v_t is identically, independently distributed (i.i.d) with zero mean, L is the lag operator, α_i is a drift term, and N is the sample size.

The test statistics are derived based on the above models by Dickey-Fuller (1979) as follows:

$$Z(\Phi_1) = (2S\alpha^2)^{-1}[NS_o^2 - NS\alpha^2] \text{ for } H_o{:}(\alpha_o,\rho) = (0,1) \quad (9)$$

$$Z(\Phi_2) = (3S_t2)^{-1}[NS_o^2 - NS_t^2] \text{ for } H_o{:}(\alpha_o{}^*,\alpha_1,\rho) = (0,0,1) \quad (10)$$

$$Z(\Phi_3) = (2S_t^2)[N\{S_o^{2-}(DR_t^N - DR_{t-1}{}^N)^2\}\text{-}NS_t^2]$$

$$\text{for } H_o{:}(\alpha_o,\alpha_1,\rho) = (\alpha_o,0,1) \quad (11)$$

where $S_o^2 = N\text{-}^1\Sigma_t = 1^N(DR_tN\text{-}DR_{t-1}{}^N)^2$, and $S\alpha^2$ and S_t^2 are, respectively, the variances of the estimated residuals from (7) and (8). The critical values are available in Dickey-Fuller (1979). If the null hypothesis of nonstationarity cannot be rejected, the tests on the first differenced series is performed again until the time series are stationary.

The inflation rate used in the test is in fact defined as the expected inflation rate, which is unobservable in contrast to the ex post inflation rate. We estimate the expected inflation rate using two different approaches. First, we use the contemporaneous inflation rate (from the inflation data) as a proxy for expected inflation. Second, we decompose the inflation rate into the expected and unexpected components by ARIMA models.[10] Once the expected inflation rate is estimated, the test of nonstationarity on the difference in the expected inflation rates between country i and j will be performed in the same procedure as in the test on the nominal interest rate.

For the purpose at hand, the two time series, DR_t^N and DEI_t, are assumed to be integrated of order 1. The cointegration test can then be performed by using equation (5). The Ordinary Least Square estimator is consistent if the residual error U_t in (5) is stationary. Again, the Dickey-Fuller test for a unit root is performed on the residual error U_t using the following regression model;

$$(1\text{-}L)U_t = \text{-}\gamma_o U_t\text{-}1 + \Sigma_{t\,=\,1}{}^N(1\text{-}L)\gamma_i U_{t\text{-i}} + \omega_t \quad (12)$$

where L denotes the lag operator, $U_{t\text{-i}}$ is the ith distributed lag residual of (5), and ω_t is an i.i.d. disturbance with zero mean.

If the two series are cointegrated of order (1,1), the estimated residual U_t must be stationary. Thus, if U_t is stationary, then the coefficient for go in (12) should be statistically different from zero. The critical values can be obtained from Engle and Granger (1987).

To generate error-correcting models, let ß be a second cointegrating scalar such that

$$DR^n{}_t - \beta DEI_t = \varepsilon_t,\ e_t = e_{t-1} + \varepsilon_t \quad (13)$$

Rewriting equation (5) yields

$$DR^n{}_t - \alpha DEI_t = U_t,\ U_t = \rho U_{t-1} + z_t,\ \rho \neq 1 \quad (14)$$

where ε_t and z_t are white noise disturbances.

The solutions to the simultaneous equations in terms of DR''_t and DEI_t can be used in the following error-correcting representation;

$$(1\text{-}L)DR''_t = -\pi DR''_{t\text{-}1} + \alpha\pi DEI_t + \zeta_t$$

$$= -\pi U_{t\text{-}}1 + \zeta_t \qquad (15)$$

$$(1\text{-}L)DEI_t = -\eta DR''_{t\text{-}1} - \alpha\eta DEI_t + \xi_t$$

$$= -\eta U_{t\text{-}1} + \xi_t \qquad (16)$$

where $\pi = \beta\eta$, $\eta = (1\text{-}\rho)/(\beta\text{-}\alpha)$, $U_{t\text{-}1} = DR''_{t\text{-}1}\text{-}\alpha D\mathrm{EI}_{t\text{-}1}$, and ζ_t and ξ_t are a combination of ε_t and ζ_t.

Error-correction models in (15) and (16) provide an appropriate and useful method for describing the linkage between the real interest rates and the expected inflation rates in two countries. In addition, the models are flexible enough to incorporate different irrational factors presumed to affect the interest rate and the inflation rate.

IV. EMPIRICAL RESULTS

We assume that the two time series, DRt and DEIt, are adequately represented by three different models in equations (6) through (8). Equation (6), which contains neither a constant nor a trend as regressors, is appropriate in a driftless case, whereas equation (7), which incorporates a constant as a regressor, allows for a nonzero mean in the series. When an economic series tends to increase over time, equation (7) is appropriate for an estimation process. If a nonzero drift is highly suspended, equation (8) is the most relevant model for testing a unit root. Most aggregate macroeconomic variables are likely to be represented by equation (8).

A. Tests of Unit Roots

The unit root test is conducted on the level and the first difference of DRt and DEIt. Following Perron (1988), we first employ $Z(\Phi_3)$ to test the null hypothesis of a unit root. If the null hypothesis is not rejected, it need not indicate non-stationarity since non-rejection could be due to the low power of the test of $Z(\Phi_3)$. $Z(\Phi_2)$ is then used to test the null hypothesis of zero drift. If the time series under consideration is suspected to have no drift (upward or downward) over time, the null hypothesis of zero drift should not be rejected. This implies that model (7) is appropriate for detecting a unit root. However, if the time series is highly suspected to have a zero mean, the estimates from equation (7) are more appropriate.

Table 1 summarizes the test results of unit roots when the contemporaneous inflation rates (obtained from the original inflation data) are employed as proxies for the expected inflation. On the level of time series DR_t, which is the difference in the nominal interest rates between countries i and j, the null hypothesis of a unit root cannot be rejected at the 1% level of significance for all pairs of countries with the U.S. and U.K. The presence of a unit root may be resulted from the poor power of the test of $Z(\Phi_3)$ statistic. As a result, the existence of a drift is next verified by $Z(\Phi_2)$. Observing Table 1, the null hypothesis ($\alpha_0^* = 0$, $\alpha_1 = 0$, and $\rho = 1$) cannot be rejected except for the pairs of countries GM/US and ND/UK whose $Z(\Phi_2)$ values are 7.46, and 7.93, respectively, with the critical value of 6.22 at the 1% level of significance as given in Dickey-Fuller (1979). This implies that the nominal interest rates, DRt, for those countries tend to increase over time. This result can be further confirmed by using $Z(\Phi_1)$. The test result again shows that the null hypothesis of a unit root without a trend is rejected at the 5% level of significance only for ND/UK but cannot be rejected at the 1% level of significance. The overwhelming evidence indicates that the level of the series, DRt, is nonstatioary. To find the next order of integration, the same test statistics are employed for the first difference of the DRt series. In Table 1, the test statistic $Z(\Phi_3)$ for the first difference of the DRt series has shown that the null hypothesis of a unit root is rejected at the 1% level of significance for all pairs of countries, except for SW/US and SW/UK. To verify the non-rejection for SW/US and SW/UK, we further test whether $Z(\Phi_2)$ and $Z(\Phi_1)$ are statistically significant. At the 1% level of significance, the null hypothesis cannot be rejected, which implies that SW/US and SW/UK must be differenced more than once to be stationary.

With the same procedure, the test of unit roots can be applied to the contemporaneous (original) inflation rates which are used as proxies for the expected inflation rates. Table 1 shows that the null hypothesis of a unit root cannot be rejected at the 1% level of significance for all pairs. But non-rejection of the null hypothesis using $Z(\Phi_3)$ statistic must be further verified by using $Z(\Phi_2)$ to test the existence of a drift. Five (FR/US, GM/US, UK/US, GM/UK, and JP/UK) out of the fifteen pairs of countries in Table 1 show that the contemporaneous inflation rates tend to increase over time. However, $Z(\Phi_1)$ statistic confirms that the null hypothesis of a unit root without a drift cannot be rejected at the 1% level of significance. Therefore, the overall test results have shown that two time series of DR_t and DEI_t are integrated of order 1, except for DR_t of SW/US and SW/UK. Since SW/US and SW/UK have two time series with different order of integration, they are excluded from the test of cointegration.

B. ARIMA Models of Expected Inflation

We also employ the ARIMA models to generate expected and unexpected components of the inflation rate using the procedure developed by Box and Jenkins (1970). The inflation forecasts from ARIMA models are then used as the estimates of expected inflation, and the forecast errors are used as the unexpected components of the inflation rates. Table 2 presents the results of the ARIMA models for the period of February 1960 through July 1991. A difficulty with using the ARIMA models to forecast inflation for long periods is that, for most countries, inflation rates are more volatile in the 1970s than in the 1960s and 1980s. It seems to have a structural shift in the inflation process in many

Table 1a
Unit Root Tests

	Interest Rate						Contemporaneous Inflation					
	Level			Difference			Level			Difference		
	$Z(\Phi_1)$	$Z(\Phi_2)$	$Z(\Phi_3)$	$Z(\Phi_1)$	$Z(\Phi_2)$	$Z(\Phi_3)$	$Z(\Phi_1)$	$Z(\Phi_2)$	$Z(\Phi_3)$	$Z(\Phi_1)$	$Z(\Phi_2)$	$Z(\Phi_3)$
BG/US	4.57	3.57	0.77	0.13	10.3*	15.4*	0.12	1.02	1.42	0.13	35*	53*
CA/US	2.82	4.48	4.32	0.35	6.53*	9.76*	0.29	0.20	0.80	0.76	157*	236*
FR/US	6.08	4.56	0.74	0.11	8.39*	12.6*	4.40	6.62*	0.92	0.54	51*	76*
GM/US	5.34	7.46*	4.40	0.31	20.7*	31.0*	4.92	7.33*	3.39	0.85	27*	41*
JP/US	0.31	0.79	1.15	0.34	3.48	8.78*	0.28	6.09	7.33	0.34	85*	127*
ND/US	3.54	2.37	0.20	0.17	11.3*	16.9*	0.90	0.60	0.10	0.91	54*	80*
SW/US	1.32	0.96	0.12	0.16	2.86	4.27	3. 9	4.08	3.86	0.20	39*	59*
UK/US	5.24	4.50	1.47	0.14	7.42*	11.1*	5. 7	9.13*	0.16	0.21	47*	71*
BG/UK	2.11	1.78	0.55	0.27	5.51	8.94*	3. 7	2.58	0.39	0.11	57*	86*
CA/UK	1.93	1.43	0.22	0.58	7.57*	11.3*	2. 8	1.99	0.14	0.15	125	186*
FR/UK	0.53	0.47	0.18	0.57	4.80	9.29*	3. 0	2.94	0.50	0.52	38*	57*
GM/UK	5.52	5.02	5.60	0.74	15.9*	23.9*	3. 5	7.70*	2.30	0.49	36*	55*
JP/UK	0.87	3.35	4.13	1.00	5.32	8.87*	5. 0	9.78*	6.01	0.58	64*	96*
ND/UK	6.32	7.93*	3.29	0.36	8.59*	12.9*	5. 0	3.36	0.14	0.54	59*	89*
SW/UK	2.42	2.27	0.97	0.98	3.34	5.14	6. 3	5.69	0.98	0.35	57*	85*

Note: BG : Belgium
CA : Canada
FR : France
GM : Germany
JP : Japan
ND : Netherlands
SW : Switzerland
UK : United Kingdom
US : United States
The numbers with the asterisk are significant at the 1% level of significance.

Table 1b
Critical Values

Statistics	*5%*	*2.5%*	*1%*
$Z(\Phi_1)$	4.63	5.45	6.52
$Z(\Phi_2)$	4.75	5.40	6.22
$Z(\Phi_3)$	6.34	7.25	8.43

Note: Source: Dickey and Fuller (1979).

countries. Thus, the results may not be robust with respect to a particular ARIMA model that is chosen in Table 2.

Table 3 shows the test results of a unit root conducted on the level and the first difference of the expected inflation rate forecasted by the ARIMA models in Table 2. The test statistic $Z(\Phi_3)$ on the level of the expected inflation rate, DEIt, reveals that the null hypothesis of a unit root cannot be rejected at the 1% level of significance for all pairs. With the test statistic $Z(\Phi_2)$, the null hypothesis is rejected in the case of UK/US where $Z(\Phi_2)$ is 6.74 with the critical value of 6.22 at the 1% level of significance. After reconfirming the null hypothesis of a unit root with a drift using $Z(\Phi_1)$, UK/US cannot reject the null hypothesis. However, the tests of unit roots based on the first difference of the expected series show that

Table 2
Arima Models for Expeccted Inflation

	ARIMA	*Parameter Estimates*			
	(p,d,q)	*AR1*	*AR2*	*MA1*	*MA2*
Belgium	(0,1,2)			.773	.120
				(.051)	(.051)
Canada	(0,1,2)			1.652	-.674
				(.040)	(.042)
France	(0,1,1)			.855	
				(.026)	
Germany	(1,1,1)	.261		.929	
		(.057)		(.024)	
Japan	(2,1,2)	-.338	-.223	.488	.410
		(.153)	(.062)	(.155)	(.150)
Netherlands	(1,0,1)	.904		.882	
		(.201)		(.221)	
Switzerland	(0,1,1)			.981	
				(.056)	
U.K.	(0,1,2)			.786	.112
				(.051)	(.051)
U.S.	(1,1,1)	.140		.839	
		(.064)		(.034)	

Note: Figures in the parenthesis are the standard deviations for the parameter estimates of the ARIMA models.

all pairs in the sample strongly reject the null hypothesis at the 1% level of significance, except for SW/US and SW/UK. Therefore, the evidence has indicated that the expected inflation rate must be first differenced to be stationary, except for SW/UK and SW/US.

D. Cointegration

Inter-country commodity and financial arbitrage ensures that deviations from a linear combination of the real interest rates between two countries should be stationary if the real interest rates are cointegrated. The tests of cointegration in this study are conducted in the overall period and four different subperiods with different lengths of time: 1960 through 1973 and 1974 through 1991 (representing the period of fixed and floating exchange rates); 1973 through 1981 and 1982 through 1991 (representing the periods of high and low inflation rates).

Table 3a
Unit Root Tests for Expected Inflation

	Expected Inflation					
		Level			*Difference*	
	$Z(\Phi_1)$	$Z(\Phi_2)$	$Z(\Phi_3)$	$Z(\Phi_1)$	$Z(\Phi_2)$	$Z(\Phi_3)$
BG/US	.013	.325	.476	.036	5.38	9.07*
CA/US	.288	.612	.628	.035	52.74*	79.10*
FR/US	2.920	2.191	.350	.015	5.29	8.91*
GM/US	2.951	4.242	3.411	.011	7.17	10.76*
JP/US	.933	5.424	4.293	.064	20.53*	30.79*
ND/US	4.323	3.042	.242	.053	3.95	8.96*
SW/US	2.274	1.617	.165	.034	2.00	2.97
UK/US	4.956	6.742*	.156	.025	7.07*	10.61*
BG/UK	8.612*	6.491	1.102	.016	6.41	9.61*
CA/UK	2.983	2.186	.281	.012	50.13*	75.20*
FR/UK	3.041	2.522	.754	.021	5.25	8.86*
GM/UK	6.832*	5.754	1.731	.084	4.36	8.54*
JP/UK	14.140*	4.322	5.481	.010	25.61*	38.41*
ND/UK	6.260	4.340	.291	.010	3.08	8.66*
SW/UK	4.370	3.140	.322	.037	2.32	3.45

Note: The country code and Critical values are shown in Table 3.1. The values with * are significant at the 1% level of significance.

Table 3b
Critical Values

Statistics	*5%*	*2.5%*	*1%*
$Z(\Phi_1)$	4.63	5.45	6.53
$Z(\Phi_2)$	4.75	5.40	6.22
$Z(\Phi_3)$	6.34	7.25	8.43

Source: Dickey and Fuller (1979).

According to Engle and Granger (1987), the estimated slope coefficient of cointegration regression (5) is a consistent estimator of α if the residual, U_t, is stationary. Table 4 reports these estimates of the coefficients. In the overall period of 1960 through 1991, the estimated values of α for the pairs of BG/US, FR/US, GM/US, JP/US, FR/UK, GM/UK, and JP/UK are not statistically significant to reject the null hypothesis of $\alpha = 1$ at the 5% level of significance, but α's for the remaining pairs are significantly below unity. However, prior to the tests of cointegration, the estimated values of α from cointegration regression (5) cannot be considered to be consistent for some pairs of countries. As seen in BG/US and FR/UK, α's in the two subsample periods of 1960-1973 and 1973-1991 are appar-

Table 4
Cointegration Regression [Equation (5)]

	BG/US	*CA/US*	*FR/US*	*GM/US*	*JP/US*	*ND/US*	*UK/US*
1960-1991							
Est. α	.774*	.022	.842*	1.02*	1.28*	.023	.582
	(.203)	(.043)	(.243)	(.303)	(.207)	(.050)	(.160)
1960-1973							
Est. α	-.197	.015	-.630	.615	.273	.05	.151
	(-.048)	(.022)	(.209)	(.375)	(.208)	(.003)	(.113)
1973-1991							
Est. α	1.244*	.171	1.58	.881*	1.29	.306	.626*
	(.283)	(.206)	(.374)	(.410)	(.266)	(.275)	(.227)
1973-1981							
Est. α	1.109*	.182	1.769*	.106	1.20*	.461	.818*
	(.446)	(.260)	(.676)	(.771)	(.386)	(.362)	(.286)
1982-1991							
Est. α	1.548*	-.018	1.598*	.717*	.310	-.351	.483
	(.326)	(.324)	(.334)	(.498)	(.299)	(.462)	(.428)
	BG/UK	*CA/UK*	*FR/UK*	*GM/UK*	*JP/UK*	*ND/UK*	
1960-1991							
Est. α	.031	.055	.907*	1.07*	1.02*	.014	
	(.039)	(.047)	(.178)	(.217)	(.210)	(.036)	
1960-1973							
Est. α	.145	.016	-.022	.212	.147	-.008	
	(.160)	(.020)	(.167)	(.307)	(.178)	(.002)	
1973-1991							
Est. α	.022	.452	1.29*	.658*	.292	.398	
	(.050)	(.185)	(.274)	(.247)	(.229)	(.152)	
1973-1981							
Est. α	.306	.503*	.729	1.09*	.376	.306	
	(.213)	(.279)	(.335)	(.429)	(.303)	(.213)	
1982-1991							
Est. α	.796*	.113	2.768	.639*	.553*	.738*	
	(.241)	(.242)	(.403)	(.281)	(.299)	(.251)	

Note: The estimates of α's are the coefficients by regressing the nominal interest rates differential on the expected inflation rates differential between each pair of countries The number in th parenthesis is the standard deviation of cointegration regression. *denotes "fail to reject H0: $\alpha = 1$ at the 5% significance level."

ently different. The different α's in these two pairs are forced into one in the overall period.

The tests of cointegration will be conducted first by the analysis of the residual in (5) such that the residual should be stationary. In addition, for the two time series, DRt and DEIt, to be cointegrated, the coefficient, γ_o, in (12) must be statistically significant different from zero, i.e., to reject the null hypothesis of no cointegration. The results of the cointegration tests are reported in Tables 5 and 6. Table 5 exhibits the results when the contemporaneous (original) inflation rates are used as the proxies for the expected inflation rates, while Table 6 reports the results when the expected inflation rates are forecasted by ARIMA models. Observing Tables 5 and 6, we have the following empirical results:

First, the real interest rates for comparable securities are not generally equal across countries in the sample, but the equality of real interest rates as a long-run relationship performs well for the overall sample period. In order for the equality of real interest rates to fit perfectly, the null hypothesis that the difference in the nominal interest rates between the pairs of countries is identical to the difference in the expected inflation rates between the two countries [i.e., $\alpha = 1$ in (5)] should not be rejected, and the hypothesis of no cointegration [H_0: $\gamma_0 = 0$ in (12)] should be rejected. As shown in Tables 4 and 5, the country pairs such as FR/US, GM/US, JP/US, FR/UK, GM/UK, and JP/UK fail to reject the null hypothesis of $\alpha = 1$ in (5), and no cointegration [H_0: $\gamma_0 = 0$ in (12)] is rejected at the 5% level of significance in the overall period. This test result shows that the difference in the nominal interest rates between the pairs of those countries is equal to the difference in the expected inflation rates for those countries. That is, international real interest rate parity can be considered as a long-run "equilibrium" relationship for those pairs of countries. The other remaining pairs of countries (BG/US, CA/US, ND/US, UK/US, BG/UK, CA/UK, and ND/UK) over 1960 through 1991 have shown that the real interest rates are not manifestly equal (according to the Fisher-open condition) due to a reject of the null hypothesis [H_0: $\alpha = 1$ in (5)], but they are strongly cointegrated because the null hypothesis [H_0: $\gamma = 0$ in (12)] is rejected. As pointed out by Granger (1986), and Engle and Granger (1987), if an equilibrium relationship exists, the deviation from the equilibrium should be stationary. Table 5 reports that all pairs of countries show a strong cointegration since the *t*-values of the coefficients, γ_o and $\gamma_o{}^4$, of the residual are greater than the critical values 3.37 for γ_o and 3.17 for go4 provided by Engle and Granger (1987) (the distributed lags are 0 and 4. See the footnote of Table 5). Therefore, the equality of real interest rates in (4) performs well as a long-run relationship of real interest rates.

Second, it is very interesting to note that the equality of real interest rates as a long-run relationship across countries is cointegrated well under both exchange rates regimes. But, it is cointegrated better under the floating exchange rate system. In Table 5, for the period of 1960 through 1973 representing the period of the fixed exchange rate system, nine out of the thirteen pairs of countries reject the null hypothesis of no cointegration, whereas, under the floating exchange rate, eleven out of thirteen pairs reject the null hypothesis. This result supports the recent argument that a high real interest rate in one country would be trans-

Table 5
Test for Cointegration Using Contemporaneous Rates

	BG/US	*CA/US*	*FR/US*	*GM/US*	*JP/US*	*ND/US*	*UK/US*
1960-1991							
γ_0^0	-.131	-.065	-.147	-.454	-.112	-.146	-.122
	(-5.08)*	(-3.54)*	(-5.47)*	(-10.46)*	(-4.74)*	(-5.29)*	(-4.92)*
γ_0^4	-.142	-.050	-.141	-.186	-.064	-.142	-.107
	(-1.75)	(-2.57)	(-4.61)*	(-3.51)*	(-2.72)	(-4.27)*	(-3.86)*
1960-1973							
γ_0^0	-1.96	-.289	-.338	-.673	-.050	-.257	-.290
	(-4.08)*	(-5.03)*	(-5.71)*	(-8.92)*	(-1.88)	(-4.64)*	(-5.15)*
γ_0^4	-.121	-.169	-.203	-.279	-.041	-.126	-.254
	(-2.25)	(-2.24)	(-2.84)	(-2.79)	(-1.53)	(-1.72)	(-3.68)*
1973-1991							
γ_0^0	-.172	-.087	-.152	-.387	-.120	-.129	-.107
	(-4.53)*	(-3.56)*	(-4.25)*	(-7.49)*	(-3.66)*	(-3.72)*	(-3.43)*
γ_0^4	-.212	-.092	-.153	-.148	-.078	-.140	-.106
	(-4.61)*	(-3.36)*	(-3.79)*	(-2.98)	(-2.28)	(-3.35)*	(-3.16)
1973-1981							
γ_0^0	-.155	-.090	-.161	-.480	-.155	-.223	-.176
	(-2.95)	(-2.31)	(-2.93)	(-5.63)*	(-2.85)	(-3.52)*	(-3.14)
γ_0^4	-.209	-.087	-.165	-.220	-.085	-.287	-.194
	(-3.16)	(-1.84)	(-2.70)	(-2.48)	(-1.44)	(-3.45)*	(-2.92)
1982-1991							
γ_0^0	-.255	-.069	-.174	-.019	-.028	-.029	-.063
	(-4.13)*	(-2.03)	(-3.38)*	(-0.62)	(-0.98)	(-0.95)	(-1.87)
γ_0^4	-.177	-.064	-.122	-.007	-.079	-.036	-.065
	(-2.43)	(-1.81)	(-2.10)	(-0.24)	(-2.53)	(-1.09)	(-1.78)

	BG/UK	*CA/UK*	*FR/UK*	*GM/UK*	*JP/UK*	*ND/UK*
1960-1991						
γ_0^0	-.074	-.089	-.116	-.343	-.075	-.109
	(-3.67)*	(-4.19)*	(-4.81)*	(-8.83)*	(-3.85)*	(-4.65)*
γ_0^4	-.085	-.120	-.092	-.140	-.045	-.085
	(-3.81)*	(-4.65)*	(-3.55)*	(-3.26)*	(-2.38)	(-3.30)*
1960-1973						
γ_0^0	-.086	-.115	-.135	-.546	-.036	-.114
	(-2.66)	(-3.17)	(-3.36)	(-7.63)*	(-1.71)	(-3.06)
γ_0^4	-.122	-.150	-.115	-.224	-.044	-.120
	(-3.39)*	(-3.61)*	(-2.55)	(-2.60)	(-1.98)	(-2.69)
1973-1991						
γ_0^0	-.073	-.123	-.124	-.383	-.051	-.226
	(-2.80)	(-3.89)*	(-3.79)*	(-7.53)*	(-2.36)	(-5.19)*
γ_0^0	-.018	-.144	-.960	-.161	-.056	-.216
	(-2.85)	(-4.01)*	(-2.71)	(-3.27)*	(-2.47)	(-3.94)*
1973-1981						
γ_0^0	-.234	-.120	-.103	-.431	-.056	-.234
	(-3.54)*	(-2.48)	(-2.28)	(-5.32)*	(-1.64)	(-3.54)*
γ_0^4	-.200	-.152	-.125	-.194	-.069	-.200
	(-2.28)	(-2.70)	(-2.45)	(-2.37)	(-1.91)	(-2.28)

(*continued*)

Table 5 (continued)
Test for Cointegration Using Contemporaneous Rates

	BG/US	*CA/US*	*FR/US*	*GM/US*	*JP/US*	*ND/US*	*UK/US*
1982-1991							
γ_0^0	-.267	-.100	-.261	-.139	-.117	-.242	
	(-4.04)*	(-2.38)	(-4.10)*	(-3.74)*	(-2.66)	(-3.75)*	
γ_0^4	-.232	-.140	-.126	-.126	-.078	-.212	
	(-2.77)	(-2.92)	(-1.93)	(-2.09)	(-1.62)	(-2.50)	

Note: The cointegration tests are conducted on the difference in the nominal interest rates, DR_{ijt}, and the difference in the expected inflation rates, DEI_{ijt}, between each pair of countries. The numbers in the parenthesis are the t-values for the estimated coefficients, γ_0^0 and γ_0^4. γ_0^0 and γ_0^4 are respectively the estimated values for regression equation (12) when the distributed lag n = 0 and 4. As given in Engle and Granger (1987), the critical values for t-statistic on γ_0^0 are respectively 4.07, 3.37, 3.03 at the 1%, 5%, and 10% significance levels, while, with four distributed lags of the residual, 3.77, 3.17, and 2.84 at the 1%, 5%, and 10% levels, respectively. Here, we test the null hypothesis of no cointegration at the 5% level of significance. *denotes significance at the level 5% level.

mitted abroad. Under the floating exchange rate system, the high real interest rate in a country tends to increase capital inflows from abroad, and an increase in capital inflows results in a rise in its currency value but a fall in currency values in other countries. Thus, this imbalance in the foreign currency markets can be resolved through an increase in real interest rates in other countries. As a result, the real interest rates are likely to be equal across countries. However, it should be emphasized that this result provides no evidence on the direction of causation of real interest rate movements.

Third, use of different proxies for the unexpected inflation rates leads to substantially different test results for the equality of real interest rates. In Table 6, where expected inflation rates are estimated by ARIMA models, the *t*-values of the coefficients on the residual increase very significantly to strongly reject the null hypothesis of no cointegration for all country pairs over the whole period as well as over most subperiods (the distributed lags are 0 and 4. See the footnote of Table 6). Especially, over the subperiod from 1973 through 1981 representing the high inflation period, with contemporaneous inflation rates (from the original inflation data) as proxies for the expected inflation rate, 45% of the pairs shows cointegration, whereas 77% of the pairs reject the null hypothesis of no cointegration when the expected inflation rates are estimated by the ARIMA models. Thus, the cointegration test results have shown that the equality of real interest rates is more likely to hold if a better proxy for the expected inflation rate is used.

Fourth, in the 1980s period when the European Monetary System (EMS) was put into effect, the real interest rates in the pairs between U.K. and the other EMS countries appear to be more strongly cointegrated than those in the pairs between U.K. and the non-EMS countries.[11] As expected, Table 5 shows that all pairs between U.K. and the other EMS countries reject the null hypothesis of no cointegration at the 5% level of significance. This means that an imbalance in the real interest rates will be adjusted by a control through the EMS so that the real interest rates can be restored to be close to equilibrium.

Table 6
Test for Cointegration Using Expected Inflation

	BG/US	*CA/US*	*FR/US*	*GM/US*	*JP/US*	*ND/US*	*UK/US*
1960-1991							
γ_0^0	-.097	-.067	-.135	-.445	-.420	-.147	-.114
	(-4.24)*	(-3.57)*	(-5.13)*	(-10.31)*	(-9.99)*	(-5.31)*	(-4.72)*
γ_0^4	-.094	-.052	-.147	-.203	-.167	-.145	-.110
	(-3.68)	(-2.61)	(-4.83)*	(-3.79)*	(-4.03)	(-4.32)*	(-4.05)*
1960-1973							
γ_0^0	-.080	-.300	-.332	-.672	-.169	-.251	-.270
	(-2.56)*	(-5.12)*	(-5.59)*	(-8.91)*	(-3.74)	(-4.53)*	(-4.86)*
γ_0^4	-.117	-.159	-.205 -	.282	-.082	-.126	-.251
	(-3.39)	(-2.28)	(-2.85)	(-2.86)	(-2.74)	(-1.71)	(-3.69)*
1973-1991							
γ_0^0	-.098	-.947	-.117	-.363	-.380	-.124	-.091
	(-3.39)*	(-3.58)*	(-3.60)*	(-7.21)*	(-6.69)*	(-3.60)*	(-3.17)*
γ_0^4	-.093	-.101	-.156	-.159	-.195	-.135	-.097
	(-2.76)*	(-3.40)*	(-4.32)*	(-3.22)	(-3.61)	(-3.27)*	(-3.07)
1973-1981							
γ_0^0	-.146	-.263	-.169	-.511	-.416	-.219	-.146
	(-2.80)	(-4.06)*	(-3.02)	(-6.15)*	(-5.05)*	(-3.49)*	(-2.79)
γ_0^4	-.214	-.251	-.227	-.323	-.257	-.300	-.193
	(-3.29)*	(-2.89)	(-3.55)*	(-3.25)*	(-2.89)	(-3.54)*	(-3.17)*
1982-1991							
γ_0^0	-.130	-.156	-.010	-.023	-.059	-.024	-.058
	(-2.55)*	(-3.08)	(-2.27)*	(-0.08)	(-1.92)	(-0.08)	(-1.87)
γ_0^4	-.156	-.141	-.115	-.010	-.088	-.036	-.061
	(-2.75)	(-2.47)	(-2.54)	(-0.30)	(-2.40)	(-1.15)	(-1.83)

	BG/UK	*CA/UK*	*FR/UK*	*GM/UK*	*JP/UK*	*ND/UK*
1960-1991						
γ_0^0	-.097	-.086	-.124	-.363	-.369	-.134
	(-4.23)*	(-4.08)*	(-4.98)*	(-9.12)*	(-9.23)*	(-5.17)*
γ_0^4	-.094	-.109	-.109	-.161	-.137	-.102
	(-3.68)*	(-4.69)*	(-3.91)*	(-3.55)*	(-3.58)*	(-3.52)*
1960-1973						
γ_0^0	-.081	-.108	-.143	-.527	-.164	-.116
	(-2.56)	(-2.99)	(-3.45)*	(-7.42)*	(-3.75)*	(-3.08)
γ_0^4	-.117	-.150	-.120	-.218	-.104	-.122
	(-3.39)*	(-3.61)*	(-2.54)	(-2.58)	(-3.62)*	(-2.70)
1973-1991						
γ_0^0	-.099	-.119	-.110	-.392	-.122	-.194
	(-3.39)*	(-3.77)*	(-3.55)*	(-7.60)*	(-3.74)*	(-4.73)*
γ_0^0	-.093	-.156	-.101	-.160	-.089	-.173
	(-2.76)	(-4.31)*	(-2.89)	(-3.25)*	(-3.61)*	(-3.46)*
1973-1981						
γ_0^0	-.139	-.106	-.091	-.550	-.176	-.269
	(-2.44)	(-2.21)	(-6.34)*	(-2.97)	(-3.99)*	
γ_0^4	-.189	-.152	-.113	-.248	-.145	-.270
	(-3.16)*	(-2.97)	(-2.38)	(-2.66)	(-2.33)	(-3.05)

(*continued*)

Table 6 (continued)

Test for Cointegration Using Expected Inflation

	BG/US	*CA/US*	*FR/US*	*GM/US*	*JP/US*	*ND/US*	*UK/US*
1982-1991							
γ_o^O	-.282	-.183	-.344	-.127	-.178	-.173	
	(-4.33)*	(-3.36)*	(-4.75)*	(-2.63)	(-3.37)*	(-3.13)	
γ_o^4	-.260	-.251	-.236	-.114	-.091	-.155	
	(-3.32)	(-3.84)*	(-2.44)	(-1.88)	(-1.49)	(-2.21)	

Note: Two time series, DR_{ijt} and DEI_{ijt}, between each pair of countries are employed for the tests of cointegration. The numbers in the parenthesis are the t-values for the estimated coefficient, γ_o^O and γ_o^4. γ_o^O and γ_o^4 are respectively the estimated values for regression equation (12) when the distributed lag n = 0 and 4. As given in Engle and Granger (1987), the critical values for t statistic on γ_o^O are respectively 4.07, 3.37, 3.03 at the 1%, 5%, and 10% significance levels, while, with four distributed lags of the residual, 3.77, 3.17, and 2.84 at the 1%, 5%, and 10% levels, respectively. Here, we test the null hypothesis of no cointegration at 5% level of significance. Thus, * denotes significant at the 5% level.

Finally, for the overall sample period, the evidence in this study suggests that the nominal interest rate and the expected inflation rate differentials in the pairs between the U.S. and the other European countries may be at least as cointegrated as those in the pairs between U.K. and the other European countries. This is due to the fact that, given in Tables 5 and 6, the *t*-values of the pairs between the U.S. and the other European countries are much higher than those of the pairs between U.K. and the other European countries. While the U.K. economy is cointegrated with the other EMS countries, this test result highlights the review that the economies of the European countries are at least as cointegrated with the U.S. economy as with the U.K. economy. This can be attributed to heavy international trade and investments between the U.S. and the European countries. In addition, Germany shows significantly higher *t*-values with the U.S. and U.K. In contrast, the *t*-values of the pair between UK/US are significantly low as compared to GM/US and GM/UK. This may positively support that the German economy is strongly cointegrated with the U.S. economy.

E. Error Correction Models

As mentioned in the theory of cointegration, the residual of the cointegration regression (5) can be used as instruments to estimate an error correction model if two time series are cointegrated. The error correction models in (15) and (16) can be determined to show how the interest and inflation rates adjust to eliminate any deviation from the equality of real interest rates which represents an equilibrium relationship.

Since the error correction models in (15) and (16) can be applied only to the case where two series are cointegrated, we will consider only two different time periods over which most pairs between countries in the study have shown strong cointegration: 1960 through 1991 and 1973 through 1991. For example, the estimated error correction models for the pair between Germany and the U.S. during 1960 through 1991 can be represented by;

$$(1\text{-L})\mathrm{DR}_t^{\mathrm{n}} = -.437\ \mathrm{U}_{t\text{-}1} \tag{17}$$
$$(.042)$$

$$(1\text{-L})\mathrm{DEI}_t = .017\ \mathrm{U}_{t\text{-}1} \tag{18}$$
$$(.009)$$

where DR_t^n is the difference in the nominal interest rates between Germany and the U.S., DEI_t is the difference in the expected inflation rates between the two countries, U_{t-1} is the residual from the cointegration regression, and the numbers in the parentheses represent the standard errors. Equation (17) indicates that the nominal interest rates in both countries quickly respond to adjust deviations from the equilibrium so as to return strongly to the equality of real interest rates. This is because approximately 44% of any positive (or negative) deviation from the equilibrium is eliminated within a month by a fall (or a rise) in the difference in the nominal interest rates between the two countries. In contrast, the error correction model in (18) shows that only 2% of any positive (or negative) deviation can be adjusted within a month by an increase (or a decrease) in the difference in the expected inflation rates between the two countries. This implies that the nominal interest rate in either Germany or the U.S. or both responds sensitively to any deviation from the equality of real interest rates, but the inflation rate barely responds to it since it is intuitive and logical that changes in the prices on goods are sticky (i.e., the slow adjustment of the prices). Thus, it will take approximately five months to return to the equality of real interest rates between Germany and the U.S. because of the sticky prices and the immediate adjustment of nominal interest rates by the two countries.

In addition, the GM/UK pair also shows a quick response of the nominal interest rates to deviations from the equality of real interest rates. As shown in Table 7, over the whole period 30% of any positive (or negative) deviation from the equality of real interest rates between GM/UK can be eliminated within a month by a fall (or a rise) in the difference in the nominal interest rates, whereas 3% of any positive (or negative) deviation can be adjusted within a month by an increase (or a decrease) in the difference in the expected inflation rates between the two countries. Therefore, the time lag to return to the equilibrium relationship (the equality of real interest rates) in GM/UK is almost seven months, which is longer than the time lag in the GM/US pair. This indicates that Germany is more closely related to the U.S. than to U.K. in terms of real interest linkage.

The overall results of the error correction models are provided in Table 7. It is very interesting to note that the contribution of the expected inflation rate to the adjustment of any deviation from the equality is higher between the pairs of the European countries than between the pairs of the U.S. and the European countries. Since transport costs among the European countries must be less than those between the U.S. and the European countries, response to any deviation from the equality of real interest rates through the adjustment of prices (or inflation) among the European countries is relatively sensitive as compared to the U.S. Another interesting finding is that Japan barely responds to deviation from

Table 7
Error Correction Model

	BG/US	*CA/US*	*FR/US*	*GM/US*	*JP/US*	*ND/US*	*UK/US*
1960-1991							
π	-.117	-.065	-.124	-.437	-.027	-.148	-.092
	(.025)	(.018)	(.024)	(.042)	(.011)	(.028)	(.021)
η	.023	.043	.028	.017	.066	.000	.054
	(.015)	(.011)	(.014)	(.009)	(.017)	(.075)	(.022)
1973-1991							
π	-.120	-.071	-.096	-.380	-.035	-.125	-.086
	(.035)	(.024)	(.029)	(.053)	(.017)	(.034)	(.028)
η	.038	.025	.033	.004	.066	.012	.042
	(.018)	(.034)	(.014)	(.010)	(.023)	(.020)	(.025)
	BG/UK	*CA/UK*	*FR/UK*	*GM/UK*	*JP/UK*	*ND/UK*	
1960-1991							
π	-.075	-.083	-.065	-.303	-.022	-.110	
	(.019)	(.020)	(.019)	(.037)	(.010)	(.023)	
η	.061	.108	-.063	.034	.051	.023	
	(.095)	(.095)	(.019)	(.015)	(.016)	(.104)	
1973-1991							
π	-.076	-.090	-.052	-.356	-.051	-.207	
	(.026)	(.026)	(.024)	(.053)	(.021)	(.042)	
η	.052	.068	.058	.014	.016	.044	
	(.136)	(.033)	(.019)	(.021)	(.026)	(.036)	

Note: The numbers in the parenthesis are the standard errors of the estimates of π and η in the error correction models in equations (15) and (16).

the equilibrium through the adjustment of the nominal interest rate. This may be accounted for by a relatively high interest rate policy in Japan.

In summary, the average time lag to return to the equilibrium of the equality of real interest rates is about two years except for GM/US and GM/UK. Therefore, the equality of real interest rates can be considered as a long-run behavior of the real interest rates between countries.

V. CONCLUDING REMARKS

Most past empirical findings which documented evidence against the equality of real interest rates can be roughly accounted for by four main problems: Changes in the unexpected inflation rate, incomparable securities, different political risks, and measurement (or specification) problems. In order to mitigate these problems, we have chosen 90-day T-bill rates (or comparable rates) from the OECD countries which can be considered as having a similar level of political risk. Since conventional regression analysis on the non-stationary level of time series results in misleading t-values, the theory of cointegration is employed to correct the problem. The expected inflation rate is estimated by using two different approaches. First, we use the contemporaneous rates as proxies for expected inflation. Second, we estimate the expected inflation rate by using ARIMA models to decompose inflation into expected and unexpected components.

The empirical evidence in this study has strongly rejected the null hypothesis of no cointegration of real interest rates across countries. It supports the international parity condition that the real interest rate equality across countries can be considered as a long-run relationship of the real interest rates among nations. The overall results are summarized below. First, the real interest rates for comparable securities are not generally equal across countries in the sample, but the international Fisher-open condition as a long-run behavior of real interest rates has performed well over the overall sample period. Second, the equality of real interest rates as a long-run relationship is cointegrated better under the floating exchange rate regime than under the fixed exchange rate regime. Third, the unfavorable evidence found in the past empirical studies against the equality of real interest rates can be accounted for by use of different proxies for the unexpected inflation rate. Fourth, the real interest rates in the pairs between the U.K. and the other EMS countries appear to be more strongly cointegrated during the 1980s period than any other period. Fifth, the nominal interest rates and the expected inflation differentials in the pairs between the U.K. and the other European countries are more strongly cointegrated than those in the pairs between U.K. and the non-EMS during the overall period. Sixth, the economic activity in the European countries is at least as cointegrated with the U.S. economy as with the U.K. economy. Seventh, Germany responds quickly to any deviation from the real interest rate equality when there exists an imbalance in the pairs between GM/US and GM/UK. Eighth, the estimated time lag to return to the real interest rate equality is approximately two years. Consequently, the empirical evidence found in this study positively supports the real interest rate equality as a long-run behavior of real interest rates between the OECD countries.

NOTES

1. This is well explained in Stein (1962), Officer and Willet (1970) and Frenkel and Levich (1975).
2. Given the following regression model,
   $E(r^j_t) - E(r^1_t) = X_{t-1}\alpha_j + u^j_i, j = 2,3,....,m$
   where $E(r^j_t)$ = the expected real interest rate for country j u^j_t = the error term with zero mean and serially uncorrelated X_{t-1} = a variable in the available information set at t-1. α_j is regression coefficient that can be obtained by the Ordinary Least Squares (OLS). But Mishkin insists that the parameter estimates from the GLS procedure will be identical to the OLS parameter estimates because the explanatory variables in each country's regression are identical.
3. See Granger (1986); Engle and Granger (1987).
4. Transaction costs in the foreign currency markets have been found to be relatively small (see McCormick 1979). In addition, taxes should be mostly not

high enough to severely impede the efficiency of international financial transactions among the OECD nations. As a result, with a sound test procedure we are able to find the cointegration of real interest rates among the OECD nations.

5. Engle and Granger (1987) insist that any time series can have an error-correction representation if it is cointegrated of order(d,b).
6. The monthly rates of 90-day T-bill seem to be more appropriate to compare the equality of real interests across countries. But we use the monthly rates derived from average of daily rates as an alternative because of data limitation. These monthly average T-bill rates will reduce the variance found in the original daily series and are inferior to using the last days of the months or the end of the month data. The authors would like to thank an anonymous referee to clear this point.
7. The collection of international financial data usually encournters nonsynchronous data points due to differences in trading time zones. The nonsynchronous data points should not cause any serious problem because the trading hours of different national financial markets are leading or lagging only few hours to each other (on the same day).
8. For more details, see Campbell and Hamao (1992).
9. Mishkin (1984b) also explains that the usual additional second-order term is not necessary in the Fisher equation if the continuous compounded rates are assumed. Furthermore, he shows that there is no significant change in the results when holding period real returns rather than continuously compounded returns are used.
10. Gultekin (1983) uses three different estimates of the expected inflation rate; contemporaneous inflation rates, inflation rates estimated by ARIMA, and short term interest rates as predictors of inflation. But in this research, the last method cannot be used because we are testing the validity of short-term interest rate as a predictor of inflation rate. Some researchers use the rational expectation to estimate the expected inflation rate. But one can cast doubt that the explanatory variables are identical across countries.
11. This result is consistent with Rogoff (1985) and Artis and Taylor (1989). They insisted that the EMS adopted in March 1979 has contributed to the reduction of the intra-EMS nominal and real exchange rate volatility relative to the pre-EMS period. They explained that the reduction of foreign exchange rate volatility over the EMS period was not accompanied by reduced inflation, but attributed by the presence of capital controls. In turn, the importance of capital controls is reflected in the appropriate movement of interest rates to adjust the imbalance of the equality of real interest rates.

REFERENCES

Aliber, R.Z. (1974). *Attributes of national monies and the independence of national monetary policies and the international financial system.* Chicago: University of Chicago Press

Artis, M.J., & Taylor, M.P. (1989). The achievement of the european montory system. *The Economic and Social Review, 20*(2), 121–146.

Ben-Zion, U., & Weiblatt, J. (1984). Purchasing power parity, interest rate parity and the modified fisher effect in the presence of tax agreement. *Journal of International Money and Finance, 3,* 67–73.

Box, G., & Jenkins, G. *(1970). Time series analysis, forecasting and control.* San Francisco: Holden Day.

Campbell, J., & Hamao, Y. (1992). The interest rate process and the term structure of interest rate in Japan. Working Paper, Princeton University and Columbia University.

Carson, J. (1977). Short term interest rate as predictor of inflation. *American Economic Review,* 469–475.

Cumby, R., & Mushikin, F. (1986). The international linkage of interest rates: The European-US connection. *Journal of International Money and Finance, 84,* 1161–1176.

Dickey, D., & Fuller, W. (1979). Distribution of the estimators for autoregressive time series with a unit root. *Journal of American Statistical Assoc. 74,* 427–431.

Dickey, D., & Fuller, W. (1981). Likelihood ratio statistics for autoregressive time series with a unit root. *Econometrica, 49,* 1057–1072.

Dornbush, R., (1976). Expectations and exchange rate dynamics. *Journal of Political Economics, 84,* 1161–1176.

Engle, R., & Granger, C. (1987). Cointegration and error correction model. *Econometrica, 55*(2), 251–276.

Engle, R., & Yoo, B. (1987). Forecasting and testing in cointegration systems. *Journal of Econometrics,* 143–159.

Fama, E. (1970, March). Multiperiod consumption-investment decision. *American Economic Review.*

Fama, E. (1975). Short-term interest rate as predictor of inflation. *American Economic Review, 69,* 601–622.

Flood, R. (1981). *Explanation of exchange rate volatility and other empirical regulatories in some popular models of the foreign exchange market.* Carnegie-Rochester Conference Series on Public Policy 15, 219–250.

Frenkel, J.A. (1981b). Flexible exchange rates, prices, and the role of 'news'; Lessons from the 1970s. *Journal of Political Economics 89,* 665–705.

Frenkel, J., & Levich, R. (1975). Covered interest arbitrage: Unexploited profit? *Journal of Political Economics, 83,* 325–338.

Granger, C. (1986). Developments in the study of cointegrated economic variables. *Oxford Bulletin of Economics and Statistics 481,* 213–223.

Granger, C., & Weiss, A. (1983). *Time series analysis of error-correction models.* Academic Press, Inc. 255–278.

Gultekin, N. (1983). Stock market return and inflation: Evidence from other countries. *Journal of Finance, 38*(1), 49–65.

Henry, D. (1986). Econometric modelling with cointegrated variables: An overview. *Oxford Bulletin of Economics and Statistics, 48,* 201–211.

Hodrick, R. (1980). Some evidence of the equality of expected real rate across countries. Working paper no.8-79-80, GSIA, Carnegie-Mellon University.

Hodson, J., & Phelps, P. The distributed impact of price level variation of floating exchange rates. *Review of Economics and Statistics,* 59–64.

Howard, D.H., & Johnson, K.H. (1983). The behavior of monetary aggregates in major industrialized countries. *Journal of Money, Credit and Banking, 15*(2) 455–458.

Hylleberg, S., & Grayham, M. (1974). Cointegration and error correction mechanisms. *The Economic Journal, 99,* 113–125.

Long, J. (1974, July). Stock prices, inflation, and the term structure of interest rates. *Journal of Financial Economics,* 256–271.

Marston, C. (1976). Interest arbitrage in the euro-currency markets. *European Economic Review, 7,* 1–13.

McClure, J. (1988). PPP, interest parity, and the modified Fisher effect in the presence of tax agreement: Comment. *Journal of International Money and Finance 7,* 347–350.

McCormick, F. (1979). Covered interest arbitrage: Unexploited profit? Comment. *Journal of Political Economics, 87,* 418–422.

Mishikin, F. (1984a). Are real interest rate equal across countries? An empirical investigation of international parity conditions. *Journal of Finance, 34,* 1345–1357.

Mishikin, F. (1984b). The real interest rate: A multicountry empirical study, *Canadian Journal of Economics,* 283–311.

Mussa, M. (1982). A model of exchange rate determination. Journal of Political Economics, 90, 74–104.

Nelson, C., & Schwert, G. (1977). Short term interest rate as predictors of inflation rate: On testing the hypothesis that the real rate of interest is constant. *American Economic Review, 67,* 478–479.

Officer, L.H., & Willet, T.D. (1970). The covered arbritrage schedule: A critical survey of recent developments. *Journal of Money, Credit and Banking,* 2(2), 247–257.

Perron, P. (1988). Trends and random walks in macroeconomics time series. *Journal of Economic Dynamics and Control, 12,* 297–332.

Phillilps, P. (1987). Time series regression with a unit root. Econometrica, 277–302.

Prachowny, M.F. (1970). A note on interest rate parity and the supply of arbitrage funds. *Journal of Political Economy, 78*(3), 540–545.

Rogoff, K. (1985). Can exchange rate predictability be achieved without monetary congergence? Evidence from the EMS. *European Economic Review, 28*(1/2), 93–116.

Stein, J.L. (1962). Forces producing disturbances in the value of output: A replay. *Journal Political Economy, 70*(2), 170–176.

Titman, S., & Warga, A. (1989). Stock returns as predictors of interest rate and inflation. Journal of Financial Quantitative Analysis, 24, 47–59.