LONG MEMORY IN INTEREST RATE FUTURES MARKETS: A FRACTIONAL COINTEGRATION ANALYSIS

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INTRODUCTION

Knowledge of the causal relationship between interest rate changes in the U.S. (domestic) and Eurodollar (external) markets is of major importance to the understanding of international financial market integration. Research analyzing this relationship includes, among others: Hendershott (1967), Kaen et al. (1983), and Fung and Isberg (1992). In the interest rate futures markets, spread trading between the U.S. Treasury bill and Eurodollar futures has become increasingly popular in recent years. Two recent articles by Fung and Lo (1995) and Tse and Booth (1995) report that these two interest rate futures are cointegrated, indicating a long-run relationship between them. This study, employing fractional cointegration techniques, reexamines this relationship by exploring the long memory between the Treasury bill and Eurodollar futures markets.

Fung and Lo (1993) use the rescaled range analysis of Mandelbrot (1972) and Lo (1991) to examine the long-run memory in Eurodollar and Treasury bill futures markets. They test each futures series individ-

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ually and find that neither futures possesses long memory. That result suggests that past movements of a particular implied futures rate are of no use in predicting future changes for that same rate. However, applying a fractional cointegration analysis, the current article provides evidence of the existence of a long-memory relationship between these two interest rate futures markets. Specifically, the deviations from the cointegrating relationship are shown to have long memory and can be well described as a fractionally integrated process. These long-memory relationship results provide a better understanding of the relationship between the two markets.

The cointegration tests used by Fung and Lo (1995) and Tse and Booth (1995) restrict the error correction terms or equilibrium errors to be an I(0) process, which is not persistent. Accordingly, the results of those studies imply that the equilibrium errors respond very quickly to innovations or shocks to the system so that deviations from equilibrium are not persistent. However, the current article shows that the equilibrium errors behave as a fractionally integrated series; or, more explicitly, I(0.75), which displays fairly slow mean reversion. That is, the equilibrium errors respond very slowly to innovations. Thus, adjustment to equilibrium between markets will take a longer time than what the results of Fung and Lo (1995) and Tse and Booth (1995) suggest.

FRACTIONAL INTEGRATION AND COINTEGRATION

Fractional Integration and Geweke–Porter–Hudak Tests

The Autoregressive Fractionally-Integrated Moving Average (ARFIMA) representation introduced by Granger and Joyeux (1980) and Hosking (1981) is a natural extension of ARIMA models:

$$\Phi(L)(1 - L)^d y_t = \Theta(L)\varepsilon_t \tag{1}$$

where L is the lag operator; ε_t is white noise; and the order, d, of the differencing operator can take on noninteger values. If $0 < d < 1$, an innovation has no permanent effect on the value of y_t, but its mean reversion properties are persistent (i.e., the dependence between distant observations is noticeable).

Geweke and Porter-Hudak (1983) (GPH) propose a semi-nonparametric method for estimating d using the following OLS regression based on the periodogram.

$$\ln\{P(\omega_j)\} = \beta_o + \beta_1 \ln\{4 \sin^2(\omega_j/2)\} + \eta_j \tag{2}$$

with $\beta_1 = -d$, where $P(\omega_j)$ is the periodogram of $\{y_t\}$ at frequency, ω_j, $\omega_j = 2\pi j/T$ $(j = 1,\ldots,T-1)$. As recommended by GPH, the number of low frequency periodogram ordinates, n, used in eq. (2) is equal to $n = T^v$ with $v = 0.50$ or 0.55. Also, to increase the efficiency in estimating β_1, the current study uses the theoretical variance of the regression error, η_j, $\pi^2/6$, proved by GPH, in eq. (2). The significance of fractional integration in the series is indicated by the standard t-statistic, β_1. The virtue of this frequency domain regression method is that it permits estimation of d without knowledge of p and q in ARFIMA (p, d, q). Moreover, Cheung (1993b), using a Monte Carlo approach, documents that the GPH test is robust to short-term dependence, conditional heteroskedastic effects, and variance shifts.[1] The GPH test has been employed by, e.g., Diebold and Rudebusch (1989), Cheung (1993a), Cheung and Lai (1993), Fang, et al. (1994), and Shea (1991).

Cointegration and Fractional Cointegration

Cointegration methodology, fully developed in Granger (1986) and Engle and Granger (1987), has been widely used to describe the existence of a stationary relationship among two or more nonstationary variables. Consider N series x_{it}, $i = 1,\ldots,N$ and let X_t be a $N \times 1$ vector of x_{it}. In the classical paradigm for cointegration, e.g., the Johansen and ADF tests, all the members of the X_t vector are assumed to be I(d) processes with $d = 1$ (or an integer), and the cointegrating linear relationship $\alpha' X_t$ with α as the cointegrating vector is presumed to be I($d - b$) with $b = 1$. The Granger Representation Theorem, nevertheless, only requires that the linear combination, $\alpha' X_t$, be stationary, and, therefore, the discrete options I(1) and I(0) are rather restrictive. Particularly, the equilibrium error, z_t, will be mean-reverting as long as it follows a fractionally integrated process, I(d), with $0 < d < 1$, as discussed in the previous subsection. In this case, the error correction term responds more slowly to shocks so that deviations from equilibrium are more persistent. Since the error correction term is depicted to possess long memory, the fractional cointegration relationship analyzed

[1]The robustness to conditional heteroskedastic effects and variance shifts is crucial, since, as pointed out by Milonas et al. (1985), results of futures price dynamics may be biased by variance nonstationarity.

herein may give a better understanding of the relationship between the Treasury bill and Eurodollar futures.

Similar to Engle and Granger's (1987) two-step cointegration procedure, z_t is obtained from the cointegrating regression,

$$x_{it} = \alpha_o + \alpha_1 x_{jt} + z_t \tag{3}$$

where $i = 1$ and $j = 2$, or vice versa. d of z_t is then estimated by eq. (2). As pointed out by Baillie and Bollerslev (1994, p. 743), "[i]t follows from Yajima (1988) and Cheung and Lai (1993) that if $[X_t]$ is I(1) but $d > 0$, so that $[X_t]$ is fractionally cointegrated, the OLS estimator of α is consistent and converges at the rate of T^{1-d}". Cheung and Lai (1993) and Baillie and Bollerslev (1994) provide a detailed description of fractional cointegration, and the empirical approaches used in the current study are analogous to Cheung and Lai (1993).

EMPIRICAL RESULTS

Daily closing index prices of the three-month Treasury bill and Eurodollar futures from March 1, 1982 to February 22, 1994 (3032 observations) are from Commodity Systems, Inc. (CSI). The nearest contract is used until the first trading day of the expiration month. Since both the Treasury bill and Eurodollar futures are traded at the International Monetary Market (IMM), the problem of nonsynchronous trading does not appear. Hereafter, the ticker symbols TB and ED are used to represent the three-month Treasury bill and Eurodollar futures, respectively.

Unlike Treasury bills, which are sold on discount, Eurodollar time deposits pay add-on interest. To compare them on an equal basis, the implied discount yield of TB futures and the implied add-on yield of ED futures are converted into the bond equivalent yield, along with logarithmic transformation.[2] Nonetheless, results reported throughout this study are qualitatively the same when the originally implied yields without any transformations are used.

Unit Root Tests

Fung and Lo (1993) evince that the TB and ED futures have a unit root while their first difference is stationary by using the ADF test.

[2]The following formulae convert (a) a discount yield and (b) an add-on yield to a bond equivalent yield respectively: Bond equivalent yield = (a) 365 × yield/[365 − yield × (days to maturity)]; (b) 365/360 × yield, where yield = (100 − index price)/100.

However, Diebold and Rudebusch (1991) and Sowell (1990) show that the ADF test has low power against fractional integration alternatives. This study uses the GPH test to estimate the differencing order $\tilde{d}$ in the first difference of the relevant series, $(1 - L)^{\tilde{d}} w_{it}$ where $w_{it} = \Delta x_{it}$ and x_{it} is TB_t or ED_t. As d of the level series equals $1 + \tilde{d}$, a value of $\tilde{d}$ (d) not significantly different from zero (one) corresponds to a unit root in x_{it}.

The values of $d = 1 + \tilde{d}$ estimated by the periodogram OLS regression (2) on w_{it} and presented in Table I are not significantly different from one. The hypothesis of a unit root is not rejected for either level series.

Fractional Cointegration

The equilibrium error, z_t, is given by the cointegrating regression (3). MacKinnon (1991, pp. 269–270) argues that including a trend time may make the resulting test statistic invariant to the value of the drift term in the data-generating process and "[a]dding a trend to the cointegrating regression often makes sense for a number of other reasons, as Engle and Yoo (1987) discuss."[3] Therefore, both of the results with and without a trend in eq. (3) are discussed. In Table II, the cointegrating coefficient, α_1, in the cointegrating regression is found to be very close to one for either model with or without trend. This one-to-one linear combination suggests that the interest rate differential between ED and TB, namely the TED spread, plays an integral role in the cointegration between markets. [See Tse and Booth (1995) for further information.]

TABLE I
Results of Unit Roots Using GPH Tests

TB			*ED*		
v	d	t-*Stat.*[a]	v	d	t-*Stat.*[a]
0.50	1.119	1.223	0.50	1.025	0.259
0.55	1.132	1.707	0.55	1.075	0.962

$d = \tilde{d} + 1$; $\tilde{d}$ is estimated by the GPH test using ΔTB or ΔED as the series examined in eq. (2) in the text.
[a]The unit root null hypothesis of $d = 1$ is tested against the alternative $d \neq 1$.

[3]In essence, adding a trend term in the cointegration regression is a statistical issue, which may not provide any economic intuition. However, Booth and Chowdhury (1991, footnote 7) remark that the trend term accounts for the possibility that the time-varying risk premia of markets might be different.

TABLE II
Results of OLS Regressions between TB and ED

	Dependent Variable (x_i)							
	TB				*ED*			
	α_o	α_1	γ (10^{-4})	R^2	α_o	α_1	γ (10^{-4})	R^2
No trend	−0.055	0.965	N/A	0.991	0.075	1.026	N/A	0.991
With trend	−0.009	0.977	0.66	0.991	0.162	0.994	−0.017	0.991

The parameters are obtained for the following cointegration regression [eq. (3) in the text]: $x_{it} = \alpha_o + \alpha_1 x_{jt} + \gamma t + z_t$.

For expository reasons, the first 120 autocorrelation coefficients ρ's of z_t, as well as the series in level and in first difference, are displayed in Table III. The extremely slow decline, if any, of the autocorrelation coefficients of both level series [with $\rho(1) = 0.99$, $\rho(60) = 0.89$, and $\rho(120) = 0.78$] and the small coefficients of any lags of the first differenced series suggest that TB and ED follow a I(1) process. In contrast, the autocorrelation coefficients of z_t at lags up to 5 are still greater than 0.9, but the coefficients afterward slowly decline to a value of 0.58 at lag 60 and 0.34 at lag 120. Of particular interest, using traditional cointegration tests that do not allow noninteger values of the differencing order d of z_t, Fung and Lo (1995) and Tse and Booth (1995) document that z_t is an I(0) series, which responds to shocks very rapidly. However, the slowly declining behavior of the autocorrelation

TABLE III
Autocorrelation Coefficients, $\rho(\tau)$

τ	*TB*	Δ*TB*	*ED*	Δ*ED*	z^a (*No Trend*)	z^a (*With Trend*)
1	0.998	0.075	0.998	0.073	0.984	0.984
2	0.996	0.032	0.997	0.033	0.967	0.968
3	0.994	−0.001	0.995	0.001	0.953	0.954
4	0.992	−0.014	0.993	0.010	0.938	0.939
5	0.991	−0.027	0.992	−0.021	0.924	0.926
6	0.989	0.013	0.990	0.007	0.911	0.913
12	0.977	0.002	0.979	0.011	0.847	0.850
24	0.954	0.026	0.956	0.023	0.756	0.757
60	0.892	−0.006	0.894	0.001	0.583	0.585
120	0.777	0.010	0.782	−0.006	0.346	0.359

[a] z's are the error correction terms derived from the cointegration regression with and without trend.

coefficients suggests that z_t responds to shocks more slowly than an I(0) series does.

Since the cointegrating vector, α, is estimated from the OLS regression (3), z_t appears more stationary than if it were derived at the true α. [See also Engle and Granger (1987) and Cheung and Lai (1993).] Hence, the critical values for the GPH tests for testing fractional cointegration between TB and ED are obtained from the Monte Carlo experiment in 10,000 replications with the same sample sizes, 3032. The results are presented in Table IV, which shows that the critical values generated for testing the null hypothesis of no (fractional) cointegration of $d = 1$ against the one-sided alternative of $d < 1$ are greater in magnitude than the standard t-statistics.[4]

The differencing order, d of z_t, according to eq. (2) is estimated to be significantly less than one (about 0.75) and larger than zero, implying that z_t is mean-reverting. As shown in Panel A of Table V, this result is robust to the normalization of eq. (3), the choice of v, and the inclusion of a trend time. Nevertheless, the t-statistics testing $d < 1$ with $v = 0.50$ are not significant at the 1% level (though they are significant at the 5% level); hence, evidence showing long memory between TB and ED is not conclusive. However, for such a large sample

TABLE IV

The Critical Values of the GPH Test for Cointegration

	T = 1500				T = 3032			
	No Trend		*With Trend*		*No Trend*		*With Trend*	
	v = 0.50	v = 0.55	v = 0.50	v = 0.55	v = 0.50	v = 0.55	v = 0.50	v = 0.55
Percentile								
1%	−2.999	−2.941	−2.906	−2.988	−2.748	−2.900	−2.887	−2.822
5%	−2.092	−2.036	−2.053	−2.074	−1.932	−2.043	−2.013	−1.989
10%	−1.663	−1.624	−1.617	−1.626	−1.538	−1.613	−1.588	−1.606
50%	−0.244	−0.222	−0.215	−0.196	−0.208	−0.273	−0.225	−0.227
90%	1.349	1.624	1.381	1.039	1.014	1.042	1.034	1.045
95%	1.018	1.039	1.049	1.380	1.354	1.382	1.368	1.372
99%	1.974	2.052	2.008	2.004	2.009	1.999	1.931	2.049
Mean	−0.288	−0.258	−0.258	−0.250	−0.244	−0.273	−0.270	−0.260
Skewness	−0.313	−0.296	−0.272	−0.299	−0.204	−0.229	−0.262	−0.149
Kurtosis	0.316	0.444	0.294	0.296	0.209	0.208	0.268	0.207

The critical values are obtained by simulating two independent random walk processes in 10,000 replications.

[4]This may be simply inferred from the negative skewness of the distribution.

TABLE V

Results of Fractional Cointegration Using GPH Tests

	Dependent Variable							
	TB				*ED*			
			t-*Statistics*				t-*Statistics*	
	v	d	d = 1[a]	d = 0[b]	v	d	d = 1[a]	d = 0[b]
(A) 3/1/1982–2/22/1994								
No trend	0.50	0.747	−2.600[c]	7.677[d]	0.50	0.751	−2.554[c]	7.703[d]
	0.55	0.769	−2.978[d]	9.914[d]	0.55	0.767	−3.003[d]	9.885[d]
With trend	0.50	0.752	−2.539[c]	7.699[d]	0.50	0.754	−2.520[c]	7.724[d]
	0.55	0.767	−3.008[d]	9.902[d]	0.55	0.763	−3.057[d]	9.842[d]
(B) 3/1/1982–9/30/1987 (Pre-crash)								
No trend	0.50	1.011	0.090	8.272[d]	0.50	0.993	−0.052	7.377[d]
	0.55	0.880	−1.211	8.881[d]	0.55	0.893	−1.086	8.906[d]
With trend	0.50	0.977	−0.182	7.731[d]	0.50	1.009	0.077	8.193[d]
	0.55	0.949	−0.516	9.602[d]	0.55	0.881	−1.203	8.906[d]
(C) 1/4/1988–2/22/1994 (Post-crash)								
No trend	0.50	0.652	−3.948[d]	7.397[d]	0.50	0.652	−2.912[c]	5.456[d]
	0.55	0.644	−3.686[d]	6.668[d]	0.55	0.701	−3.097[d]	7.261[d]
With trend	0.50	0.532	−3.919[d]	4.455[d]	0.50	0.657	−2.867[c]	5.492[d]
	0.55	0.647	−3.655[d]	6.699[d]	0.55	0.703	−3.072[d]	7.271[d]

$d = \tilde{d} + 1$; $\tilde{d}$ is estimated by the GPH test using Δz_t as the series examined in eq. (2) in the text. Results are virtually the same if d is directly estimated from the level z_t. Critical values are based on the simulated values reported in Table IV.
[a]The null hypothesis of $d = 1$ is tested against the alternative $d < 1$.
[b]The null hypothesis of $d = 0$ is tested against the alternative $d > 0$.
[c]Significant at the 5% level.
[d]Significant at the 1% level.

size, the 1% significance level is more appropriate than the 5% level. See Connolly (1989) for a detailed explanation on this issue.[5]

Several studies find that the October 1987 stock crash (the crash) changed the structure of international movements among financial markets. [See, e.g., Malliaris and Urrutia (1992), and Arshanapalli and Doukas (1993).] In addition, both TB and ED experienced the greatest (absolute) yield changes on October 19, 1987 for the period examined. The whole period, therefore, is divided into two subperiods, the pre- (March 1, 1983 to September 30, 1987) and post-crash (January 2, 1988 to February 22, 1994) periods, for the analysis of long memory between TB and ED. Panels B and C of Table V indicate that the long

[5]As Connolly (1989) contends, the significance level should be adjusted downward with sample size. Otherwise, spurious significant results are incurred by large sample size distortion.

memory evidence is much stronger during the post-crash period. In particular, d in the post-crash period is about 0.66, on average, and significantly less than one in each case, while that in the pre-crash period is 0.95 and not significantly different from one even at the 10% level. Note that the critical values used for the subperiods are obtained from generating two random walk processes with a sample size of 1500 (see Table IV). In short, TB and ED are fractionally cointegrated, but this long-memory relationship is only obtained in the post-crash period.

Derivations of Long Memory

Since TB and ED are found to be fractionally cointegrated in the post-crash period with the TED spread (ED minus TB) being the cointegrating vector, one may interpret that yields adjust because the TED spread is out of equilibrium.[6] Alternatively, fractional cointegration may imply that the TED spread provides agents with information for forecasting changes in TB and ED yields.[7] Indeed, trading the interest rates differential can be a means to speculate on general economic conditions and on the soundness of banks in particular. More specifically, the TED spread may be considered a quality spread that reflects the risk premium of holding a Eurodollar deposit versus a Treasury bill. For example, the spread increases significantly on the date of the crash.

Regardless of what reasons are given for the use of (fractional) cointegration, Granger and Escribano (1986) state that there must be Granger causality in at least one direction through the knowledge of lagged z_t in an error correction model. But Granger and Newbold (1986, p. 226) argue that if the series are cointegrated, a further possible explanatory variable to account for changes in the long run is suggested. Risk premium may be one such variable to account for changes in the long run is suggested. Risk premium may be one such variable, and this is consistent with market efficiency. Dwyer and Wallace (1992) further point out that whether series are cointegrated is a function of the relevant model and cointegration can be consistent with market efficiency in various contexts.

[6] As suggested by Engle and Granger (1987), cointegration reflects the behavior of economic forces interacting to obtain an equilibrium. Booth and Chowdhury (1991) extend this explanation by discussing the long-run dynamics in terms of equilibrium overshooting [in the Dornbusch et al. (1983) sense] between two similar assets.

[7] This argument is illustrated in Campbell and Shiller's (1987, 1988) present value models in the context of term structure.

Moreover, an out-of-sample forecasting investigation may be required to confirm the fractional cointegration result for the post-crash period, because the fractionally integrated error correction term should furnish forecasting improvement in future yield changes. However, as argued by Baillie and Bollerslev (1994), for fractional cointegration, any improvement in forecasting may only be evident several years later due to the slow adjustment to equilibrium. As a result, this forecasting experiment is reserved for future research when a much longer span of data is available.

CONCLUSIONS

Long memory between the Treasury and Eurodollar futures is examined by using fractional cointegration techniques. The Geweke–Porter–Hudak tests show that each individual series follows a I(1) process and does not have long memory. These results are consistent with Fung and Lo (1993). However, the GPH tests provide evidence that the deviations from the cointegrating relationship with the TED spread as the cointegrating vector possess long memory. The results suggest that TB and ED may be tied together through a fractionally integrated I(d) type process such that the equilibrium errors between them exhibit slow mean reversion. Accordingly, adjustment to equilibrium between markets will take a longer time than what the results of Fung and Lo (1995) and Tse and Booth (1995) suggest. Investors should incorporate this slow adjustment process into their spread trading strategies.

A more detailed examination reveals that there is a structural change in the relationship: the two markets are fractionally cointegrated during the post-crash period, but not during the pre-crash period. These results are comparable to previous studies which find that the linkage among international financial markets has increased after the crash. In conclusion, although this article supports the long memory relationship between the Treasury and Eurodollar futures after the crash, further research is needed to examine the sources of the long memory relationship between them.

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