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Review of Applied Economics, Vol. 4, No. 1-2, (January-December 2008) : 93-112

MATURITIES, NONLINEARITIES, AND THE INTERNATIONAL TRANSMISSION OF SHORT-TERM INTEREST RATES

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This paper employs linear and nonlinear Granger causality tests to re-examine the dynamic relation between daily Eurodollar and U.S. certificates of deposit rates during the July 16, 1973 to May 1, 2006 period. This study also conducts sub-period analysis based on the switching regression technique of Goldfield and Quant (GQSRT) (1972, 1973, and 1976). The main empirical findings are: (1) Full-sample results show significant bi-directional linear causality between the EURO and CD interest rates for one-month maturities and unidirectional linear causality from the CD rate to the EURO rate for three-month and six-month maturities. Furthermore, full-sample results reveal strong bi-directional nonlinear Granger causality between the EURO and CD interest rates for all three maturities. (2) Sub-sample results based on linear tests show a unidirectional causal relation from the CD rate to the EURO rate during the first sub-period for all three maturities. During the second sub-period, however, linear tests uncover a strong bi-directional relation between the CD and the EURO rates for all three maturities. The linear results for the third sub-period reveal mostly unidirectional causality from the EURO rate to the CD for three maturities. (3) Finally, sub-sample nonlinear causality tests reveal mostly a unidirectional causality from the CD rate to the EURO rate for all three maturities during the first sub-sample, a strong significant bi-directional causality between the two rates for all three maturities during the second sub-period, and an uneven bi-directional causality between the two rates for all three maturities during the third sub-period. Overall, the results of this study show that the EURO rate's role is becoming more prominent compared to that of the CD rate.

JEL Classification: *F3; C1.*

Keywords: *Eurodollar interest rates, CD interest rates, linear and nonlinear causality, financial market integration.*

INTRODUCTION

The dynamic linkage between domestic and offshore yields on a currency is an important barometer of financial market integration especially for currencies with significant offshore

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markets. As noted by several previous researchers [see, for example, Swanson (1988)], this linkage depends on regulatory barriers and other distortions between markets and therefore responds to evolving global financial market conditions. Understanding the nature of the causal linkage between domestic and offshore yields on a given currency is important for several reasons. For example, as noted by Mougoué and Wagster (1997), the ability of international investors to maximize their returns without the knowledge of international interest rate linkages and how they respond to exogenous shocks is at best doubtful. Furthermore, business managers must understand the cross-market causality of interest rates especially if their firm's cost of capital is sensitive to interest rate changes. Finally, cross-market interest rate causality is crucial in assessing the ability of national monetary authorities to pursue exogenous monetary policies.

This study re-examines the dynamic linkage between a sample of Eurodollar interest rates and U.S. certificate of deposit rates. This re-examination is important for several reasons. First, most of the previous empirical work on the relation between Eurodollar and domestic interest rates relies mainly on traditional linear tests [see, for example, Kaen and Hachey (1983), Hartman (1984), Swanson (1987, 1988), Mougoué and Wagster (1997), Lipoids (2001), Anoruo, Ramchander, and Thiewes (2002, and Nieh and Yau (2004)]. While these previous investigations may be well suited for uncovering linear causal relations, they are not designed to uncover nonlinear causal linkages. As noted by Hsieh (1991) and Brock (1993), the recent focus on nonlinear structures in asset prices in both the financial press and the academic literature is motivated by the richer types of asset behavior that nonlinear models reveal. In addition, Hinich and Patterson (1985), Scheinkman and LeBaron (1989), Brock, Hsieh, LeBaron (1992), Fujihara and Mougoué (1997), among others, reported evidence of significant nonlinear dependence in asset returns. For these reasons, the need to investigate nonlinearities in the Eurodollar and U.S. deposit rates and understand how these nonlinearities might affect the interrelation between the two interest rates is one of the gaps this study sets out to fill. To accomplish that goal, this study not only tests for standard linear Granger causality but it also uses a nonparametric to test for nonlinear Granger causality between Eurodollar and U.S. deposit rates of differing maturities.

Second, unlike most previous research in this area that use interest rates with only one maturity, such as the three-month Eurodollar and the US CD rates [see, *e.g.*, Fung and Isberg (1992), Swanson (1987, 1988), and Mougoué and Wagster (1997)], this study covers a much greater proportion of the short end of the term structure by using one-, three-, and six-month maturity instruments. Using all three rates is more enriching because it might shed light on the important issue of whether or not the robustness of the international transmission of interest rates is maturity-dependent. In other words, this study answers the question of whether the nature of the causal relation between Eurodollar and U.S. interest rates along the maturity spectrum determines the maturity of the borrowing and lending needs of corporations, money managers, and other market participants engaged in cross-border activities. Also, analyzing interest rates with different maturities can help determine whether the ability of national monetary authorities to conduct an autonomous policy hinges on the maturity of the instruments they target.

Finally, unlike most prior related research, this study does not rely on an ad hoc approach to determine possible breakpoints in the data because such ad hoc techniques have generally led to spurious statistical results. This study circumvents this problem by using the switching regression technique of Goldfeld and Quandt (1972, 1973, and 1976) to generate its sub-samples.

DATA

This study uses daily quotations for one-, three-, and six-month U.S. certificates of deposit and Eurodollar deposit rates from July 16, 1973 to May 1, 2006. The data source is the Federal Reserve Bank's International Data tape.

Since the validity of regression results depends crucially on the stationarity of the data employed, this study uses three different unit root tests, namely the Augmented Dickey-Fuller (*ADF*)(1981), the Phillips and Perron (PP) (1988) and the Sims' (1988) tests to ascertain the stationarity of the interest rates. The null hypothesis in the Augmented Dickey-Fuller (1981) and the Phillips and Perron (1988) test is that the series contains a unit root. Because this approach has been recently criticized for lacking the power to distinguish between a unit root and weakly stationary alternatives, the Sims (1988) Bayesian posterior odds ratio procedure is also used. The Sims' test is important because Sims (1988) suggests that rejecting or failing to reject the null should result in the consideration of some set of nearby parameter settings and that a unit root test failing to address this issue may be misleading because it is difficult to distinguish between unit root models and those containing roots that lie near the unit circle.

The unit root test results are reported in Panels A and B of Table 1. The 1 per cent, 5 per cent, and 10 per cent critical values for the *ADF* and *PP* tests are, –3.43, –2.86, and –2.57, respectively. With the lone exception of the *ADF* estimate for the one-month *EURO* interest rate, all the estimates in Panel A are insignificant, indicating non-rejection of the null hypothesis of a unit

Table 1
ADF, Phillips-Perron, and Sims Unit Root Test Results for *CD* and *EURO* Rates

	1-Month Rates		*3-Month Rates*		*6-Month Rates*	
	CD	*EURO*	*CD*	*EURO*	*CD*	*EURO*
***Panel A**: Augmented Dickey Fuller (ADF) and Phillips and Perron (PP) Test Results*						
ADF	–2.4874	–2.6739[c]	–2.4824	–2.5103	–2.4057	–2.0621
PP[1]	–1.9120	–2.5166	–1.8731	–1.9528	–1.8170	–1.8545
***Panel B**: Sims' Test Results*						
Squared t[2]	4.966	5.428	4.777	5.058	4.185	4.252
Schwarz L[3]	15.931	14.342	15.862	15.311	15.699	15.527
Small Sample L[3]	9.934	8.345	9.865	9.314	9.702	9.530
Marginal α[4]	0.9796	0.9451	0.9807	0.9711	0.9844	0.9825

The sample period is from July 16, 1973 to May 1, 2006. *CD* = US certificate of Deposit Rate; *EURO* = Eurodollar Deposit Rate

[1] These results use six lagged terms for the regression and exclude the trend. Results are similar if the trend is included or if four or twelve lags are used.

[2] Squared t is the t-statistic used as the test statistic.

[3] Schwarz L is the Schwarz limit used as the asymptotic critical value for the test statistic, while small sample L is the small sample limit used as the finite sample critical value.

[4] The marginal α is the threshold value at which the posterior odds for and against the unit root are even. A small value of the marginal α indicates evidence against a unit root. For all tests, the null hypothesis is that the series contain a unit root, *i.e.*, $\rho = 1$.

[c] Indicates significance at the 10 per cent level.

root. The empirical estimates in Panel B for the Sims test also show that all the interest rate series contain a unit root. In fact, the estimates of (marginal) α's, which are the threshold values at which the posterior odds for and against a unit root are even, all exceed 97 per cent. Overall, the results in Table 1 imply that all the interest rate series contain a unit root and, therefore, are nonstationary, implying that tests for causality should rely on changes in the interest rates.

In light of the results in Table 1, several descriptive statistics for changes in the interest rate series are computed and provided in Table 2. The distributions of the daily changes for five of the six interest rate series display significant negative skewness while the distribution of the daily change for the remaining series (1-month *EURO* rate) is positively skewed. Additionally, the distributions of the daily changes for all six interest rate series display significant excess kurtosis relative to the normal distribution. Hall, Brorsen, and Irwin (1989) suggest that the excess kurtosis may be due to a possible time-varying variance in the evolution of the data.

Table 2
Summary Statistics for Changes in the *CD* and *EURO* Rates

	1-Month Rates		*3-Month Rates*		*6-Month Rates*	
	ΔCD	$\Delta EURO$	ΔCD	$\Delta EURO$	ΔCD	$\Delta EURO$
Mean($\times$100)	–0.0005	–0.0006	–0.0005	–0.0006	–0.0005	–0.0006
t-statistics	–0.4142	–0.1672	–0.4101	–0.3212	–0.3482	–0.3648
Median (%)	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
SD (%)	0.0012	0.0034	0.0013	0.0018	0.0016	0.0016
Skewness	–0.5876[a]	0.2059[a]	–0.8874[a]	–0.3099[a]	–1.4328[a]	–0.6977[a]
Kurtosis	38.8922[a]	13.4855[a]	33.1423[a]	17.6516[a]	31.6763[a]	17.5126[a]
Minimum (%)	–1.6500	–2.5600	–1.7300	–1.6200	–2.1000	–1.3200
Maximum (%)	1.2700	3.3100	1.1400	1.6300	1.1400	1.3100

The sample period is from July 16, 1973 to May 1, 2006. ΔCD = Changes in US certificate of Deposit Rate; $\Delta EURO$ = Changes in Eurodollar Deposit Rate.
SD is the Standard deviation of the mean
The *t*-statistics is for the null hypothesis that the mean equals zero.
[a]Indicates significance at the 1 per cent level.

RESEARCH DESIGN AND EMPIRICAL RESULTS

Tests for linear Granger causality

The linear causality test this study uses is attributed to Granger (1969). Other causality methodologies reported in the literature include those of Sims (1972) and Pierce and Haugh (1977). However, Granger's causality tests are employed because they are superior to Sims' (Geweke, Meese, and Dent (1983)) and perform well for small samples (Guilkey and Salemi, 1982). The Granger test involves the estimation of the following vector autoregressive (VAR) model:

$$\begin{bmatrix} \Delta CD_t \\ \Delta EURO_t \end{bmatrix} = \begin{bmatrix} \Psi_{11}(L) & \Psi_{12}(L) \\ \Psi_{21}(L) & \Psi_{22}(L) \end{bmatrix} \begin{bmatrix} \Delta CD_t \\ \Delta EURO_t \end{bmatrix} + \begin{bmatrix} \alpha_{CD} \\ \alpha_{EURO} \end{bmatrix} + \begin{bmatrix} \xi_{CD,t} \\ \xi_{EURO,t} \end{bmatrix} \quad (1)$$

where L is the lag operator, ΔCD_t and $\Delta EURO_t$ are changes in the daily U.S. *CD* deposit and Eurodollar interest rates, respectively, ψ_{ij}'s are the lag polynomials, α_{CD} and α_{EURO} are constant or intercept terms, and $\xi_{CD,\,t}$ and $\xi_{EURO,\,t}$ are uncorrelated errors for the U.S. *CD* and Eurodollar rates, respectively. The appropriate lag lengths of the VAR process in (1) are obtained using the Schwarz (1978) information criterion (SIC).

The process in (1) is an unconstrained bivariate causal system and, therefore, is appropriately used to test for linear causality between the *CD* and Eurodollar rates. For example, if $\Sigma\psi_{12,\,i} = 0$ ($\Sigma\psi_{21,\,i} = 0$), *i.e.*, if the sum of the coefficients of the lag polynomial on the Eurodollar rate (*CD* rate) in the *CD* rate (Eurodollar rate) equation is zero, then past Eurodollar rates (*CD* rates) have no influence on *CD* rates (Eurodollar rate), *i.e.*, Eurodollar rates (*CD* rates) do not "cause" *CD* rates (Eurodollar rates). By contrast, if $\Sigma\psi_{12,\,i} \neq 0$, then Eurodollar rates are said to Granger-cause *CD* rates, implying that the information set used to predict next-period *CD* rates should also include current Eurodollar rates. If $\Sigma\psi_{12,\,i} \neq 0$ and $\Sigma\psi_{21,\,i} \neq 0$ in (1), then bi directional causality exists, implying that knowledge of the past values of either time series is useful in the prediction of the other. Finally, if $\Sigma\psi_{12,\,i} = 0$ and $\Sigma\psi_{21,\,i} = 0$, then no causality exists between the Eurodollar and U.S. *CD* rate series.

The results of the linear causality tests are reported in Panels A to C of Table 3. The *t*-statistics test for the aggregate impact of each right-hand side variable on the left-hand side variable.

The results for the one-month maturity interest rates (Panel A, Table 3) show that *EURO* rates significantly linearly Granger cause *CD* interest rates at the 5 per cent level ($\Sigma\psi_{12,\,i} = 0.0220$, $t = 2.3362$). This finding is boosted by the significant χ^2 statistic (5.46), which tests the exclusion of the *EURO* rate from the *CD* rate equation. The estimates in Panel A also show that there is feedback from the one-month *CD* rate to the one-month *EURO* rate since both the sum of the coefficients ($\Sigma\psi_{21,\,i} = 4.2637$) on the *CD* rate in the *EURO* rate equation and the χ^2 statistic (341.97) are significant at the 1 per cent level. The results in Panel A of Table 3 imply that a statistically significant bi-directional linear causality exists between the one-month *EURO* and the one-month *CD* interest rates.

The empirical findings for the three-month maturity interest rates (Panel B, Table 3) are dissimilar to those for the one-month interest rates. The *EURO* rate appears to have no causal impact on the *CD* rate whereas the *CD* rate continues to exert a significant causal impact on the *EURO* rate. The sum of the coefficients on the *CD* rate in the *EURO* rate equation is positive ($\Sigma\psi_{21,\,i} = 2.0252$) and significant at the 1 percent level ($t = 10.3392$), implying that the three-month *CD* rate uni-directionally causes the three-month *EURO* rate. Once again, the fact that χ^2 is significant (305.44) in the *EURO* rate equation confirms this finding.

The results for the six-month maturity interest rates (Panel C, Table 3) are in agreement with those for the three-month rates given in Panel B of Table 3. That is, there is evidence of significant unidirectional causality from the six-month *CD* rate to the six-month *EURO* rate since the sum of the coefficients on the *CD* rate in the *EURO* rate equation is positive ($\Sigma\psi_{21,\,i} = 1.3120$) and statistically significant at the 1 per cent level ($t = 10.9048$) and the χ^2 statistic in the *EURO* equation is also highly significant.

Table 3
Test Results of Linear Causality between *CD* and *EURO* Rates

Panel A: *1-Month Rates*			
	Independent variable		
Dependent variable	ΔCD	$\Delta EURO$	χ^2-*Stat*
ΔCD	$\sum_{i=1}^{10} \Psi_{11,i} = \underset{(5.0399)^a}{0.3936}$	$\sum_{i=1}^{1} \Psi_{12,i} = \underset{(2.3362)^b}{0.0220}$	5.46^b
$\Delta EURO$	$\sum_{i=1}^{14} \Psi_{21,i} = \underset{(10.1646)^a}{4.2637}$	$\sum_{i=1}^{15} \Psi_{22,i} = \underset{(-9.2852)^a}{-3.6318}$	341.97^a
Panel B: *3-Month Rates*			
	Independent variable		
Dependent variable	ΔCD	$\Delta EURO$	χ^2-*Stat*
ΔCD	$\sum_{i=1}^{10} \Psi_{11,i} = \underset{(4.5013)^a}{0.3913}$	NA	NA
$\Delta EURO$	$\sum_{i=1}^{8} \Psi_{21,i} = \underset{(10.3392)^a}{2.0252}$	$\sum_{i=1}^{14} \Psi_{22,i} = \underset{(-7.8391)^a}{-1.5437}$	305.44^a
Panel C: *6-Month Rates*			
	Independent variable		
Dependent variable	ΔCD	$\Delta EURO$	χ^2-*Stat*
ΔCD	$\sum_{i=1}^{10} \Psi_{11,i} = \underset{(3.6710)^a}{0.3214}$	NA	NA
$\Delta EURO$	$\sum_{i=1}^{5} \Psi_{21,i} = \underset{(10.9048)^a}{1.3120}$	$\sum_{i=1}^{5} \Psi_{22,i} = \underset{(-8.6392)^a}{-0.9049}$	235.37^a

The sample period is from July 16, 1973 to May 1, 2006. ΔCD = Changes in US certificate of Deposit Rate; $\Delta EURO$ = Changes in Eurodollar Deposit Rate.

1. *t*-statistics are reported in parentheses below the sum of the estimated coefficients. The sum represents the cumulative effect. For a given sum of coefficients *s*, the *t*-statistic is calculated as $t = s/\sigma_s$, where $s = \Sigma^n a_i$ and n = number of lags on the independent variable whose impact is being investigated. For example, if $n = 3$ then, $s = \Sigma^n a_i = a_1 + a_2 + a_3$ and

$$\sigma_s = \sigma_{(a_1 + a_2 + a_3)} = \sqrt{\sigma^2_{a_1} + \sigma^2_{a_2} + \sigma^2_{a_3} + 2\sigma_{a_1 a_2} + 2\sigma_{a_2 a_3} + 2\sigma_{a_1 a_3}}$$

2. χ^2-Stat is the chi-square statistic testing for the joint significance of the lags of the large (small) firms in the equation of the small (large) firms and is calculated as $\chi^2_{pi} = (c - c\hat{\beta})' [C \Sigma_X^{C'}]^{-1} (c - C\hat{\beta})$, $i = 2, 4$; where c is a $(p \times 1)$ vector of known constants, C is a $(p \times k)$ hypothesis design matrix of known constants, β is a $(k \times 1)$ vector of the regression coefficients, and Σ_X is the estimated covariance matrix of the regression coefficients.

[a], [b], and [c] indicate significance at the 1 per cent, 5 per cent, and 10 per cent respectively.

Test for Nonlinear Dependence in the Univariate Interest Rate Series

The empirical results based on model (1) rely on the untested presumption that any causal relation between the *EURO* and the *CD* rates is of a *linear* nature. Another problem with relying exclusively on model (1) is that the interest rate series being investigated may exhibit significant

nonlinearities that, in turn, may yield significant *nonlinear* causal linkages. To investigate this last possibility, this paper applies a formal *nonlinearity* test to the *EURO* and *CD* rate series.

The nonlinear dependence test is based on the work of Brock, Dechert, and Scheinkman (1987) who propose a test (*BDS*) for deviations from independent and identically distributed (iid) behavior. For a sequence of observations $\{x_t : t = 1, ..., T\}$ that are i.i.d., an m-dimensional vector $X_t^m = (x_t, x_{t+1}, ..., x_{t+m-1})$ can be formed. The test computes a statistic based on the correlation integral defined by:

$$C_m(\varepsilon, T) = \frac{2}{n(n-1)} \sum_{t<s} I_\varepsilon\,(X_t^m, X_s^m)$$

where $n = T - m + 1$ and $I_e(X_t^m - X_s^m)$ is an indicator function defined as,

$$I_\varepsilon\,(X_t^m, X_s^m) = \begin{cases} 1, & \text{if} \quad \left\| X_t^m - X_s^m \right\| < \varepsilon \\ 0, & \text{otherwise} \end{cases}$$

and || || denotes the maximum norm.

The test statistic is given by:

$$BDS_m(\varepsilon, T) = \frac{n^{1/2} \times \left\lfloor C_m(\varepsilon, T) - C_1(\varepsilon, T)^m \right\rfloor}{\sigma_m(\varepsilon, T)}$$

Under the null hypothesis that $\{x_t\}$ is i.i.d. the term $n^{1/2} \times [C_m(e, T) - C_1(\varepsilon, T)^m]$ has a normal limiting distribution with mean zero and standard deviation $\sigma_m(\varepsilon, T)$. The null hypothesis of a random i.i.d. process is rejected if the probability of any two m-histories being close together exceeds the mth power of the probability of any two points being close together.

Table 4 provides the results of the *BDS* test applied to the six interest rate series. The 1 per cent, 5 per cent, and 10 per cent critical values of the *BDS* test are 2.575, 1.960, and 1.645, respectively. The empirical estimates in Table 4 are all positive and statistically significant, leading to a rejection of the null hypothesis of i.i.d. behavior for the interest rates. This rejection implies that the interest rate series are characterized by significant nonlinearities, opening up the possibility that a complex nonlinear relation might also tie the *CD* and *EURO* rates together.

Test for Nonlinear Causality

Using the *BDS* test statistic, Section 3.2 demonstrates that the univariate *EURO* and *CD* rates exhibit significant nonlinearities, implying that a complete investigation of the causal relation between the two interest rates should embody tests for both linear and nonlinear dependence.

Baek and Brock (1992) propose a nonparametric statistical technique for uncovering nonlinear causal relations. Consider the two time series of $\{EURO_t\}$ and $\{CD_t\}$. Let the m-length lead vectors of $EURO_t$ and CD_t be denoted as $EURO_t^m$ and CD_t^m respectively and the L_{euro}-length and L_{cd}-length lag vectors of $EURO_t$ and CD_t, respectively, by $EURO_{t-L_{euro}}^{L_{euro}}$ and $CD_{t-L_{cd}}^{L_{cd}}$. For

Table 4
BDS Statistics for Changes in the *CD* and *EURO* Rates

		1-Month Rates		*3-Month Rates*		*6-Month Rates*	
M	ε/σ	ΔCD	$\Delta EURO$	ΔCD	$\Delta EURO$	ΔCD	$\Delta EURO$
2	0.5	46.3227	62.9364	45.6656	49.4715	43.0557	39.6016
3	0.5	55.7696	82.8594	56.6002	64.2454	53.4254	51.4101
4	0.5	64.2072	101.6193	67.7425	79.5856	63.2523	62.5636
5	0.5	74.1646	124.6892	81.3915	99.5479	75.4854	76.8344
6	0.5	86.6702	155.3425	99.7208	126.9626	92.1092	95.2640
2	1.0	39.1807	56.1658	40.6918	43.8201	40.8931	37.1334
3	1.0	45.4593	70.0720	47.2534	52.9105	47.0894	45.2089
4	1.0	48.3289	80.5714	51.2222	60.1121	51.0253	50.7051
5	1.0	50.4802	91.6021	54.6922	68.0420	54.8987	55.9264
6	1.0	52.5376	104.4418	58.1459	77.4321	59.1139	61.7975
2	1.5	36.9011	49.6089	37.1494	39.5356	39.7184	36.0921
3	1.5	42.7694	58.9966	43.2000	47.0995	44.2710	42.7235
4	1.5	44.5995	64.8926	45.4497	51.4152	46.5262	46.6598
5	1.5	45.4312	69.6297	47.1903	55.8786	48.4984	49.7105
6	1.5	46.1815	74.2332	48.4663	60.5831	50.4968	52.4751
2	2.0	34.3798	44.6705	35.7426	33.2830	37.9501	34.3090
3	2.0	40.2198	51.5859	40.9372	39.5260	41.8619	39.7625
4	2.0	41.4613	55.7959	42.8967	42.5596	43.6103	42.9507
5	2.0	41.7450	58.8399	43.9727	44.8406	44.6677	45.2771
6	2.0	41.9277	61.7045	44.6798	46.9379	45.6821	46.8214

The sample period is from July 16, 1973 to May 1, 2006. ΔCD = Changes in US certificate of Deposit Rate; $\Delta EURO$ = Changes in Eurodollar Deposit Rate. The numbers in the table are the *BDS* test statistics and are computed as,

$$BDS_m(\varepsilon, T) = \frac{n^{1/2} \times \lfloor C_m(\varepsilon, T) - C_1(\varepsilon, T)^m \rfloor}{\sigma_m(\varepsilon, T)}$$

The *BDS* statistic has a standard normal limiting distribution. The null hypothesis of a random iid process is rejected if the probability of any two *m*-histories being close together exceeds the mth power of the probability of any two points being close together, where *m* is the embedding dimension.

known values of m, L_{euro}, and $L_{cd} \geq 1$ and for $d > 0$, *CD* does not strictly Granger cause *EURO* if:

$$\begin{aligned}&\text{Prob}\left(\left\|EURO_t^m - EURO_s^m\right\| < d, \left\|EURO_{t-L_{euro}}^{L_{euro}} - EURO_{s-L_{euro}}^{L_{euro}}\right\| < d, \left\|CD_{t-L_{cd}}^{L_{cd}} - CD_{s-L_{cd}}^{L_{cd}}\right\| < d\right)\\ &= \text{Prob}\left(\left\|EURO_t^m - EURO_s^m\right\| < d, \left\|EURO_{t-L_{euro}}^{L_{euro}} - EURO_{s-L_{euro}}^{L_{euro}}\right\| < d\right) \qquad (2)\end{aligned}$$

where Prob(·) denotes probability and || || denotes the maximum norm. The probability on the left hand side (*LHS*) of equation (2) is the conditional probability that two arbitrary *m*-length lead vectors of $\{EURO_t\}$ are within a distance d of each other, given that the corresponding L_{euro}-length lag vectors of $\{EURO_t\}$ and L_{cd}-length lag vectors of $\{CD_t\}$ are within d of each other. The probability on the right hand side (*RHS*) of equation (2) is the conditional probability that two arbitrary m-length lead vectors of $\{EURO_t\}$ are within a distance d of each other, given that their corresponding L_{euro}-length lag vectors are within a distance d of each other. The strict noncausation condition in equation (2) is expressed as

$$\frac{CI1(m + L_{euro}, L_{cd}, d)}{CI2(L_{euro}, L_{cd}, d)} = \frac{CI3(m + L_{euro}, d)}{CI4(L_{euro}, d)} \tag{3}$$

where $CI(\cdot)$ are the correlation-integral estimators of the joint probabilities. Assuming that $\{EURO_t\}$ and $\{CD_t\}$ are strictly stationary, weakly dependent, and obey the conditions of Denker and Keller (1983), if $\{CD_t\}$ does not strictly Granger cause $\{EURO_t\}$ then,

$$\sqrt{n}\left(\frac{CI1(m + L_{euro}, L_{cd}, d, n)}{CI2(L_{euro}, L_{cd}, d, n \sim N(0, \sigma^2(m, L_{euro}, L_{cd}, d))} - \frac{CI3(m + L_{euro}, d, n)}{CI4(L_{euro}, d, n)}\right) \tag{4}$$

Hiemstra and Jones (1994) show that a consistent estimator of the variance $\sigma^2(m, L_{euro}, L_{cd}, d)$ in equation (4) is $\hat{\sigma}^2(m + L_{euro}, L_{cd}, d, n) = \hat{\delta}(n)\hat{\Sigma}(n)\hat{\delta}(n)'$.

The test statistics given in equations (3) and (4) are applied to the two estimated residual series from the VAR model (1), $\{\hat{\xi}_{CD,t}\}$ and $\{\hat{\xi}_{EURO,t}\}$. The null hypothesis is that $\{CD_t\}$ does not nonlinearly strictly Granger cause $\{EURO_t\}$, and that equation (4) holds for all m, L_{euro}, and $L_{cd} \geq 1$ and for all $d > 0$. By removing linear predictive power with a linear VAR model, any remaining incremental predictive power of one residual series for another can be considered nonlinear predictive power (see Baek and Brock (1992)).

Values for the lead length m, the lag lengths L_{euro} and L_{cd}, and the scale parameter d must be selected in order to conduct the Baek and Brock test. However, because there is no literature on the appropriate way to specify optimal values for lag lengths and the scale parameter in nonlinear causality tests, this study relies on the Monte Carlo results found in Hiemstra and Jones (1993) by setting, for all cases, the lead length to $m = 1$ and $L_{euro} = L_{cd}$. Common lag lengths of 1 to 6 lags are also used. Additionally, for all cases, the test is applied to standardized series using a common scale parameter $d = 0.4 \times \sigma$, where $\sigma = 1.0$ denotes the standard deviation of the standardized time series.

The empirical results for nonlinear Granger causality tests are reported in Table 5. DIFF and NORM, respectively, denote the difference between the two conditional probabilities in equation (3) and the standardized test statistic in equation (4). Under the null hypothesis of nonlinear Granger noncausality, the NORM test statistic is asymptotically distributed $N(0, 1)$.

Table 5 reveals significant bi-directional nonlinear Granger causality between the *EURO* and the *CD* interest rates. This result holds for all six common lag lengths used in the tests. The results for the one-month maturity interest rates (Panel A) show that the nonlinear causality from the *EURO* interest rate to the *CD* rate is significant at the one percent level for all the lags examined, while the nonlinear causality from *CD* to *EURO* is also significant at the one percent level, although the size of the NORM statistics is less prominent. The results for the three-month maturity interest rates (Panel B) and the six-month interest rates (Panel C) indicate that all of the standardized test statistics (NORM) are significant at the one percent level, implying strong evidence of bi-directional nonlinear Granger causality between the *CD* and the *EURO* interest rates.

Overall, the results of the nonlinear Granger causality tests for the one-month interest rates concur with those of the linear tests in that they reveal a significant bi-directional linear and

Table 5
Results of Nonlinear Granger Causality Test between the *CD* and *EURO* Rates

Panel A: *1-Month Rates*

	H_0 : *CD Does Not Cause EURO*		H_0: *EURO Does Not Cause CD*	
$L_{euro} = L_{cd}$	*DIFF*	*NORM*	*DIFF*	*NORM*
1	0.0037	5.2088[a]	0.0080	5.4813[a]
2	0.0050	3.9139[a]	0.0128	6.2314[a]
3	0.0044	3.6881[a]	0.0155	6.1228[a]
4	0.0032	3.4231[a]	0.0158	6.4808[a]
5	0.0016	3.2508[a]	0.0143	6.1924[a]
6	0.0016	2.8080[a]	0.0133	5.9639[a]

Panel B: *3-Month Rates*

	H_0 : *CD Does Not Cause EURO*		H_0: *EURO Does Not Cause CD*	
$L_{euro} = L_{cd}$	*DIFF*	*NORM*	*DIFF*	*NORM*
1	0.0035	4.5131[a]	0.0034	4.3007[a]
2	0.0052	4.5368[a]	0.0063	5.3159[a]
3	0.0064	5.1946[a]	0.0095	6.2912[a]
4	0.0067	5.3156[a]	0.0129	6.3463[a]
5	0.0061	4.3023[a]	0.0126	5.7924[a]
6	0.0047	4.0983[a]	0.0114	6.2752[a]

Panel C: *6 Month Rates*

	H_0 : *CD Does Not Cause EURO*		H_0: *EURO Does Not Cause CD*	
$L_{euro} = L_{cd}$	*DIFF*	*NORM*	*DIFF*	*NORM*
1	0 .0042	4.0232[a]	0.0034	4.9646[a]
2	0.0064	5.0657[a]	0.0050	5.3331[a]
3	0.0085	6.0473[a]	0.0064	5.0857[a]
4	0.0086	5.6228[a]	0.0090	5.2199[a]
5	0.0080	5.6816[a]	0.0092	5.9191[a]
6	0.0080	5.6070[a]	0.0109	6.4531[a]

The sample period is from July 16, 1973 to May 1, 2006. ΔCD = Changes in US certificate of Deposit Rate; $\Delta EURO$ = Changes in Eurodollar Deposit Rate. This table reports the results of the Baek and Brock nonlinear Granger causality test applied to the estimated VAR residuals from equations (1) and (2). $L_{euro} = L_{cd}$ denotes the number of lags on the residuals series used in the test. In all cases, the tests are applied to the unconditional standardized series. The lead length, m, is set to unity, and the length scale, d, is set to 0.5. DIFF and NORM, respectively, denote the difference between the two conditional probabilities in equation (3) and the standardized test statistic in equation (4). Under the null hypothesis of nonlinear Granger noncausality, the test statistic is asymptotically distributed $N(0, 1)$. [a], [b], and [c] indicate significance at the 1 per cent, 5 per cent, and 10 per cent levels, respectively.

nonlinear causality between the *CD* and the *EURO* interest rates during the July 16, 1973 to May 1, 2006 period. These findings for the one-month interest rates are similar to those reported earlier in the literature (see, for example, Mougoué and Wagster (1997) and Fung and Isberg (1992)).

The nonlinear causality test results for the three-month and the six-month rates are dissimilar to those from the linear causality tests. As noted earlier, linear tests uncover only a unidirectional

causality running from the three-month (six-month) *CD* rate to the three-month (six-month) *EURO* rate whereas nonlinear tests reveal a significant bi-directional causal relation between the three-month (six-month) *CD* rate and the three-month (six-month) *EURO* rate.

Further Investigation

This section examines whether the causal relation between the Eurodollar markets and the U.S. markets has changed over time. This examination is important because financial markets in general, and the Eurodollar markets in particular, have witnessed several significant developments during the time period covered by this study. First, the size of the Eurodollar markets has grown significantly over the years, with the growth accelerating since 1984. Second, trading in the Eurodollar futures contracts, which began in December 1981 on the International Money Market of the Chicago Mercantile Exchange, has been accelerating, jumping from 909,100 contracts in 1983 to approximately 3,6363,000 million contracts by 1984 and has been rising continually. In fact, by early 2006, Eurodollar futures and options had reached over 32 million in open interest and 2.8 million in average daily trading volume. Other fundamental changes took place in the Eurodollar market in the 1990s including a dramatic fall-off in foreign exchange trading due to the introduction of the Euro, banks' increasing links with financial firms like hedge funds and securities houses, and the consolidation in the banking industry (McGuire, 2004). These changes may have impacted tremendously the Eurodollar market, as claims out of British banks increasingly shifted towards non-bank borrowers in the US. Finally, the growth of interest rate swap markets combined with Eurodollar futures and option trading has increased liquidity in the Eurodollar markets by enhancing arbitrage opportunities. This increased liquidity may, in turn, impact the causal relations between the Eurodollar markets and U.S. markets.

Rather than imposing a set of prior beliefs as to when interest-rate regime changes might have occurred, this study examines the temporal variability of the parameters of model (1) using Goldfeld and Quandt's switching regression technique (GQSRT). The GQSRT identifies changes in model parameters by simultaneously specifying the number of effective regimes, the parameter values in each regime, the switch dates at which one regime supplants another and the gradualness of each regime. The GQSRT is illustrated for an *n*-regime specification of the *CD* equation in model (1) as follows,

$$\Delta CD_t = \alpha_{CDk} + \Psi_{11}(L)\,\Delta CD_t + \Psi(L)\,\Delta EURO_t + \xi_{CDkt}, \quad k = 1, ..., n \tag{5}$$

In (5), k indexes the n regimes and ξ_{CDkt} is a noise term that is assumed to be normally distributed. Since the number and timing of regime shifts is not known a priori, equation (5) cannot be estimated as a dummy-variable regression. An extraneous classifying variable V_j with unknown cutoff values, V_j^*, drives the transition between regimes. To convert equation (5) into a single regression model, the GQSRT employs a series of transitional dummy variables, D_{tj}. If the observations are from n regimes, there exist $n-1$ switch dates V_j^* and $n-1$ gradualness parameters σ_j^*. The $n-1$ sets of variables D_{tj} may be approximated by a normal cumulative density function (*CDF*) with mean V_j^* and variance σ_j^{*2} as follows,

$$D_{tj} = \int_{-\infty}^{V_t} [(2\pi)^{1/2}\sigma_j^*]^{-1} \exp\left(\frac{-1}{2}\left(\frac{\eta - V_j^*}{\sigma_j^*}\right)^2\right) d\eta \tag{6}$$

where j runs from 1 to $n - 1$ and the endpoint values are specified as $D_{tn} = 0$ and $D_{t0} = 1$. The estimate of σ_j^* provides knowledge about the smoothness of the structural change. The smaller σ_j^* is, the more abrupt the transition between regimes. If σ_j^* is significantly different from zero, the hypothesis that the structural change is abrupt in the vicinity of V_j^* should be rejected. After multiplying the equation representing the gth regime in (5) $\lambda_{tg} = \prod_{j=0}^{g-1} D_{tj} \prod_{j=g}^{n} (1 - D_{tj})$ it with the consolidated equation to be estimated is derived by adding up the resulting equations for the n regimes as follows,

$$\sum_{g=1}^{n} \Delta CD_t \lambda_{tg} = \sum_{g=1}^{n} \left(\left(\alpha_{CDg} + \Psi_{11}(L)\Delta CD_t + \Psi_{12}(L)\Delta EURO_t + \xi_{CDgt} \right)(\lambda_{tg}) \right) \tag{7}$$

The log-likelihood function for an n-regime specification is obtained under the assumption that $\Delta CD_t \lambda_{tg}$ is normally distributed with mean and standard deviation specified as, $\sum_{g=1}^{n} \left((\alpha CD_g + \Psi_{11}(L)\Delta CDt + \Psi_{12}(L)\Delta EURO_t + \xi CDgt)(\lambda_{tg}) \right)$ and $\sigma_{\xi t}^2 = \sum_{g=1}^{n} (\sigma_{\xi g}^2)(\lambda_{tg}^2)$ respectively. The resulting log-likelihood function is defined as,

$$Ln(L) = -\frac{T}{2} Ln(2\pi) - \frac{1}{2} \sum_{t=1}^{T} \left[\frac{\left(\sum_{g=1}^{n} \xi_{tg} \lambda_{tg} \right)^2}{\sum_{g=1}^{n} (\sigma_{\xi g}^2)(\lambda_{tg}^2)} \right] - \frac{1}{2} \sum_{t=1}^{T} \sum_{g=1}^{n} Ln(\sigma_{\xi g}^2)(\lambda_{tg}) \tag{8}$$

The maximum-likelihood estimates of the parameters of equation (5), the mean switch dates V_j^*, and the gradualness parameters σ_j^*, are obtained by maximizing equation (8) with respect to the unknown parameters.

To detect the number of regimes during the period of analysis, the maximum-likelihood values for one-regime (L_1) and two-regime (L_2) models using equation (8) are first obtained. The null hypothesis of no regime switch is tested against the alternative hypothesis that two regimes are present (implying one switch point). The log-likelihood-ratio test statistic is given as $-2Ln(L_1/L_2)$ and is asymptotically distributed as a χ^2 with degrees of freedom equal to the number of restrictions under the null hypothesis. If the null hypothesis is rejected in favor of the alternative, then the possibility of three regimes is tested. Once again, if the null hypothesis is rejected, the possibility of four regimes is investigated. This process continues until failure to reject the null hypothesis occurs.

The maximization of the likelihood functions in this study necessitates numerical optimization. Two routines from Princeton University's GQOPT package are used: the NMSIMP (Nelson-Mead Simplex Method) are used to secure starting points, which are then be employed as input into the second routine known as GRADX (an algorithm that uses the quadratic hill-climbing procedure) to produce parameter estimates and t-values.

The empirical results, shown in Table 6, identify August 12, 1985 and October 27, 1995, September 24, 1985 and November 28, 1995, and November 13, 1985 and January 4, 1996 as the most likely break points in model (1) for the one-month, the three-month, and the six-month interest rates, respectively. These dates are used to break up the sample into three non-overlapping sub-samples. Tests for linear and nonlinear causality are re-conducted based on these split samples.[1]

Table 6
Results of Goldfeld-Quandt Tests Identifying the Number of Switches, Most Likely Switch Dates, and Gradualness of Switches in the Granger Causality Model

Panel A: *Likelihood-Ratio Test Results*

	$-2Ln(L^r/L^u)^2$		
Regime[1]	*1-Month Rates*	*3-Month Rates*	*6-Month Rates*
R_1 vs. R_2	131.55[a]	110.39[a]	97.09[a]
R_2 vs. R_3	71.63[a]	46.17[a]	53.05[a]
R_3 vs. R_4	6.77	5.09	4.93

Panel B*: Switch Dates and Gradualness*[3]

	Switch Point (V^*)	*Implied Switch Date*	*Gradualness Parameter (σ^*)*
1-Month	2961	08/12/1985	1.48
	(112.5)		(2.33)
	5522	10/27/1995	1.08
	(89.3)		(0.80)
3-Month	2999	09/24/1985	0.98
	(131.3)		(1.22)
	5543	11/28/1995	0.98
	(77.0)		(0.25)
6-Month	3032	11/13/1985	1.19
	(155.1)		(1.55)
	5568	01/04/1996	1.01
	(95.9)		(0.69)

[1] R_i represents the hypothesis that the Granger causality model switches exactly $i - 1$ times during the 7/16/73 to 5/1/2006 period.

[2] The test statistic is $-2Ln(L^r/L^u)$, where L^r and L^u are restricted and unrestricted maximum likelihood values. The 1 per cent and 5 per cent critical values for 2 degrees of freedom are 9.21034 and 5.99146, respectively.

[3] Asymptotic standard errors are in parentheses.

[a] indicates significance at the 1 per cent level.

Table 7 shows the results of the empirical tests for linear causality for the three sub-periods. The results for Period 1 (1973-1985) reveal a significant unidirectional causal relation from the *CD* rate to the *EURO* rate during for all three maturities. For all three maturities, the aggregate impact of the *CD* rate on the *EURO* rate is highly significant. Furthermore, the χ^2 statistics used to test for the exclusion of the *CD* rate from the *EURO* rate equation are highly significant. The findings for Period 2 (1986-1995) show a strongly significant bi-directional causal relation between the *CD* rate and the *EURO* for all the three maturities studied. Period 3 (1996-2006) results are somewhat surprising as they reveal that the pattern of causality between the *CD* rate

Table 7
Test Results of Linear Causality between CD and EURO Rates: Sub-Sample Analysis

Period 1: *1973-1985*
Panel A: *1-Month Rates*

Dependent Variable	*Independent Variable* ΔCD	$\Delta EURO$	χ^2*-Stat*
ΔCD	$\Sigma\Psi_{11,i}$ = 0.4174 (5.9431)[a]	NA	NA
$\Delta EURO$	$\Sigma\Psi_{21,i}$ = 2.5124 (14.1905)[a]	$\Sigma\Psi_{22,i}$ = –2.0065 (–14.6507)[a]	280.70[a]

Period 2: *1986-1995*
Panel A: *1-Month Rates*

Dependent Variable	*Independent Variable* ΔCD	$\Delta EURO$	χ^2*-Stat*
ΔCD	$\Sigma\Psi_{11,i}$ = –0.8866 (–3.0273)[a]	$\Sigma\Psi_{12,i}$ = 1.0068 (3.7503)[a]	16.75[a]
$\Delta EURO$	$\Sigma\Psi_{21,i}$ = 0.6881 (3.6271)[a]	$\Sigma\Psi_{22,i}$ = –0.6972 (–3.2674)[a]	29.47[a]

Period 3: *1996-2006*
Panel A: *1-Month Rates*

Dependent Variable	*Independent Variable* ΔCD	$\Delta EURO$	χ^2*-Stat*
ΔCD	$\Sigma\Psi_{11,i}$ = 0.1860 (3.3914)[a]	$\Sigma\Psi_{12,i}$ = 0.5986 (5.9974)[a]	17.70[a]
$\Delta EURO$	$\Sigma\Psi_{21,i}$ = 0.6501 (1.6002)	$\Sigma\Psi_{22,i}$ = –0.3997 (–2.5977)[a]	7.60

Panel B: *3-Month Rates*

Dependent Variable	*Independent Variable* ΔCD	$\Delta EURO$	χ^2*-Stat*
ΔCD	$\Sigma\Psi_{11,i}$ = 0.3305 (4.3593)[a]	NA	NA
$\Delta EURO$	$\Sigma\Psi_{21,i}$ =1.6211 (11.0724)[a]	$\Sigma\Psi_{22,i}$ = –1.1807 (–9.1872)[a]	249.68[a]

Panel B: *3-Month Rates*

Dependent Variable	*Independent Variable* ΔCD	$\Delta EURO$	χ^2*-Stat*
ΔCD	$\Sigma\Psi_{11,i}$= 01465 (3.4576)[a]	$\Sigma\Psi_{12,i}$ = 0.1965 (3.5662)[a]	97.94[a]
$\Delta EURO$	$\Sigma\Psi_{21,i}$= 0.9069 (11.8089)[a]	$\Sigma\Psi_{22,i}$ = –0.6477 (–9.0320)[a]	181.72[a]

Panel B: *3-Month Rates*

Dependent Variable	*Independent Variable* ΔCD	$\Delta EURO$	χ^2*-Stat*
ΔCD	$\Sigma\Psi_{11,i}$ = 0.1834 (2.9419)[a]	$\Sigma\Psi_{12,i}$ = 0.4119 (4.0007)[a]	79.75[a]
$\Delta EURO$	$\Sigma\Psi_{21,i}$ = 0.9809 (1.9157)[c]	$\Sigma\Psi_{22,i}$ = –0.7207 (–7.2468)[a]	4.80[c]

Panel C: *6-Month Rates*

Dependent Variable	*Independent Variable* ΔCD	$\Delta EURO$	χ^2*-Stat*
ΔCD	$\Sigma\Psi_{11,i}$ = 0.0802 (2.0943)[b]	NA	NA
$\Delta EURO$	$\Sigma\Psi_{21,i}$ = 1.3314 (8.7103)[a]	$\Sigma\Psi_{22,i}$ = –0.8737 (–7.4378)[a]	173.58[a]

Panel C: *6-Month Rates*

Dependent Variable	*Independent Variable* ΔCD	$\Delta EURO$	χ^2*-Stat*
ΔCD	$\Sigma\Psi_{11,i}$ = 0.0731 (2.3226)[b]	$\Sigma\Psi_{12,i}$ = 0.2518 (7.7732)[a]	255.98[a]
$\Delta EURO$	$\Sigma\Psi_{21,i}$ = 0.7406 (10.5818)[a]	$\Sigma\Psi_{22,i}$ = –0.4160 (–11.8571)[a]	316.76[a]

Panel C: *6-Month Rates*

Dependent Variable	*Independent Variable* ΔCD	$\Delta EURO$	χ^2*-Stat*
ΔCD	$\Sigma\Psi_{11,i}$ = 0.0296 (0.3870)	$\Sigma\Psi_{12,i}$ = 0.2089 (3.1895)[a]	11.05[a]
$\Delta EURO$	$\Sigma\Psi_{21,i}$ = 0.8404 (1.8785)[c]	$\Sigma\Psi_{22,i}$ = –0.5529 (–5.9337)[a]	5.26[c]

Notes: See Table 3

Table 8
Results of Nonlinear Granger Causality Test Between the *CD* and *EURO* Rates: Sub-Sample Analysis

	Period 1: *1973-1985* ***Panel A:*** *1-Month Rates*					***Period 2:*** *1986-1995* ***Panel A:*** *1-Month Rates*					***Period 3:*** *1996-2006* ***Panel A:*** *1-Month Rates*			
	H_0 : CD does not cause EURO		*H_0 : EURO does not cause CD*			*H_0 : CD does not cause EURO*		*H_0 : EURO does not cause CD*			*H_0 : CD does not cause EURO*		*H_0: EURO does not cause CD*	
$L_{euro} = L_{cd}$	*DIFF*	*NORM*	*DIFF*	*NORM*	$L_{euro} = L_{cd}$	*DIFF*	*NORM*	*DIFF*	*NORM*	$L_{euro} = L_{cd}$	*DIFF*	*NORM*	*DIFF*	*NORM*
1	0.0048	3.9068[a]	0.0144	1.9567[b]	1	0.0048	4.9967[a]	0.0144	4.9567[a]	1	0.0005	1.7603[b]	0.0012	2.6529[a]
2	0.0140	3.3088[a]	0.0286	1.0450	2	0.0140	5.3096[a]	0.0286	6.0450[a]	2	0.0013	1.3573[c]	0.0021	3.3173[a]
3	0.0142	3.7887[a]	0.0382	1.0767	3	0.0142	4.7777[a]	0.0382	6.0767[a]	3	0.0014	1.7664[b]	0.0033	2.0342[b]
4	0.0127	4.0003[a]	0.0392	2.0140[b]	4	0.0127	4.6760[a]	0.0392	6.0140[a]	4	0.0011	1.3997[c]	0.0029	2.1522[b]
5	0.0096	2.0367[b]	0.0343	1.1250	5	0.0096	4.0740[a]	0.0343	5.1250[a]	5	0.0013	1.5144[c]	0.0034	2.4293[a]
6	0.0086	2.7384[a]	0.0333	1.3272	6	0.0086	3.8540[a]	0.0333	4.8272[a]	6	0.0015	1.6849[b]	0.0033	2.3578[a]
	Panel B: *3-Month Rates*					***Panel B:*** *3-Month Rates*					***Panel B:*** *3-Month Rates*			
	H_0 : CD does not cause EURO		*H_0 : EURO does not cause CD*			*H_0 : CD does not cause EURO*		*H_0 : EURO does not cause CD*			*H_0 : CD does not cause EURO*		*H_0: EURO does not cause CD*	
$L_{euro} = L_{cd}$	*DIFF*	*NORM*	*DIFF*	*NORM*	$L_{euro} = L_{cd}$	*DIFF*	*NORM*	*DIFF*	*NORM*	$L_{euro} = L_{cd}$	*DIFF*	*NORM*	*DIFF*	*NORM*
1	0.0084	3.2706[a]	0.0090	1.0331	1	0.0084	5.1507[a]	0.0090	5.0201[a]	1	0.0001	1.4171[c]	0.0007	2.1899[b]
2	0.0154	3.6451[a]	0.0180	1.9729[b]	2	0.0154	5.6451[a]	0.0180	5.9498[a]	2	0.0001	1.4013[c]	0.0009	2.6239[a]
3	0.0203	4.0541[a]	0.0293	1.1182	3	0.0203	6.2289[a]	0.0293	7.0211[a]	3	0.0002	1.9751[b]	0.0014	2.6995[a]
4	0.0214	2.2889[b]	0.0382	0.7846	4	0.0214	6.3629[a]	0.0382	7.4496[a]	4	0.0003	2.3860[a]	0.0020	2.4870[a]
5	0.0188	2.7551[a]	0.0377	0.9974	5	0.0188	5.6943[a]	0.0377	6.9123[a]	5	0.0004	2.5419[a]	0.0018	2.1239[b]
6	0.0170	3.2894[a]	0.0361	1.2964	6	0.0170	5.4715[a]	0.0361	7.1627[a]	6	0.0005	2.3507[a]	0.0024	2.5816[a]
	Panel C: *6-Month Rates*					***Panel C:*** *6-Month Rates*					***Panel C:*** *6-Month Rates*			
	H_0 : CD does not cause EURO		*H_0 : EURO does not cause CD*			*H_0 : CD does not cause EURO*		*H_0 : EURO does not cause CD*			*H_0 : CD does not cause EURO*		*H_0: EURO does not cause CD*	
$L_{euro} = L_{cd}$	*DIFF*	*NORM*	*DIFF*	*NORM*	$L_{euro} = L_{cd}$	*DIFF*	*NORM*	*DIFF*	*NORM*	$L_{euro} = L_{cd}$	*DIFF*	*NORM*	*DIFF*	*NORM*
1	0.0113	2.6693[a]	0.0090	2.0007[b]	1	0.0113	5.3247[a]	0.0090	6.0305[a]	1	0.0000	2.1722[b]	0.0000	2.9859[a]
2	0.0188	3.0752[a]	0.0162	1.3962	2	0.0188	6.1535[a]	0.0162	6.0336[a]	2	0.0001	1.9484[b]	0.0001	2.7158[a]
3	0.0265	2.1117[b]	0.0214	1.0772	3	0.0265	7.1675[a]	0.0214	6.1191[a]	3	0.0003	2.2108[b]	0.0004	2.6024[a]
4	0.0289	3.7431[a]	0.0279	1.3087	4	0.0289	6.9902[a]	0.0279	6.4407[a]	4	0.0004	2.0695[b]	0.0007	2.4178[a]
5	0.0290	3.8663[a]	0.0293	0.9969	5	0.0290	6.9052[a]	0.0293	6.6272[a]	5	0.0006	2.2155[b]	0.0016	2.6662[a]
6	0.0283	2.6770[a]	0.0335	0.8943	6	0.0283	6.7754[a]	0.0335	6.8875[a]	6	0.0012	2.5105[a]	0.0028	2.7993[a]

Notes: See Table 5

and the EURO rate has changed dramatically. More specifically, the results for Period 3 show unidirectional directional causality from the *EURO* rate to the *CD* rate, a finding that contradicts the result for Period 1 that revealed unidirectional causality from the *CD* rate to the *EURO* rate instead.

The sub-period results of the nonlinear causality tests reported in Table 8 are generally in line with the linear results in Table 7. For example, Table 8 shows that during Period 1 (1973-1985) nonlinear causality ran mostly uni-directionally from the *CD* rate to the *EURO* for all three maturities, a finding that mirrors the results in Table 7 for Period 1. The results for Period 2 (1986-1995) reveal strong bi-directional nonlinear causality between the *EURO* and the *CD* deposit rates for all three maturities. As seen in Table 8, all the NORM statistics are highly significant during Period 2 for all the three maturities considered. Finally, like Period 2, Period 3 (1996-2006) uncovers bi-directional nonlinear causality between the *CD* rate and the *EURO* rate for all three maturities. However, unlike Period 2, Period 3 shows that the causality effect of the *EURO* rate on the *CD* rate is greater that the (causality) effect of the *CD* rate on the *EURO* rate. More specifically, the NORM statistics testing the hypothesis that the "*EURO* does not cause *CD*" are uniformly greater than the NORM statistics testing the hypothesis that the "*CD* does not cause *EURO*". This may be indicative of the growing importance of the Eurodollar market.

DISCUSSIONS

The full-sample and Period 1 (1973-1985) results of this study using linear tests show a unidirectional causality from the *CD* rate to the *EURO* rate for the three- and six-month maturities. Period 2 (1986-1995) reveals a strong bi-directional causal relation between the *CD* rate and the *EURO* rate for the three- and six-month maturities while Period 3 (1996-2006) shows mostly a unidirectional causal relation from the *EURO* rate to the *CD* rates for the three- and six-month maturities. This later finding implies that the influence of US markets on the Eurodollar market has totally dissipated. One reason for this may be due to the significant shift that occurred in the Eurodollar market in the second half of the 1990s. As noted by McGuire (2004), the rate at which banks channel funds back into the London's inter-bank market declined dramatically following the introduction of the Euro currency and the subsequent contraction in foreign exchange trading. Banks in London continue to receive US dollar deposits from banks abroad but are using increasingly large portions of these surplus dollars to finance non-bank borrowers, primarily in the United States. These non-bank borrowers actively engaged in cross-border activities appear to be securities houses, hedge funds and other non-bank financials that have relied on the banks in London to leverage their capital in taking positions in fixed income securities.

It is interesting to note that the findings for the three-month rates by and large disagree with those of other researchers (Swanson (1987, 1988), and Fung and Isberg (1992)). For example, Hartman (1984) uses weekly data and the Granger-Sims test to explore the causal link between the three-month U.S. *CD* and the Eurodollar deposit rates during the 1971 to 1978 period. During the 1971-1974 sub-period he finds unidirectional causality and during the 1975-1978 sub-period he reports bi-directional causality. Consequently, Hartman concludes the influence of the Eurodollar market on the U.S. money market is growing because of the change in interest-rate causality that he attributes to increasing financial-market integration.

Swanson (1987, 1988) uses daily data on the same deposit rates and the Granger methodology to conduct similar tests during the mid-1973 through 1984 period. Overall, her results show unidirectional causality from 1974 through 1980, and reverse unidirectional causality for 1982 and 1983. During 1981, she finds bi-directional causality. Swanson's results contradict Hartman's (1984) finding of bi-directional causality during the 1975-1978 period. Swanson concludes, as did Hartman, that her results indicate the influence of the Eurodollar market on the U.S. money market is growing because of increasing financial-market integration.

Fung and Isberg (1992) use daily data and an error correction model (ECM) to test for causality during the 1981 to 1988 period. During the 1981-1983 sub-period they find unidirectional causality and during the 1984-1988 sub-period they find bi-directional causality. Fung and Isberg's results contradict Swanson's (1987, 1988) findings of bi-directional causality during 1981 and reverse unidirectional causality during 1982 and 1983. Fung and Isberg conclude, as did Hartman (1984) and Swanson, that their results indicate the influence of the Eurodollar market on the U.S. money market is growing because of increasing financial-market integration.

As for the one-month yield, there is evidence of significant bi-directional causal relation between the *CD* and the *EURO* rates over the entire sample period (Panel A, Table 3) while sub-period analysis results (Table 7) are similar to those for the three- and six-month rates.

The results of this study suggest that market integration, as gauged by the relation between Eurodollar yields and U.S. yields, may be period-driven. As shown in this study, there is an irrefutable relation between the *EURO* and U.S. *CD* rates for all three maturities during the second sub-period (1986-1995) when both linear and nonlinear tests are used. By contrast, linear results using all three maturities show only a unidirectional link between the two rates during the first (1973-1985) and third (1996-2006) sub-periods and thus cast doubt on the notion that these two markets have become increasingly integrated. The linear results, therefore, seem to cast doubt on the notion that the Eurodollar and U.S. *CD* markets have become more integrated.

The findings of this study also seem to suggest that successful independent national monetary policies may be period specific. As shown in Table 7, the conduct of an autonomous U.S. monetary policy would have been impossible during the second and third sub-periods (1986-1995 and 1996-2006, respectively) because the two markets were highly intertwined during the first sub-period, while the third sub-period was characterized by unidirectional causality from the *EURO* rate to the *CD* rate. By contrast, the lack of a causal relation from the *EURO* rate to the *CD* rate during the first sub-period (1973-1985) seems to suggest that independent U.S. monetary policies had a greater chance of success during this sub-period.

The linear results of this study are robust since they are mostly also supported by nonlinear tests. The only noticeable divergence between linear and nonlinear results occurs in Period 3. As already noted, linear tests for Period 3 show only a unidirectional causality from the *EURO* rate to the *CD* rate (Table 7) while nonlinear tests show a bi-directional relation between the two rates for all three maturities even though the impact of the *EURO* rate on the *CD* rate is stronger than the other way around.

Finally, the results of the causality tests reported in this study are consistent with the presence of market imperfections and/or transaction costs. The identification of such imperfections and

costs and an explanation for their existence would seem desirable in light of the relatively newly established Eurodollar futures market.

CONCLUSIONS

This paper employs linear and nonlinear Granger causality tests to re-examine the dynamic relation between daily Eurodollar and U.S. certificates of deposit rates during the July 16, 1973 to May 1, 2006 period. This study also conducts sub-period analysis based on the switching regression technique of Goldfeld and Quandt (GQSRT) (1972,1973, 1976). The main empirical findings are summarized as follows:

1. Full-sample results show a significant bi-directional linear and nonlinear Granger causality between the *EURO* and *CD* interest rates for the one-month maturity. As for the three- and six-month rates, linear tests reveal a significant unidirectional causality from the *CD* rate to the *EURO* rate while nonlinear causality tests show that the relation between these two rates is significantly bi-directional.
2. Sub-sample results based on linear tests show significant unidirectional causal relation from the *CD* rate to the *EURO* rate during the first sub-period (1973-1985), bi-directional causality between the two rates for all three maturities during the second sub-period (1986-1995), and unidirectional causality from the *EURO* rate to the *CD* rate for all three maturities.
3. Finally, nonlinear causality test results for the sub-periods are similar to those from the linear tests excepting the third period. Linear results show unidirectional causality from the *EURO* rate to the *CD* rate while nonlinear results reveal bi-directional causality between the two rates for all three maturities.

An interesting line of future inquiry would be an attempt to investigate whether the results of this study are confined to the short end of the maturity spectrum or whether they carry over to the long end of the maturity spectrum. Another valuable extension of this study would include other currency denominations such as the Euro Yen, the Euro Canadian Dollar or the Euro Swiss Franc.

ACKNOWLEDGEMENT

The authors thank Prashob Viruthiamparambath for able research assistance and Deborah R. Kaplan, Catherine Kirchmeyer, Jeff Madura, and Ike Mathur for constructive comments.

NOTE

1. Following the recommendation of an anonymous referee, we also ran all of our tests using the periods from 1973 to 1985, 1986 to 1996, and 1997 to 2006. The results, not given here but available upon request, are qualitatively similar to those reported in the paper.

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