Noninformative and Informative Tests of Efficiency in Three Energy Futures Markets

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This article presents a critique of tests of market efficiency commonly applied to energy futures markets. Most of this literature fails to deal adequately with the endogeneity, nonstationarity, and cointegration characteristics of spot and futures prices, resulting in tests that are not informative about market efficiency. Consistent with this literature, application of these noninformative tests to spot and futures prices from three energy markets is generally not supportive of market efficiency. The article also presents two alternative tests of efficiency that properly deal with the stochastic features of these price series. Nonstationary data can be modeled with cointegration methods, allowing valid tests of the efficiency hypothesis. Alternatively, testing the equivalence of the data generation processes of the spot and futures prices provides an informative approach to efficiency testing with stationary data that does not suffer from endogeneity problems. Application of these two tests is largely supportive of weak and semi-strong efficiency in three energy futures markets. © 1998 John Wiley & Sons, Inc. Jrl Fut Mark 18: 939–964, 1998

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INTRODUCTION

A substantial literature has presented tests of the efficiency of futures markets using a variety of methodologies. The results of these tests across various commodity and financial futures markets are as varied as the test procedures they employ. Unfortunately, most tests applied to the energy futures markets fail to allow for the fundamental stochastic properties of spot and futures prices. In this study evidence is presented on prices from three energy markets in support of the following propositions: (i) these price series are integrated stochastic processes, requiring differencing to become stationary; (ii) spot and futures prices of a given commodity are cointegrated, sharing a common stochastic trend; and (iii) in regressions involving spot and lagged future prices, the presence of a risk premium introduces a correlation between the regressor and the error term, inducing a potential simultaneity bias. These three characteristics of commodities prices require the use of special procedures for modeling and testing the relations between spot and future price series.

The objectives of this research are twofold. First, tests that have been applied to energy futures markets are reviewed, in order to establish that these procedures are noninformative about market efficiency. Tests are noninformative if they are based on regressions in which either (i) parameter estimates are biased and inconsistent, or (ii) distributions governing regression test statistics have nonstandard distributions. This article describes how the failure to deal with the three characteristics of commodity prices described above renders previous tests noninformative. Generally, these tests of energy futures markets are representative of the wider literature that examines market efficiency in financial and commodities futures markets, and the critique of these tests applies with equal force to these nonenergy markets. Second, two test procedures that are informative about market efficiency are described, and these tests are applied to three energy futures markets. These informative tests do not suffer from problems of inconsistent estimators or nonstandard distributions. Contrary to the evidence from the noninformative tests, these two test procedures provide evidence that is generally consistent with market efficiency in these three markets.

METHODOLOGICAL ISSUES IN TESTS OF MARKET EFFICIENCY

The efficient markets hypothesis (EMH) is a joint hypothesis of rational expectations and the absence of risk-free excess returns. For a commodity spot and futures market, define

S_t = the natural logarithm of the spot price at time t;

F_{t-1} = the natural logarithm of the price at time $t - 1$ of a one month futures contract for delivery at time t;

$E_{t-1}(S_t) = E(S_t/I_{t-1})$, which expresses the expectation of the spot price at time t, conditional on the information set I_{t-1}.

The rational expectations hypothesis states that participants in the market exploit all available information in forming their expectations about the future spot price. In this case the forecast error,

$$e_t = S_t - E_{t-1}(S_t) \tag{1}$$

must be uncorrelated with the information set I_{t-1}. Assuming zero transactions costs and risk neutrality, the futures price will be forced by market agents to equal the expected spot price. However, under risk aversion speculators may be compensated for holding a risky futures contact by a risk premium (v_{t-1}), so that the futures price is discounted relative to the expected future spot price:

$$F_{t-1} = E_{t-1}(S_t) - v_{t-1} \tag{2}$$

In this formulation the risk premium is time varying and it may have an expected value that differs from zero. An alternative interpretation of v_{t-1} is as a measure of the noise in F_{t-1} as a proxy for the expected future spot rate (Frenkel, 1981).

Combining eqs. (2) and (1) yields

$$S_t - F_{t-1} = e_t + v_{t-1}$$

which may be expressed in the form of a testable model,

$$S_t = \alpha + \beta F_{t-1} + e_t + v_{t-1} \tag{3}$$

In this model the futures price is an unbiased predictor of the spot price if $\alpha = 0$ and $\beta = 1$.[1] If, in addition, the equation error, $e_t + v_{t-1}$, is unpredictable from past information, including its own past values, then F_{t-1} is an efficient predictor of S_t. The markets governing the spot and futures prices are said to satisfy the efficient markets hypothesis (EMH).

[1]If the risk premium has a nonzero mean, this component would appear in the intercept of eq. (3). Rejection of the hypothesis that α equals zero is often taken as evidence of a risk premium, rather than a rejection of market efficiency. However, it is the presence of a time-varying component of the risk premium that creates the econometric problems discussed throughout this article. These econometric problems arise from the presence of an error term in eq. (2), whether this term reflects a risk premium or noise in the use of the futures price as a measure of the expected spot price.

The EMH can be expressed in terms of the data generating processes (DGP) of the spot and futures prices. Following Barnhart, McNown, and Wallace (1996), market efficiency implies that the data generating processes of spot and futures prices should be equivalent, except for the forecast error and the risk premium.

$$S_t = g(X) + e_t \tag{4}$$

$$F_{t-1} = g(X) - v_{t-1} \tag{5}$$

where g(X) is a function of lagged variables that determine the formation of spot and futures prices[2] Subtracting eq. (5) from eq. (4) again yields eq. (3).

Tests of the efficient markets hypothesis require careful attention to the stochastic properties of the series in eq. (3). There is considerable evidence that spot and futures prices from commodity and financial markets are integrated processes, requiring first differencing to become stationary. Formally, a time series, x_t, is integrated of order d [I(d)] if it must be differenced exactly d times to become stationary. A process that is integrated of order one is also said to contain a unit root in its autoregressive component. Statistics derived from such time series are not governed by standard probability distributions, so that tests for a unit root involve special distributions. Most unit root testing employs the procedures developed by Dickey and Fuller (1979), which Schwert (1989) has established have reasonable size and power characteristics relative to available alternatives. Applications of the Dickey-Fuller test to energy futures have been unable to reject the null hypothesis of a unit root in these price series; see for example, Quan (1992), Crowder and Hamed (1993), Schwarz and Szakmary (1994), Serletis (1994).[3]

Early tests of unbiasedness in the relation between spot and futures prices employed simple regressions of spot prices on futures prices at a lag equal to the contract length,

[2]These could be lagged spot and futures prices of the market considered, lagged spot and futures prices of related markets, or other relevant macroeconomics variables.

[3]Schwarz and Szakmary (1994) use the augmented Dickey-Fuller (ADF) test for daily data of spot and futures prices (one month contract) of heating oil no. 2 (HO) and west Texas intermediate (WTI), for the period: 1 January 1984–15 May 1991, and for unleaded gasoline (UG), for the period: 1 January 1985–15 May 1991. Serletis (1994) uses the ADF test for daily data of futures prices (one month contract) of HO, WTI and UG for the period: 12 March 1984–30 April 1993. Quan (1992) uses the ADF test for monthly data of spot and futures prices (1, 3, 6, 9 monthly contracts) of WTI, for the period: January 1984–July 1989. Crowder and Hamed (1993) use the ADF test for monthly data of spot and futures prices (one month contract) of WTI, for the period: March 1983–September 1990. All these studies found that the series investigated are I(1).

$$S_t = \alpha + \beta F_{t-1} + \mu_t \tag{6}$$

In this relation the null hypothesis of market efficiency can not be rejected if $\alpha = 0$, $\beta = 1$, and μt is free of serial correlation. Goss (1981), Tomek and Gray (1970), and Bigman, Goldfarb, and Schechtman (1983) estimated such relations for various commodity markets by ordinary least squares, testing the null hypothesis with conventional regression statistics. However, if the series in this relation are I(1), these test statistics do not adhere to conventional distribution theory (Elman and Dixon, 1988; Shen and Wang, 1990).

The most severe consequences of nonstationary time series are ameliorated in regressions involving stationary linear combinations of integrated variables. In the context of the present model, even if S_t and F_{t-1} are I(1), it is possible that the linear relation between them,

$$S_t - \alpha - \beta F_{t-1} = \mu_t \tag{6'}$$

is stationary. In this case S_t and F_{t-1} are said to be cointegrated, and (6') defines the cointegrating equation. Evidence in support of cointegration between spot and lagged futures prices for West Texas intermediate crude is found by Quan (1992), Crowder and Hamed (1993) and Schwarz and Szakmary (1994). The last article also presents evidence of cointegration in the market for one-month heating oil and unleaded gasoline contracts.

Cointegration carries three important implications for testing market efficiency: (i) least squares estimators of the parameters of a cointegrating equation are (super)consistent, although they are biased in small samples (Stock, 1987); (ii) if the regressor is endogenous or if the equation errors are autocorrelated, then test statistics from this least squares regression do not have conventional distributions[4] and (iii) there exists an error correction model that describes the dynamic relation between cointegrated time series.

The second implication has generally been overlooked in testing the efficiency of futures markets. Chowdhury (1991) rejected the unbiasedness hypothesis ($\alpha = 0$ and $\beta = 1$) for markets in several nonferrous metals using least squares regression statistics without acknowledgment of their nonstandard distributions. Serletis and Banack (1990) used the Engle-Granger methodology to test for cointegration between spot and

[4]Hamilton (1994, pp. 602–603) describes the distributional properties of least squares estimators in the special case of exogenous regressors and serially uncorrelated errors. The analytical argument is summarized on page 603: "...despite the I(1) regressors and complications of cointegration, the correct approach for this example would be to estimate by OLS and use standard t or F statistics to test any hypotheses about the cointegrating vector."

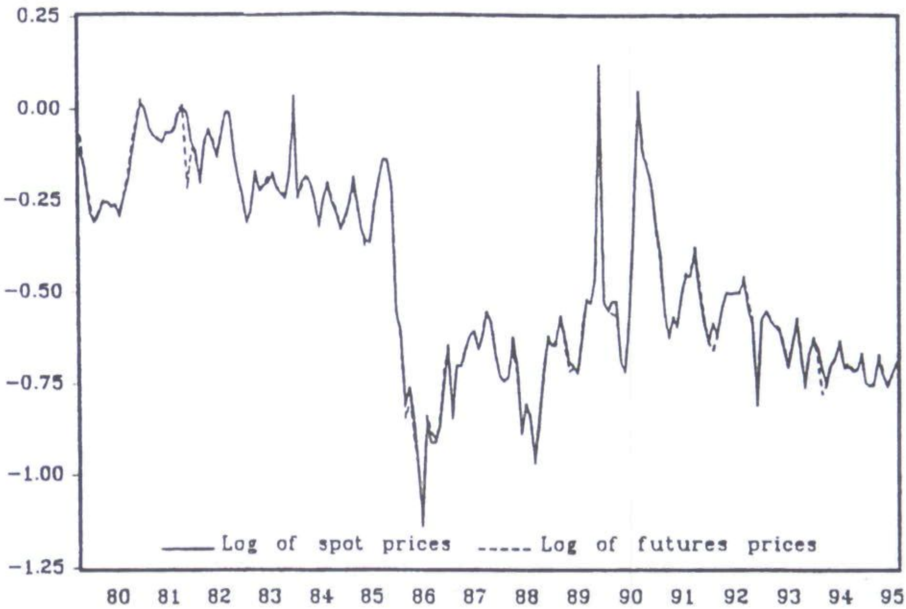

FIGURE 1
Log of current spot and futures prices of heating oil: 11/1979–8/1995.

futures prices of heating oil, West Texas intermediate and unleaded gasoline.[5] They found that the series were cointegrated and therefore they concluded that the markets are efficient, and the futures price is an unbiased predictor of the spot price. However, as Hsieh (1990) pointed out in his commentary on the article, they did not test the restrictions of the unbiasedness hypothesis in the cointegrating vector. Moreover, they did not test as a further condition for market efficiency the absence of serial correlation in the errors of the cointegrating equation.

Moosa and Al-Loughani (1994) test the three and six month futures markets for West Texas intermediate using West's (1988) adjustment of the standard errors. This procedure is designed to correct for serial correlation in the equations errors to produce asymptotically normal test statistics, even with integrated regressors. However, the West procedure requires that the nonstationary regressor have a nonzero drift component, which is not apparent in any of the series displayed in Figures 1–3 presented here. Moosa and Al-Loughani's rejection of unbiasedness for these markets is therefore subject to question.

Phillips and Loretan (1991) present a nonlinear estimation procedure that corrects for both serial correlation in the errors and endogeneity

[5]Serletis and Banack (1990) used the futures prices at one month contract as a proxy for the spot prices, and futures prices at two month contracts as a proxy for futures prices at one month.

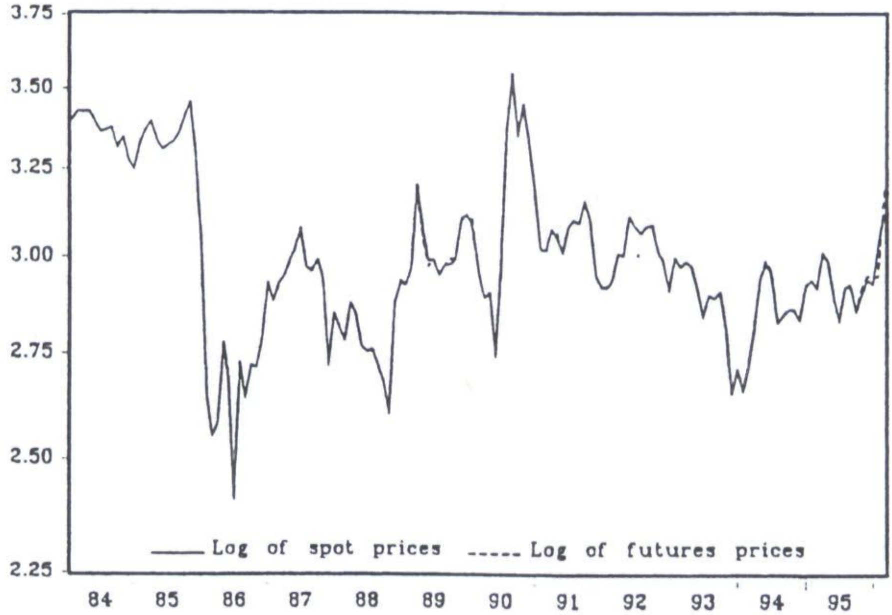

FIGURE 2
Log of current spot and futures prices of West Texas Intermediate: 1/1934–3/1996.

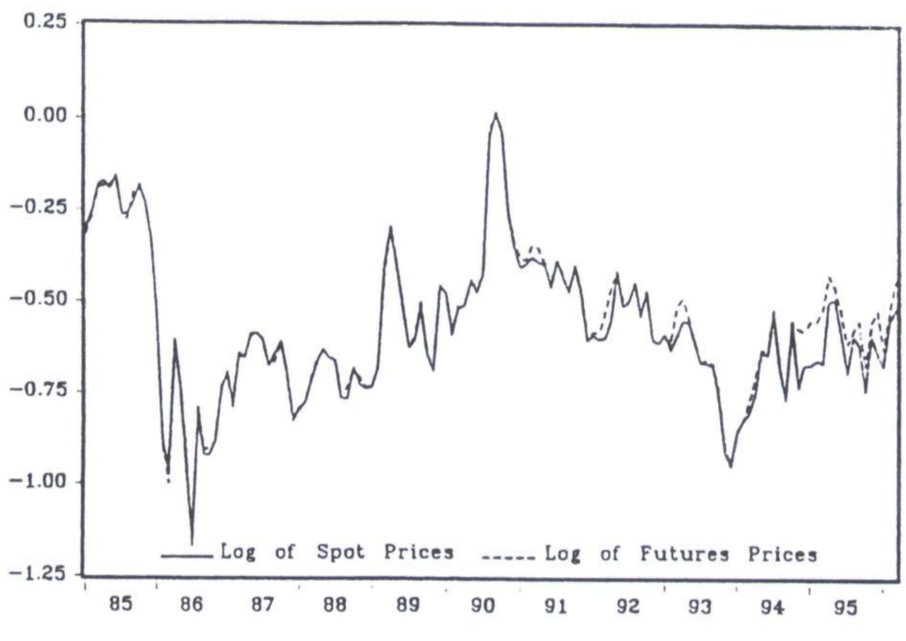

FIGURE 3
Log of current spot and futures prices of unleaded gasoline: 1/1985–3/1996.

of the regressors in cointegrating regressions involving general I(1) regressors, allowing informative tests of the market efficiency hypothesis.

$$S_t = \alpha + \beta F_{t-1} + \sum_{i=1}^{n} \eta_i (S_{t-i} - \alpha - \beta F_{t-i-1}) + \sum_{j=-L}^{L} \phi_j \Delta F_{t-j} + u_t \qquad (7)$$

Nonlinear least squares estimation of this equation produces regression test statistics with standard asymptotic distributions, allowing tests of the unbiasedness hypothesis that $\alpha = 0$ and $\beta = 1$. The additional condition that the error in the spot-futures equation not be autocorrelated is tested with the Ljung-Box (1978) Q-statistics applied to the residuals constructed from eq. (6), with α and β replaced by their Phillips-Loretan estimate.[6]

An alternative approach to handling the nonstationarity of spot and futures prices in eq. (6) is to transform the equation to a form that contains only stationary variables. One such approach is suggested by subtracting S_{t-1} from both sides of (6), yielding the "basis regression"

$$S_t - S_{t-1} = \alpha + \beta(F_{t-1} - S_{t-1}) + \mu_t \qquad (8)$$

As in eq. (6') unbiasedness requires that $\alpha = 0$ and $\beta = 1$. However, as demonstrated by Liu and Maddala (1992), this test equation suffers from simultaneity bias, rendering tests based on (8) noninformative. As shown by comparing eqs. (3) and (6), the error in eq. (8) contains the risk premium, which is correlated with the regressor through the lagged futures price. Serletis (1991), following Fama and French (1987), estimates this relationship for heating oil, unleaded gasoline and crude oil and he rejects the hypothesis that $\beta = 1$, which is interpreted as evidence for a time varying risk premium.

Fujihara and Mougou' (1997),[7] building on Kaminsky and Kumar (1990),[8] tested the weak form of efficiency for the crude oil, the heating oil and the unleaded gasoline futures markets. They regressed the excess

[6]Johansen's (1988) maximum likelihood procedure provides an appropriate alternative to testing hypotheses on cointegrating equations. Crowder and Hamed (1993) apply Johansen's method to the market for West Texas intermediate crude oil, finding evidence in support of market efficiency.

[7]Fujihara and Mougou' (1997) used daily data for the period: from 12 March 1984–30 September 1993.

[8]Kaminsky and Kumar (1990) tested the weak and semistrong form of efficiency for the following commodities futures markets: corn, soybean, wheat, cocoa, coffee, copper, and cotton using futures contracts for 1,3,6, and 9 months. They found mixed evidence for market efficiency.

futures returns from a futures market on own lagged excess futures returns,

$$\Delta F_{t-1} = \alpha + \sum_{t=2}^{n} \beta_i \Delta F_{t-i} + \varepsilon_t \tag{9}$$

They tested if all $\beta_i = 0$, claiming that market efficiency requires unpredictability in futures returns. Their tests rejected market efficiency for the heating oil and unleaded gasoline futures markets, but not for the crude oil market. However, Dwyer and Wallace (1992) have shown that predictability in futures returns depends on the data generating process (DGP) governing futures prices, and, therefore, it has no relevance to questions of market efficiency.

Alternatively, Hakkio and Rush (1989) suggest tests based on a renormalization of the cointegrating equation of the form,

$$\begin{aligned} S_t - S_{t-1} &= \alpha + \beta_1(F_{t-1} - F_{t-2}) + \beta_2(F_{t-2} - S_{t-1}) \\ &\quad + \sum_{i-1}^{n} \eta_i \, \Delta S_{t-i} + \sum_{j=2}^{m} \delta_j \, \Delta F_{t-j} + \mu_t \end{aligned} \tag{10}$$

or

$$S_t - S_{t-1} = \alpha + \beta_1(F_{t-1} - F_{t-2}) + \beta_2(F_{t-2} - S_{t-1}) + \mu_t \tag{11}$$

Market efficiency requires that $\beta_1 = \beta_2 = 1$ in both equations, and $\eta_i = \delta_j = 0 \ \forall$ i and j in eq. (10). These conditions correspond to a slope coefficient of unity in (6), and the absence of predictability of the error in (6) based on past values of spot and futures prices. Moosa and Al-Loughani (1994) apply this model to three and six month futures markets for West Texas intermediate crude, rejecting the stated conditions for market efficiency. However, the Hakkio and Rush equations are plagued with the same endogeneity problem as the basis regression (8). The risk premium component of μ_t, which contains v_{t-1}, is correlated with terms involving F_{t-1} as implied by eq. (5). Barnhart, McNown, and Wallace (1996) demonstrate analytically and through simulations that the Hakkio-Rush procedure will generally reject unbiasedness in markets that are truly characterized by market efficiency. The simultaneity problem in eqs. (10) and (11) renders these tests noninformative.

Informative tests using stationary data can be based on the principle that market efficiency requires the equivalence of the data generation processes (DGPs) of the spot and lagged futures price series. Cointegration implies that the DGPs of these two series are of the error correction form[9] (Engle and Granger, 1987)

$$S_t - S_{t-1} = \alpha + \delta(F_{t-2} - S_{t-1}) + \sum_{i=1}^{m} \eta_i \, \Delta S_{t-i} + \sum_{j=2}^{n+1} \phi_j \Delta F_{t-j} + \varepsilon_t \tag{12}$$

$$F_{t-1} - S_{t-1} = \alpha + \delta(F_{t-2} - S_{t-1}) + \sum_{i=1}^{m} \eta_i \Delta S_{t-i} + \sum_{j=2}^{n+1} \phi_j \Delta F_{t-j} + v_{t-1} \tag{13}$$

Here the DGPs are expressed only in term of own lagged spot and futures prices. More general forms are considered in the empirical testing.

This review of methods and models for testing efficiency in energy futures markets demonstrates that failure to take into account the nonstationarity and endogeneity characteristics of spot and futures prices undermines the credibility of the results presented in this literature. Excepting Crowder and Hamed (1993), researchers who have employed energy price data in nonstationary form have not applied methods that yield valid regression test statistics. Moreover, tests based on stationary transformations of the price series have ignored the problems of endogeneity endemic to these specifications. Fortunately, informative test procedures do exist, as described above. One objective of this study is to apply these informative tests to three energy futures markets. In addition, to demonstrate the marked differences between results based on informative and noninformative procedures, the complete battery of tests described above is applied to data from the same three energy markets.

STOCHASTIC PROPERTIES OF ENERGY SPOT AND FUTURES PRICES

This study employs monthly observations[10] on spot prices and futures prices from one month contracts for West Texas intermediate (crude oil), heating oil no. 2 and unleaded gasoline, traded at NYMEX. The futures

[9]These DGPs are renormalized by subtracting S_{t-1} from both S_t and F_{t-1} equations, so that the basis, rather than the change in F_{t-1}, appears on the left side of the second error correction equation.
[10]Using monthly observations instead of daily observations, eliminates the overlapping forecast errors problem described by Hansen and Hodrick (1981).

and spot prices series for the three commodities cover the following periods:

—from 01-1984 to 03-1996 for West Texas intermediate;

—from 11-1979 to 08-1995 for heating oil no. 2;

—from 01-1985 to 03-1996 for unleaded gasoline.

The futures price series for the West Texas intermediate and the unleaded gasoline are from Tick Data Inc., excepting the period from 10-1995 to 3-1996, which are from the *Wall Street Journal*. The futures and spot prices for heating oil n. 2 are also from Tick Data Inc. The spot price series of West Texas intermediate and unleaded gasoline are from Platt's Commodities Division of Standard and Poor.

The monthly futures and spot prices for heating oil n. 2 and unleaded gasoline are the closing prices for the last trading day, which is the last business day of the month. For West Texas intermediate futures and spot prices, closing prices are for the last trading day, which is the third business day before the 25th of the month preceding the delivery month.[11] All series are analyzed in natural logarithmic form.

Visual inspection of the plots in Figures 1–3 suggests the prices series are individually nonstationary, but the spot and futures prices for each market appear to share common stochastic trends. Augmented Dickey-Fuller (ADF) tests are applied to all six series following the lag search procedure suggested by Ng and Perron (1995). Although other tests of nonstationarity are available, the simulation evidence produced by Schwert (1989) indicates the augmented Dickey-Fuller test has reasonable size and power characteristics, except in the case of integrated series with a strong moving average component. Examination of the correlogram of the first differences of the six price series does not indicate the presence of a large moving average component, so the Dickey-Fuller test is appropriate in these cases. Starting with a 24-lag specification of the Dickey-Fuller equation,

$$\Delta y_t = a_0 + \delta t + \gamma y_{t-1} + \sum_{j=1}^{p} \eta_j \Delta y_{t-j} + \varepsilon_t \tag{14}$$

insignificant lags are eliminated, based on t- and F-statistics. The final

[11]Usually, the delivery period lasts one month. For heating oil and unleaded gasoline futures contracts, the buyer (long position) specifies a 5 day delivery window inside the delivery month, and usually the delivery takes place on the first day of this window (Ma, 1989). For the West Texas intermediate, the date of delivery depends on the dispatching schedule of the pipeline companies. However, as Ma (1989, pp. 398) has pointed out: "An examination of the convergence patterns of crude oil futures prices indicates that the future price predicts the spot price on the last trading day." Therefore, the data used in this study is consistent with this assessment.

specifications, summarized in Table I are tested for residual autocorrelation with the Ljung-Box Q-statistic (Ljung and Box, 1978).

For all six series the test statistics reported in Table I fail to reject the unit root hypothesis. Repeating the tests without the trend term leads to the same conclusion for the spot and futures prices of heating oil and West Texas intermediate crude. However, the results for unleaded gasoline are less clear. The Ng-Perron criterion for choice of lagged differences leads to a model with a single term at lag two, and this specification rejects the unit root hypothesis for spot prices, but not for futures prices, at the 5% level. Based on the test results in Table I, the strong similarity between the spot and futures prices shown in Figure 3, and other evidence in support of the unit root hypothesis for this series reported by Schwarz and Szakmary (1994) and Serletis (1994), the analysis proceeds under the assumption that spot prices of unleaded gasoline are also integrated of order one.

NONINFORMATIVE AND INFORMATIVE TESTS OF ENERGY MARKET EFFICIENCY

In the relation between spot and futures prices in their nonstationary (levels) form, market efficiency holds if the following three conditions are satisfied: (i) spot and futures prices are cointegrated; (ii) the slope coefficient in this cointegrating equation equals one; and (iii) the errors in this equation are not predictable using information available in any prior period. Cointegration in all three markets is supported by the ADF test on the residuals from the cointegrating equation, as reported in Table II. In addition, the nonlinear Phillips-Loretan estimates of the cointegrating equation yield estimates of slope parameters that are numerically very close to 1.0 (Table III). Nor can one reject the individual hypotheses that the slopes and intercepts are equal to one and zero, respectively, or the joint hypothesis on the slope and intercept. The lack of signficance of the intercepts in each model implies the absence of a nonzero mean in the risk premium, although this does not rule out the existence of a time-varying risk premium or other stochastic term, v_{t-1}, in the process generating the futures prices. Finally, there is no evidence of serial correlation in the residuals of the equation for West Texas intermediate crude based on Ljung-Box Q-statistics computed with 4 and 12 lagged autocorrelations. There are weak indications of autocorrelation for heating oil and unleaded gasoline. However, Q-statistics are significant at only the 10% level and in both cases the significance is sensitive to the lag length.

TABLE I

Dickey-Fuller Tests for Unit Roots in the Logarithms of Spot and Futures Prices of Heating Oil, West Texas Intermediate, and Unleaded Gasoline

		ADF with trend and constant			ADF with constant	
	Lags	Trend δ	$H_0{:}\gamma=0$ τ	Q_{16}[a]	$H_0{:}\gamma=0$ τ	Q_{16}[a]
Heating oil spot prices	7	-0.0003 (-1.63)	-2.14	9.94	-1.38	10.32
Heating oil futures prices	7	-0.0003 (-1.59)	-2.11	7.09	-1.37	7.44
West Texas Int. spot prices	16	0.0001 (0.54)	-2.43	1.43	-2.73	1.35
West Texas Int. futures prices	16	0.0001 (0.66)	-2.39	1.36	-2.72	1.27
Unleaded gasoline spot prices	2[b]	9.388E-05 (0.39)	-2.91	6.12	-2.98*	6.04
Unleaded gasoline futures prices	2[b]	0.0001 (0.740)	-2.76	7.85	-2.79	7.51

MacKinnon critical values with trend and constant, τ_τ: -4.01 (1%), -3.43 (5%) and -3.14 (10%) for lags 7 and 2*. For lags 16, τ_τ: 4.03 (1%), -3.45 (5%) and -3.15 (10%).
MacKinnon critical values with constant only, τ_μ: $-$ 3.47 (1%), -2.88 (5%) and -2.7 (10%) for lags, 7 and 2*. For 16 lags, τ_μ: -3.48 (1%), -2.88 (5%) and -2.58 (10%).
** = significant at 0.01; * = significant at 0.05; + = significant at 0.1.
For the trend *t* statistics are in parenthesis.
[a]Ljung-Box Q statistics
[b]The first lag is dropped because it is insignificant.
Reported test statistics are based on the Dickey-Fuller regression,

$$\Delta y_t = a_0 + \delta t + \gamma y_{t-1} + \sum_{j=1}^{p} \eta_j \Delta y_{t-j} + \varepsilon_t$$

Although the Phillips-Loretan estimates are consistent with market efficiency, simple least squares estimation of cointegrating regressions yields misleading results. As shown in the second line of each panel in Table III, the point estimates of the slope coefficients are substantially less than 1, and the t-statistics reject the hypothesis of unity in every case. The joint hypothesis, $\alpha = 0$ and $\beta = 1$ is also rejected at the 5% level by the F-statistics. A comparison of least squares and Phillips-Loretan results points out the importance of appropriate cointegration methods in modeling nonstationary data.

TABLE II

Engle-Granger Test of Noncointegration–Dickey-Fuller Tests of the Residuals from the Cointegrating Equations

Residuals	H_0: $\gamma=0$ τ	Lags	Q_{12}	Prob.
Heating Oil	-3.91**	8	1.12	1
	-14.78**	0	15.96	0.19
Crude Oil (WTI)	-4.18**	15	0.90	1
	-10.33**	0	7.88	0.79
Unleaded Gasoline	-9.18**	1	5.62	0.99
	-10.41**	0	8.28	0.76

Engle and Yoo critical values for the model with nonzero lags: −3.73 (1%), −3.17 (5%) and −2.91 (10%); Engle and Yoo critical values for the model without lags −4.07 (1%), −3.37 (5%) and −3.03 (10%).
** = significant at 0.01; * = significant at 0.05; + = significant at 0.10.
Ljung-Box Q-statistics.
Reported test statistics are based on the residuals from the cointegrating regression, $S_t - \alpha - \beta F_{t-1} = \mu_t$

Pursuing further the question of predictability of the forecast errors from the cointegrating equation, the forecast errors from eq. (6) are regressed on lagged forecast errors from all three markets. Since the intercept and slope in each cointegrating equation are not statistically different than zero and one, respectively, the forecast errors may be computed as $(S_t - F_{t-1})$ for each market. Tables IV and IVa present joint tests of significance (F-tests) of these regressions containing both one-lag and two-lag specifications. These semi-strong tests are unable to reject market efficiency for the heating oil and unleaded gasoline markets, but do show 5% level significance for both the one- and two-lag specifications of the crude oil regressions. Overall, the cointegration based tests provide evidence of semi-strong efficiency for the heating oil and unleaded gasoline markets, but only weak-form efficiency for the market for West Texas intermediate.

This evidence in favor of market efficiency contradicts the results of the noninformative tests based on regressions with stationary data. Table V reports the results of a replication of the test used by Fujihara and Mougoue' (1997). This Table displays F-tests of the significance of regressions such as eq. (9) based on a search for significant lags out to a maximum of 14 months. In particular, nonpredictability of returns in spot

TABLE III

Market Efficiency Tests Using the Estimates from the Nonlinear and Linear (OLS) Least Squares Estimates of Cointegrating Equations

Spot Prices (S_t)	α	F_{t-1}	t_1 $H_0{:}\beta=1$	F test $H_0{:}\alpha=0\ \beta=1$	Q_4[a]	Q_{12}[a]
Heating oil	0.0027	1.0024	0.3199	0.41	5.09	19.43+
	(-0.68)	(132.23)				
	-0.0358	0.9243	-2.74	3.78		
	(2.42)	(33.50)				
West Texas Int.	-0.0001	0.9999	0.0233	0.02	2.61	12.66
	(-0.009)	(269.24)				
	0.3344	0.8878	-3.05	4.69		
	(3.02)	(24.18)				
Unleaded Gasoline	-0.0173	0.9987	-0.0246	1.98	9.36+	11.29
	(-0.59)	(20.37)				
	-0.0984	0.8538	-3.21	6.76		
	(-3.62)	(18.77)				

For every commodity, the first line shows the estimates from the Phillips-Loretan equation

$$S_t = \alpha + \beta F_{t-1} + \sum_{i=1}^{n} \eta_i (S_{t-i} - \alpha - \beta F_{t-i-1}) + \sum_{j=-L}^{L} \phi_j \Delta F_{t-j} + u_t.$$

The second line in each panel displays the least squares estimates of $S_t - \alpha - \beta F_{t-1} = \mu_t$.
The t statistics are in parenthesis.
The t critical values are: 2.576 (1%), 1.96 (5%) and 1.645 (10%).
** = significant at 0.01; * = significant at 0.05; + = significant at 0.10.
[a]Ljung-Box Q-statistics for autocorrelation in the residuals computed as $e_t = S_t - \alpha - \beta F_{t-1}$, where α and β are the estimates from the non linear least squares estimates of the cointegrating equations.

and futures markets is strongly rejected for heating oil and unleaded gasoline. These results conform with those of Fujihara and Mougoue, but conflict with those from the informative tests. This confirms the irrelevance of tests of market efficiency based on unpredictability of returns (Dwyer and Wallace, 1992).

Following Moosa and Al-Loughani (1994) efficiency is tested using Hakkio and Rush's (1989) renormalization of the cointegrating equation, eqs. (10) and (11). Hakkio and Rush contend that market efficiency requires that $\beta_1 = \beta_2 = 1$ under both specifications, and that $\eta_i = \delta_j = 0$, for all i and j in the first specification. Specifications of eq. (10) are based on the lag lengths found appropriate for the error correction models specified below.

TABLE IV

Semi-Strong Test of Efficiency: Testing Predictability Using Past Information from Other Markets; Period: 1 February 1985–1 August 1995

(S_t-F_{t-1})	Intercept	$(S_{t-1}-F_{t-2})_{HO}$	$(S_{t-1}-F_{t-2})_{WTI}$	$(S_{t-1}-F_{t-2})_{UG}$	F test	Prob.	Q_{12}
Heating oil	-0.0002 (-0.01)	-0.1976 (-1.84)	0.1463 (1.18)	0.0284 (0.22)	0.98	0.42	19.47+
West Texas int.	-0.0004 (-0.04)	0.1927 (2.04)	-0.1359 (-1.25)	0.2224 (2.04)	3.45*	0.01	8.96
Unleaded gasoline	-0.0153 (-1.54)	0.1167 (1.16)	-0.0182 (-0.15)	0.0012 (0.01)	1.04	0.38	10.55

The *t* statistics are in parenthesis.
The null hypothesis is that all coefficients are jointly zero.
** = significant at 0.01; * = significant at 0.05; + = significant at 0.10.
Ljung-Box Q-statistics.

TABLE IVa

Semi-Strong Form of Efficiency: Specification with Two Lags

	F Statistics	Prob.
Heating oil	1.58	0.14
West Texas int.	2.6*	0.015
Unleaded gasoline	1.69	0.11

Test statistics are based on regressions involving the same variables as in Table 4, but with one additional lag on each right hand side variable. F-statistics test the hypothesis that all slope coefficients are jointly zero.

TABLE V

Predictability of Energy Spot and Futures Returns

Heating oil				Unleaded gasoline				West Texas crude			
Spot returns		Futures returns		Spot returns		Futures returns		Spot returns		Futures returns	
Lags	F stat	Lags	F stat	Lags	F stat	Lags	F stat	Lags	F stat	Lags	F stat
1	2.82 (0.09)			2	3.58 (0.03)	2	4.39 (0.014)	NO PREDICTABILITY from lag 1 to 14			
6	2.11 (0.054)	6	2.11 (0.053)	3	2.52 (0.06)	3	3.13 (0.027)				
7	2.72 (0.010)	7	2.72 (0.010)			4	2.39 (0.053)				
8	2.72 (0.007)	8	2.59 (0.010)			5	1.94 (0.091)				

Probability in parenthesis.
For Heating Oil checked lags 1–8.
For Unleaded gasoline checked lags 1–14.
For West Texas crude checked lags 1–14.
Reported test statistics are based on regressions of first differences of spot (futures) prices on their own lagged differences,

$$\Delta F_{t-1} = \alpha + \sum_{i=2}^{n} \beta \Delta F_{t-i} + \varepsilon_t$$

TABLE VI

Estimates of the Hakkio and Rush Renormalization of the Cointegrating Equation

(S_t-S_{t-1})	Lags	Intercept α	$(F_{t-2}-S_{t-1})$ β_1	$(F_{t-1}-F_{t-2})$ β_2	F_1	t_1	t_2	Q_{12}	R^2	F
Heating oil	8	-0.0012 (-0.157)	1.2274 (1.253	1.41 (4.152)	1.70*	0.23	1.20	1.22	0.24	2.91**
	0	1.827E-05 (0.002)	1.6368 (5.13)	1.5351 (4.80)	2.7+	2*	1.68+	18.56+	0.12	13.5**
West Texas int.	14	-0.0048 (-0.520)	-4.2709 (-0.596)	-0.9439 (-0.572)	1.37	-0.73	-1.17	0.86	0.29	1.35
	0	-0.0016 (-0.19)	-0.7555 (-0.82)	-0.6597 (-0.72)	2.33	-1.92+	1,82+	11.97	0.02	0.93
Unleaded gasoline	2	-0.0115 (-1.079)	0.5198 (1.510)	0.5922 (1.523)	2.11+	-1.39	-1.05	2.69	0.08	1.91+
	14	-0.0081 (-0.685)	0.5902 (1.023)	0.9073 (1.974)	0.97+	-0.71	-0.20	4.85	0.24	0.97
	0	-0.0084 (-0.80)	0.4311 (1.47)	0.4314 (1.44)	1.9	-1.94+	-1.9+	11.63	0.01	1.09

The t statistics are in parenthesis.
Under column F_1: F statistics for the null hypothesis of $\beta_1 = 1$, $\beta_2 = 1$, $\eta_i = \delta_j = 0 \; \forall \; i$ and j (equations with lags) and for the null hypothesis of $\beta_1 = 1$, $\beta_2 = 1$ (equations without lags).
Under column t_1: t statistics for the null hypotheses of $\beta_1 = 1$.
Under column t_2: t statistics for the null hypotheses of $\beta_2 = 1$.
Estimates and tests are based on the Hakkio and Rush equation,

$$S_t - S_{t-1} = \alpha + \beta_1(F_{t-1} - F_{t-2}) + \beta_2(F_{t-2} - S_{t-1}) + \sum_{i=1}^{n} \eta_i \, \Delta S_{t-i} + \sum_{j=2}^{m} \delta_j \Delta F_{t-j} + \mu_t$$

Although the results presented in Table VI do not soundly reject market efficiency in every case, these regressions do not appear to be informative. The only regressions with significant overall explanatory power (5% level) are the equations for heating oil, and here at least one test statistic rejects market efficiency at the 5% level. For the other two commodities no regressions are significant at the 5% level. The standard errors of the estimated coefficients are too large to convincingly reject the stated null hypotheses, although at least one test statistic is significant at the 10% level in every equation but one (the long lag specification for West Texas intermediate). Overall, one would not interpret the results in Table VI as supportive of market efficiency, a conclusion that reflects the problem of simultaneity bias plaguing this model.

The evidence from the basis regressions (eq. 8) is equally unconvincing (Table VII). The explanatory power of these regressions is abysmally low, and point estimates of the slope coefficient are not numerically close to one. For West Texas intermediate the estimated coefficient is negative, and for unleaded gasoline the hypothesis of a slope of unity is rejected at the 0.05 level of significance. Simultaneity bias in these equations destroys the validity of these tests of hypothesis, whether supportive or contrary to the market efficiency hypothesis.

One valid test of efficiency that employs stationary data is the test of equivalence of the DGPs of the spot and futures prices. In particular, the equivalence of corresponding parameters in the spot and futures DGP equations is tested in their modified error correction form, eqs. (12) and (13). Tests are based on specifications of the DGPs that are limited to past information from the same market, as well as more general processes that also include lagged rates of change in futures prices from the other two energy markets. To accommodate the maintained hypothesis, lag lengths for the spot and futures price equations were forced to be identical within each market. The lag structure was determined by a standard search procedure. Starting with a maximum possible lag length of 16, insignificant lagged terms were eliminated based on likelihood ratio tests. Beginning with the most distant lags, terms were tested and eliminated in groups of four, with the search terminated if the tested group was found to be significant, or if evidence of autocorrelation arose.

Table VIII presents likelihood ratio tests of the equivalence of the DGPs, including only same-market information in the equations. Allowing a nonzero risk premium, the tests reported in the right-hand panels do not restrict the intercepts to be equal across the two equations. For West Texas intermediate and unleaded gasoline the equivalence of the two DGPs cannot be rejected, regardless of the treatment of the inter-

TABLE VII

Regressions of Rate of Depreciation on the Basis

$(S_t - S_{t-1})$	α	$(F_{t-1} - S_{t-1})$ (β)	t_1 $H_0{:}\beta=1$	F test $H_0{:}\alpha=0$ $\beta=1$	Prob.	R^2	σ_ε	Q_{16}	F statistics
Heating Oil	-0.0004 (-0.05)	1.5397 (4.88)	1.71	1.48	0.22	0.11	0.10	21.07	23.9**
West Texas Int.	-0.001 (-0.21)	-0.6639 (-0.72)	-1.82	1.69	0.18	0.003	0.10	21.82	0.53
Unleaded Gasoline	-0.007 (-0.75)	0.4175 (1.44)	-2.54*	3.56*	0.03	0.02	0.10	17.25	2.09

σ_ε = standard errors of regression.
The *t* statistics are in parenthesis.
The *t* critical values are: 2.576 (1%), 1.96 (5%) and 1.645 (10%).
** = significant at 0.01; * = significant at 0.05; + = significant at 0.10.
Ljung-Box Q-statistics.
Reported test statistics are based on the basis regression, $S_t - S_{t-1} = \alpha + \beta(F_{t-1} - S_{t-1}) + \mu_t$

TABLE VIII

Joint Tests of Equality of the Coefficients in the Data Generating Process of Spot and Futures Prices: Error Correction Equations

	Lags	$\alpha_1=\alpha_2$						$\alpha_1\neq\alpha_2$					
		λ_{LR}	$\chi^2(r)$ at 10%	$\chi^2(r)$ at 5%	T	C	r	λ_{LR}	$\chi^2(r)$ at 10%	$\chi^2(r)$ at 5%	T	C	r
Heating Oil													
$F_{t-2}-S_{t-1}$	8	25.348	25.989	28.869	180	18	18	25.348^{+}	24.769	27.587	181	18	17
West Texas I.													
$F_{t-2}-S_{t-1}$	14	33.754	40.256	43.773	131	30	30	33.708	39.087	42.557	131	30	29
Unleaded Gasoline													
$F_{t-2}-S_{t-1}$	14	27.860	40.256	43.773	119	30	30	25.529	39.087	42.557	119	30	29

** = significant at 0.01; * = significant at 0.05; + = significant at 0.10.
Null hypothesis for column $\alpha_1 = \alpha_2$: all coefficients in the spot rate equation are equal to the coefficients in the basis (forward premium) equation; null hypothesis for column $\alpha_1 \neq \alpha_2$: all coefficients in the spot rate equation are equal to the coefficients in the basis (forward premium) equation except for the intercept terms.
λ_{LR} = likelihood ratio statistic proposed by Sims (1980), which has an asymptotic χ^2 distribution with degrees of freedom equal to the number of restrictions (r).
Test statistics are based on the modified error correction equations,

$$S_t - S_{t-1} = \alpha + \delta(F_{t-2} - S_{t-1}) + \sum_{i=1}^{m} \eta_i \,\Delta S_{t-i} + \sum_{j=2}^{n+1} \phi_j \,\Delta F_{t-j} + \varepsilon_t$$

$$F_{t-1} - S_{t-1} = \alpha + \delta(F_{t-2} - S_{t-1}) + \sum_{i=1}^{m} \eta \Delta S_{t-i} + \sum_{j=2}^{n+1} \phi \Delta F_{t-j} + v_{t-1}$$

cepts. However, there is evidence against market efficiency for the heating oil market, but only for the case of differing intercepts at the 10% level.

The validity of these tests of DGP equivalence depends upon correct specification of these models. Generalizing these models to include lagged futures rates from the other two markets yields the tests reported in Table IX. Under these specifications the equality of each pair of DGPs is not rejected for any of the three markets. Generally, the evidence from these tests of equivalence of the DGPs is consistent with market efficiency.

CONCLUSION

In previous tests of efficiency in energy futures markets, most researchers have employed procedures that are not informative about the efficiency hypothesis. Early applications ignored the statistical problems inherent in regressions with nonstationary time series, using test statistics that do not have standard distributions. A more recent literature has treated the problem of nonstationarity by renormalizing the spot-futures price equation to a form containing only stationary variables. Unfortunately, when applied to markets with a futures risk premium, these renormalizations are plagued by the problem of simultaneity bias. The correlation between the regressor(s) and the error term in these stationary renormalizations causes inconsistent least squares estimators and consequently invalid inferences concerning the efficiency hypothesis.

In this article such problems are confirmed in applications of these noninformative procedures to tests of efficiency in three energy futures markets. In particular, simultaneity bias leads to an apparent rejection of market efficiency for futures markets for heating oil no. 2, unleaded gasoline, and West Texas intermediate crude oil. Tests of nonpredictability of returns also indicate inefficiency in spot and futures markets for two of three commodities examined, although this test is also noninformative.

However, application of two informative tests to these three markets yields evidence that is more supportive of market efficiency. Current methods and models of cointegration permit valid tests of the efficiency hypothesis in regressions with nonstationary time series. Application of these methods yields strong evidence that the spot and futures prices in these three markets are cointegrated in a relation with an intercept of zero and a slope equal to unity. In two of the markets there is weak evidence of serial correlation in the residuals of this relation that could be an indication of either market inefficiency or an autocorrelated risk premium.

TABLE IX

Joint Tests of Equality of the Coefficients in the Data Generating Process of Spot and Futures Prices: Error Correction Equations with Included Futures Rates of Other Markets; Period 4:1985–8:1995

		$\alpha_1=\alpha_2$						$\alpha_1\neq\alpha_2$					
	Lags	λ_{LR}	$\chi^2_{(r)}$ at 10%	$\chi^2_{(r)}$ at 5%	T	C	r	λ_{LR}	$\chi^2_{(r)}$ at 10%	$\chi^2_{(r)}$ at 5%	T	C	r
Heating Oil													
$F_{t-2}-S_{t-1}$	8	23.372	30.813	33.924	116	22	22	23.372	29.615	32.670	116	22	21
West Texas I.													
$F_{t-2}-S_{t-1}$	14	37.622	40.256	43.773	110	34	34	37.594	39.087	42.557	110	34	33
Unleaded Gasoline													
$F_{t-2}-S_{t-1}$	14	34.500	40.256	43.773	110	34	34	33.368	39.087	42.557	110	34	33

** = significant at 0.01; * = significant at 0.05; + = significant at 0.10.
Null hypothesis for column $\alpha_1 = \alpha_2$: all coefficients in the spot rate equation are equal to the coefficients in the basis (forward premium) equation; null hypothesis for column $\alpha_1 \neq \alpha_2$: all coefficients in the spot rate equation are equal to the coefficients in the basis (forward premium) equation except for the intercept terms. λ_{LR} = likelihood ratio statistic proposed by Sims (1980), which has an asymptotic χ^2 distribution with degrees of freedom equal to the number of restrictions (r).
Test statistics are based on the error correction equations in Table VIII, augmented by two lagged first differences of the futures prices from each of the other two markets.

An alternative procedure based on stationary data is to test for the equivalence of the data generating processes of the spot and lagged futures prices. These equations are specified to avoid the simultaneity problem that plagues the other tests with stationary variables. When these data generation equations contain only lagged spot and futures prices of the market under consideration, efficiency cannot be rejected for two of the three markets and is weakly rejected for the third. Considering a more general specification of the data generating processes to include lagged futures rates from other markets, the efficiency hypothesis is supported for all three energy futures markets.

Valid tests of market efficiency must take into account the stochastic properties of the variables involved in the test equations. In particular, issues of nonstationarity and endogeneity must be confronted with appropriate procedures that yield consistent estimators with standard statistical distributions. When appropriate procedures are applied to three energy futures markets, in contrast to the existing literature the tests are generally supportive of the efficiency hypothesis.

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