A Cointegration Test for Oil Market Efficiency

Article *in* Journal of Futures Markets · December 1993

DOI: 10.1002/fut.3990130810

2 authors, including:

William Crowder
University of Texas at Arlington

A Cointegration Test for Oil Futures Market Efficiency

William J. Crowder
Anas Hamed

INTRODUCTION

The efficiency of the futures or forward market in commodities or currency exchange has been extensively tested using the model

$$S_t = \beta_0 + \beta_1 F_{t-1} + \varepsilon_t \tag{1}$$

where S_t is the natural logarithm of the spot price in period t and F_{t-1} is the analogously defined forward or futures price on a one period contract in period $t - 1$. The joint restriction of market efficiency and risk neutrality implies coefficient values of $\beta_0 = 0$ and $\beta_1 = 1$. These restrictions are based on a definition of market efficiency that argues that price changes from one period to the next should be unpredictable given current information. If the futures price, F_{t-1}, contains all relevant information to forecast the next period's spot price, S_t, as this definition of market efficiency implies, then F_{t-1} should be an unbiased predictor of the future spot price. This represents Fama's (1970) notion of weak form efficiency.

Several studies have examined this efficient market definition for commodities and foreign exchange futures markets [e.g., Baillie and Myers (1991); Chowdhury (1991); Kroner and Sultan (1991)] with mixed results. In each study, the non-stationarity of the underlying univariate data generating processes (DGP) required use of cointegration techniques to obtain valid inference on the relationship between S_t and F_{t-1}. [See Engle and Granger (1987); Hakkio and Rush (1989); Chowdhury (1991); and Lai and Lai (1991) for discussion.] A common explanation for the rejection of the simple efficiency hypothesis has been the existence of a risk premium. Such a risk premium can account for the existence of non-zero speculative returns in the futures market. This does not imply markets are inefficient, only that investors require compensation for the risk they undertake.

The authors would like to thank Hank Bessembinder, Dan Himarios, Dennis Hoffman, Mike Melvin, and two anonymous referees for the helpful comments. Any remaining errors must be attributed, solely, to the authors.

William J. Crowder is an Assistant Professor of Economics at the University of Texas at Arlington.

Anas Hamed is a Lecturer in the Department of Economics at Arizona State University.

The Journal of Futures Markets, Vol. 13, No. 8, 933–941 (1993)

CCC 0270-7314/93/080933–09

In an effort to reconcile the empirical findings with the notion of efficient markets, Brenner and Kroner (1992) derive the conditions for market efficiency under the no arbitrage profit rule.[1] Consider an investor in the oil futures market who should be indifferent between taking an open spot position (buying) and agreeing to sell at the futures price (shorting a futures contract), or purchasing a risk-free bond that matches the maturity of the futures contract, i.e., the expected return of a hedged position in any commodity market should equal the risk-free return. This no risk-free profit condition generalizes readily to short futures positions. This arbitrage condition implies the following relationship between spot price, futures price, and the return on the risk-free bond;

$$S_t = \beta_0 + \beta_1 F_{t-1} + \beta_2 R_t + \varepsilon_t \tag{2}$$

where S_t and F_{t-1} are defined as in eq. (1) and R_t is the continuously compounded rate of return on the risk-free bond, i.e., $R_t = \ln(1 + r_t)$.[2] The constant term need not be zero, but the error term should be serially uncorrelated for market efficiency. That is, profitable arbitrage deviations should not be predictable based on past information. A necessary condition for futures market efficiency, assuming all variables are non-stationary, is that the system be cointegrated with cointegration vector $[1, -1, -1]'$, i.e., $\beta_1 = 1$ and $\beta_2 = 1$. Dwyer and Wallace (1992) maintain that this no arbitrage profit definition for market efficiency has greater economic meaning than the simple efficiency definition discussed above, especially in the presence of an unobservable risk premium. They advocate that the definition of market efficiency should rely on the condition that risk-free arbitrage profits should not persist. This definition allows for some predictability of asset price changes (for example, through a possibly time varying risk premium) but precludes markets from exhibiting persistent risk-free arbitrage profit opportunities. A necessary condition for market efficiency is that the variables comprising the arbitrage condition form a stationary equilibrium.[3]

This study analyzes the cointegration properties of the oil futures market to allow valid inference on market efficiency. Unlike most analyses of commodities futures markets, the results presented imply that monthly spot and futures prices are cointegrated with cointegration vector insignificantly different from $[1, -1]'$, supporting the simple efficiency hypothesis. The evidence, surprisingly, weighs in against the arbitrage relationship given in eq. (2). Several possible explanations for this result are offered. It may be that the unobserved convenience yield is best characterized with a stochastic trend in its univariate representation and is cointegrated with the risk-free rate. This would be completely consistent with the results from both the simple efficiency and arbitrage models. Another possibility is that the oil futures market might be characterized by a preferred habitat motive. Finally, it may be that this market is inefficient with risk-free profit opportunities occurring with probability one, clearly a violation of efficient markets regardless of the definition used. The second section discusses econometric issues. The third section presents the empirical results and is followed by the conclusion.

[1] Brenner and Kroner (1992) demonstrate that under quite general assumptions about the underlying forcing variables, the arbitrage condition for futures markets, with mark-to-market pricing, and forward markets, that do not mark-to-market, are analogous. The arbitrage condition itself is a well-known condition for simultaneous equilibrium in the spot, futures, and credit markets, e.g., Garbade (1982). For markets to be efficient, the deviations from this equilibrium must be mean zero, with allowances for transactions costs, and have little or no persistence.

[2] This is eq. (13) from Brenner and Kroner (1992).

[3] A sufficient condition is that no persistence exists in equilibrium deviations.

COINTEGRATION AND EFFICIENCY TESTS

It is a characteristic of asset prices, in general, that their univariate time series representation follow non-stationary processes. Specifically, asset prices appear to be integrated of order one, denoted $X_t \sim I(1)$.[4] The use of ordinary least squares (OLS) in estimating models characterized by integrated variables has received criticism due to the potential spurious results [see Granger and Newbold (1974)]. In general any linear combination of non-stationary variables will itself be non-stationary. A special case arises when the variables are linked by a long-run equilibrium that prevents them from drifting too far apart. The models of futures market efficiency, eqs. (1) and (2), represent two such equilibrium relationships. Such time series are said to be cointegrated and the OLS regression of cointegrated variables is not spurious. Even when the regression is not spurious, the inferences based on OLS standard errors are generally not valid [see Stock (1988)]. Even though OLS produces asymptotically superconsistent estimates of the cointegrating vector, simulation experiments provide evidence of significant small sample bias due to the endogeneity and simultaneity of the variables [see Banerjee et al. (1986)]. To estimate the hypothesized equilibrium implied by market efficiency and obtain valid inference, Johansen's (1989) cointegration technique is used. Johansen has developed the maximum likelihood estimator for a cointegrated system with gaussian errors. Gonzalo (1989) uses Monte Carlo methods to show that this estimator is robust to differing error structures and produces superior inference with respect to other estimators.

The Granger representation theorem implies that any cointegrated system may be written as an error correction model, as in eq. (3),

$$\Delta X_t = \mu + \Pi X_{t-1} + \sum_{j=1}^{k-1} \Gamma_j \Delta X_{t-j} + \epsilon_t \qquad (3)$$

where X_t is a $p \times 1$ vector of $I(1)$ variables and Π is $p \times p$ matrix that has reduced rank when the variables in X_t are cointegrated. The matrix Π can be decomposed into two $p \times r$ matrices α and β such that $\alpha\beta' = \Pi$. The columns of β represent the r linear combinations of X_t that are stationary or cointegrated. The corresponding columns of α represent the corresponding error correction coefficients that can be loosely interpreted as speed of adjustment parameters. Johansen develops test statistics to examine the null of r cointegration vectors. The reader is referred to Johansen (1989) for a complete discussion of the estimation technique.

The term μ contains all deterministic components of the X_t system such as a constant or deterministic time trend. If μ contains a time trend then X_t has a quadratic trend in its DGP. If μ contains a constant only, then X_t has a linear deterministic trend. If $\mu = 0$ it is still possible to include a constant in the cointegration vector by adding a constant to the vector X_t. Johansen (1991) describes a sequential testing procedure to determine the properties of the term μ. This is important for two reasons. First, the asymptotic distributions of the cointegration test statistics are dependent upon the presence of trends and/or constants in μ. Second, the speculative efficiency hypothesis requires the constant term to be zero. This represents a testable hypothesis on the parameter μ.[5]

[4]A time series is said to be integrated of order d ($X_t \sim I(d)$) if it is necessary to difference X_t d times to make it stationary. A time series that is non-stationary is necessarily integrated, but an integrated series need not be non-stationary. The condition for non-stationarity is that $d \geq 0.5$. In this article the terms integrated and non-stationary are used interchangably.

[5]Lai and Lai (1991) impose the restriction of no trends without testing for its validity in an application to the foreign currency forward market.

If $X_t = [S_t, F_{t-1}, R_t]'$, then the deviations from the arbitrage condition in eq. (2) should be a mean zero stationary process. This implies that X_t is a cointegrated system if $X_t \sim I(1)$, with cointegrating vector $[1, -1, -1]'$. If the speculative efficiency hypothesis holds, then $X_t = [S_t, F_{t-1}]'$ with implied cointegration vector of $[1, -1]'$ and the risk-free rate should be a stationary process.[6] These represent testable restrictions within the Johansen cointegration framework. By expressing the restrictions as $\beta = H\varphi$, where H is a known $p \times s$ matrix, a likelihood ratio test can be formed analogous to that of eq. (4) below, where $\hat{\lambda}_i^*$ are the restricted eigenvalues from the Johansen estimation methodology. This statistic is asymptotically distributed as $\chi^2\ (r(p - s))$.

EMPIRICAL ANALYSIS

The data used to test the efficiency of the oil futures market are monthly observations from the New York Mercantile Exchange (NYMEX) for the period from March, 1983, to September, 1990, a total of 90 observations. The futures price is defined as the closing price of a futures contract 30 days prior to the last day of trading on that contract. The future spot price is the cash price on the last trading day of the corresponding futures contract. The next futures price observation then begins on the next trading day using the same data matching method for all spot and futures price combinations. The risk-free rate series is defined as the interest rate on the 3-month U.S. Treasury Bill that matures nearest to the last trading day of the relevant futures contract. Therefore, the holding periods for the T-Bill and the futures contract are matched as closely as possible. All variables are transformed into natural logarithms.[7]

The first step in the analysis is to determine the order of integration of the relevant variables.[8] Table I provides augmented Dickey–Fuller (ADF) tests of the null that each series has a unit root. See Dickey and Fuller (1979) and Said and Dickey (1984).

The results from the unit root tests provide evidence of non-stationarity for each of the three series examined. This result is quite robust to lag length specification and inclusion

Table I
UNIT ROOT TESTS

Series	**Spot**		**Futures**		**Risk-Free Rate**	
ADF lag	**w/o trend**	**with trend**	**w/o trend**	**with trend**	**w/o trend**	**with trend**
0	−1.5399	−0.7185	−1.7715	−1.2230	−2.1676	−2.0921
1	−1.8659	−1.2676	−1.7859	−1.2934	−1.6120	−1.4704
2	−1.9937	−1.3375	−1.9415	−1.4244	−1.4742	−1.2634
3	−1.9504	−1.2760	−2.1365	−1.6908	−1.4383	−1.1570
4	−1.8285	−1.0689	−1.9328	−1.3324	−1.4136	−1.0967
5	−1.7199	−0.8928	−1.7581	−1.0140	−1.4333	−1.0967
6	−1.5993	−0.6917	−1.6424	−0.8023	−1.5622	−1.2699

Five percent critical values of −2.89 and −3.46 are taken from Dickey and Fuller (1979) for the w/o trend and with trend specifications, respectively.

[6] Or the risk-free rate is cointegrated with some omitted variable, such as the unobserved convenience yield.

[7] By using monthly observations on 30-day futures contracts, the overlapping forecast error problem described by Hanson and Hodrick (1981), is avoided.

[8] The Johansen estimation procedure does not require pre-testing of the variables for unit roots since the estimator can differentiate among variables of differing orders of integration.

of a time trend in the regression. Therefore, it can be concluded that the series contain a unit root and the use of standard OLS procedures should be avoided for the reasons outlined earlier.[9] The next step is to proceed to the Johansen estimation procedure as discussed above.

The first specification analyzed is the speculative efficiency model. As stated earlier, this version of the efficient markets hypothesis has received the greatest treatment in the literature but has little empirical support. To test the restrictions implied by the simple efficiency hypothesis, it is first necessary to determine if the system can be written in a form that excludes deterministic trends. The procedure described by Johansen (1991) is employed, in which the system is estimated with a deterministic trend,[10] denoted as $H_1(r)$ where r is the maintained number of cointegrating vectors and without a trend but with a constant in the cointegrating vector, denoted $H_1(r)^*$. The test proposed by Johansen is a likelihood ratio test based on the $p - r$ smallest eigenvalues of the estimated Π matrix.

The two specifications, $H_1(r)$ and $H_1(r)^*$, are estimated using lag length of $k = 1$ in eq. (3), which is chosen based on the Akiake information criterion (AIC). The eigenvalue associated with the $H_1(1)$ is 0.033897 while that associated with $H_1(1)^*$ is 0.034707. The test statistic proposed by Johansen is given as,

$$Q = -T \sum_{i=r+1}^{p} \ln\left\{ \frac{(1 - \hat{\lambda}_i^*)}{(1 - \hat{\lambda}_i)} \right\} \tag{4}$$

where $\hat{\lambda}_i^*$ are the eigenvalues associated with the $H_l(1)^*$ specification and the $\hat{\lambda}_i$ are those that are associated with the $H_l(1)$ specification and the test statistic, Q, is asymptotically distributed as a $\chi^2(p - r)$. The calculated test statistic is 0.074651 which is distributed as a $\chi^2(1)$. Therefore, the null that there are no deterministic trends in the data is not rejected at conventional levels of significance. The analysis of the cointegrating relations proceeds under the specification of $X_t = [1, S_t, F_{t-1}]'$.

Table II displays results of the cointegration analysis. The null of zero cointegrating vectors is rejected at high significance levels, but the null of no more than one cointegrating vector cannot be rejected, even at the 50% level of significance.[11] Furthermore, the estimated parameter on the futures price is insignificantly different from unity, the estimate of the constant is insignificantly different from zero, and the residuals are not significantly different from white noise. All of the restrictions implied by the speculative efficiency hypothesis cannot be rejected. Testing the joint hypothesis of $\beta_0 = 0$ and $\beta_1 = 1$, yields an asymptotic $\chi^2(2)$ test statistic of 1.6659 with a p-value of 0.435.[12]

The analysis of the simple efficiency hypothesis leads to several interesting implications. First, the futures price is an unbiased predictor of the corresponding future spot price. There is no evidence of a risk premium and there is no evidence that past forecast errors are useful in predicting the future spot price. Inference on the causal relationship between spot and futures prices in the oil futures market can be made by analyzing the

[9]The first differences of each series were tested for a unit root and all were found to exhibit stationary behavior at high levels of significance. Further analysis of the underlying DGPs, as proposed by Schwert (1987), revealed no significant serial correlation that might bias the tests for unit roots. For a complete discussion of these issues see Schwert (1987, 1989) and Pantula (1991).

[10]This is operationalized by including a constant term in μ.

[11]The test statistic presented is Johansen's Trace test, computed as

$$-T \sum_{j=r+1}^{p} \ln(1 - \hat{\lambda}_j)$$

where $\hat{\lambda}_j$ are the p largest eigenvalue estimates from the Johansen estimation procedure.

[12]These results are robust to sample period and lag length in the vector error correction model.

Table II
COINTEGRATION ANALYSIS

H_0: $r = 0$	H_0: $r \leq 1$	$\hat{\beta}_0$	$\hat{\beta}_1$	$\hat{\alpha}_s$	$\hat{\alpha}_F$	B.L. Q-Stat
193.2355[a]	3.1084	−0.0391	−0.9956	−0.0954	−0.8454	14.3658
		(0.1210)	(0.0735)	(0.1846)	(0.0291)	[0.278]

[a]Denotes significance at the 5% level. r represents the hypothesized number of cointegration vectors in X_t. Critical values for the cointegration tests are taken from Table D.3. of Osterwald–Lenum (1990). $\hat{\beta}_j$ are the cointegrating parameter estimates normalized on S_t. $\hat{\alpha}_j$ are the error correction coefficient estimates implied by the normalized cointegration parameters. Numbers in parentheses are asymptotic standard errors. B.L. Q-Stat is the Box–Ljung statistic for 12th-order serial correlation in the residuals. Number in brackets is the p-value for the Q-Stat.

error correction coefficients in the system X_t. The error correction term on the spot price is statistically insignificant implying that past equilibrium errors do not Granger cause spot price changes. But, the error correction coefficient on the futures price is significant, which implies that past equilibrium errors do Granger cause the futures price. This is very appealing intuitively. It means that past forecast errors only affect the current forecast of the spot price, not the spot price itself which is presumably being driven by fundamentals.

How does the arbitrage equilibrium fit into the above equilibrium relationship? Recall from eq. (2) that the presence of the risk-free rate in the arbitrage equilibrium implies that either the risk-free rate is stationary or that it is non-stationary but cointegrated with spot and futures prices. The results from Table I demonstrate that the null of a non-stationary risk-free rate is not rejected at reasonable levels of confidence. Therefore, for the arbitrage equilibrium to hold, the risk-free rate must be cointegrated with the spot and futures prices. But, it has already been shown that the spot and futures prices are cointegrated without the inclusion of the risk-free rate in the system. Table III presents cointegration results for the arbitrage equilibrium system, i.e., $X_t = [1, S_t, F_{t-1}, R_t]'$.[13]

The null of zero cointegration vectors is easily rejected at high levels of significance but the null of no more than one cointegration vector cannot be rejected at up to the

Table III
ARBITRAGE EQUILIBRIUM COINTEGRATION RESULTS

H_0: $r = 0$	H_0: $r \leq 1$	H_0: $r \leq 2$	$\hat{\beta}_0$	$\hat{\beta}_1$	$\hat{\beta}_2$	B.L. Q-Stat
206.7569[a]	14.8518	2.2237	0.0593	−1.0131	0.5229	13.1347
			(0.1486)	(0.0765)	(0.4449)	[0.359]

[a]Denotes significance at the 5% level. r represents the hypothesized number of cointegration vectors in X_t. Critical values for the cointegration tests are taken from Table D.3. of Osterwald–Lenum (1990). $\hat{\beta}_j$ are the cointegrating parameter estimates normalized on S_t. Numbers in parentheses are asymptotic standard errors. B.L. Q-Stat is the Box–Ljung statistic for 12th-order serial correlation in the residuals. Number in brackets is the p-value for the Q-Stat.

[13]The analysis of trends in the DGPs is conducted in a manner analogous to the simple efficiency model. The two smallest eigenvalues for the unrestricted specification (trended case) are 0.13313 and 0.02456. The corresponding eigenvalues for the restricted case (no trends) are 0.13368 and 0.024953. The computed likelihood ratio statistic is 0.0923525 which is asymptotically distributed as $\chi^2(2)$. Therefore, as in the simple efficiency specification, the null of no trends in the DGPs cannot be rejected.

25% level of significance. The estimated cointegrating parameters normalized on the spot price, i.e., the coefficient on the spot price is set equal to one, reveal that the futures price parameter is insignificantly different from minus one, as predicted by the model, but the risk-free rate parameter is insignificantly different from zero and has the wrong sign.[14] This result implies that the estimated cointegration vector, $\hat{\beta}$, is insignificantly different from $[0, 1, -1, 0]'$, but is significantly different from $[0, 1, -1, -1]'$, the arbitrage model's predicted cointegration vector, where $X_t = [1, S_t, F_{t-1}, R_t]'$. The likelihood ratio test of the restriction $\beta = [0, 1, -1, 0]'$ is 3.0944 which is distributed $\chi^2(3)$ with a p-value of 0.377.

This is truly a puzzling result. It implies that either oil futures traders ignore the availability of a positive risk-free return or that there exists an important omitted variable from the arbitrage relationship. This is true since it is demonstrated that the expected return to futures speculation is zero. There is no evidence of a risk premium. This result seemingly rejects rationality on the part of oil futures market participants. What could possibly explain such a result?

One possible explanation concerns the use of ex post realizations to make ex ante inferences. It is possible that the ex ante return to oil futures speculation is non-zero, but that the ex post return is zero. But, if this is a persistent market outcome, eventually speculators should realize it and incorporate it in their expectations. Another possibility, although unlikely, is that the oil futures market is governed by some preferred habitat motive. Thus, markets for oil futures and risk-free debt may be segmented driving a wedge between their respective returns. Finally, it is possible that the unobserved convenience yield from holding the commodity is non-stationary and cointegrated with the risk-free rate. Gibson and Schwartz (1990) estimate the unobserved convenience yield in the oil commodity market using weekly data from January, 1984, to November, 1988. They demonstrate that the convenience yield has significant mean reversion tendencies and conclude that it is a stationary process. This conclusion is incompatible with cointegration between the risk-free rate and the convenience yield, making this explanation of the failure of the arbitrage model suspect, as well.

CONCLUSIONS

The simple efficiency hypothesis implies that the expected return to futures speculation in the oil futures market should be zero. The arbitrage equilibrium hypothesis implies that the expected return to speculation in the oil futures market should equal the risk-free rate of return. The evidence presented in this article supports the simple efficiency hypothesis but not the arbitrage equilibrium hypothesis. Several explanations are posited for this result. Perhaps semi-strong form market inefficiency prevails since oil futures traders are ignoring risk-free returns above what they expect to earn in the oil futures market. This is, of course, based on ex post observations, not ex ante. If the oil futures market is characterized by a preferred habitat motive on the part of traders, they may ignore possible returns in other markets and trade exclusively in the oil futures market. Or, the unobserved convenience yield may be cointegrated with the risk-free rate. This would contradict the mean reversion properties of the convenience yield cited by Gibson and Schwartz. This suggests the need for an extensive analysis of the time series properties of the unobserved convenience yield to resolve the issues discussed here.

[14]The coefficient estimate on the risk-free rate is significantly different from the hypothesized value of minus one.

Bibliography

Ahn, S., and Reinsel, G. (1988): "Nested Reduced-Rank Autoregressive Models for Multiple Time Series," *Journal of the American Statistical Association,* 83:849–856.

Baillie, R., and Myers, R. (1991): "Bivariate GARCH Estimation of the Optimal Commodity Futures Hedge," *Journal of Applied Econometrics,* 6:109–124.

Banerjee, A., Dolado, J., Hendry, D., and Smith, G. (1986): "Exploring Equilibrium Relationships in Econometrics Through Static Models: Some Monte Carlo Evidence," *Oxford Bulletin of Economics and Statistics,* 48:253–277.

Bilson, J. (1981): "The Speculative Efficiency Hypothesis," *Journal of Business,* 54:435–451.

Brenner, R., and Kroner, K. (1992): "Arbitrage and Cointegration," Discussion Paper, University of Arizona.

Chowdhury, A. (1991): "Futures Market Efficiency: Evidence from Cointegration Tests," *The Journal of Futures Markets,* 11:577–589.

Dickey, D., and Fuller, W. (1979): "Distribution of Estimators for Autoregressive Time Series with a Unit Root," *Journal of the American Statistical Association,* 74:427–431.

Dwyer, G., and Wallace, M. (1992): "Cointegration and Market Efficiency," *Journal of International Money and Finance,* 11:318–328.

Engle, R., and Granger, C. (1987): "Cointegration and Error Correction: Representation, Estimation, and Testing," *Econometrica,* 50:987–1008.

Garbade, K. (1982): *Securities Markets.* New York: McGraw-Hill.

Gibson, R., and Schwartz, E. S. (1990): "Stochastic Convenience Yield and the Pricing of Oil Contingent Claims," *Journal of Finance,* 45:959–976.

Gonzalo, J. (1989): "Comparison of Five Alternative Methods of Estimating Long Run Equilibrium Relationships," Working Paper, UCSD.

Granger, C. W. J., and Newbold, P. (1974): "Spurious Regressions in Econometrics," *Journal of Econometrics,* 2:111–120.

Hakkio, C., and Rush, M. (1989): "Market Efficiency and Cointegration: An Application to the Sterling and Deutschemark Exchange Markets," *Journal of International Money and Finance,* 8:75–88.

Hansen, L., and Hodrick, R. (1980): "Forward Exchange Rates as Optimal Predictors of Future Spot Rates: An Econometric Analysis," *Journal of Political Economy,* 88:829–853.

Johansen, S. (1989): "Statistical Analysis of Cointegrated Vectors," *Journal of Economic Dynamics and Control,* 12:231–254.

Johansen, S. (1991): "The Role of the Constant Term in Cointegration Analysis of Non-Stationary Variables," Discussion Paper, University of Copenhagen.

Kroner, K., and Sultan, J. (1991): "Time Varying Distributions and Dynamic Hedging with Foreign Currency Futures," Manuscript, Center for the Study of Futures Markets, Columbia Business School.

Lai, K., and Lai, M. (1991): "A Cointegration Test for Market Efficiency," *The Journal of Futures Markets,* 11:567–576.

Osterwald-Lenum, M. (1990): "Recalculated and Extended Tables of the Asymptotic Distribution of some Important Maximum Likelihood Cointegration Test Statistics," Discussion Paper, University of Copenhagen.

Pantula, S. (1991): "Asymptotic Distributions of Unit-Root Tests when the Process is Nearly Stationary," *Journal of Business and Economic Statistics,* 9:63–71.

Said, S., and Dickey, D. (1984): "Testing for Unit Roots in Autoregressive-Moving Average Models of Unknown Order," *Biometrika,* 599–608.

Schwert, G. (1987): "Effects of Model Specification on Tests for Unit Roots in Macroeconomic Data," *Journal of Monetary Economics,* 20:73–103.

Schwert, G. (1989): "Tests for Unit Roots: A Monte Carlo Investigation," *Journal of Business and Economic Statistics,* 7:147–160.

Stock, J. (1987): "Asymptotic Properties of Least Squares Estimates of Cointegration Vectors," *Econometrica,* 55:1035–1056.