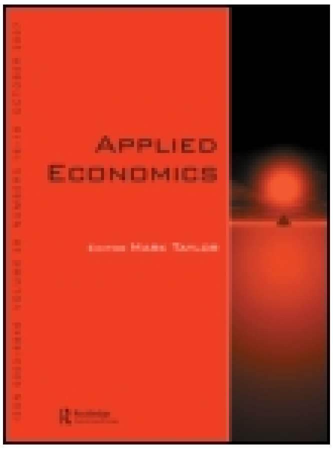

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Cointegration and market efficiency in commodities futures markets

Stacie E. Beck [a]
[a] Department of Economics , University of Delaware , Newark, DE, 19716, USA
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Cointegration and market efficiency in commodities futures markets

STACIE E. BECK

Department of Economics, University of Delaware, Newark, DE 19716 USA

The hypothesis that futures prices are unbiased predictors of spot prices is a joint hypothesis that markets are efficient and risk premia are absent. Rejection of unbiasedness could be caused by the failure of either premise. Here cointegration techniques are used to test market efficiency while permitting the presence of risk premia. Five commodity markets were tested at the eight and twenty-four week horizon. Results showed that all five were sometimes inefficient but no market was inefficient always. Moreover, rejections of the unbiasedness hypothesis were nearly always caused by market inefficiency rather than the presence of risk premia.

I. INTRODUCTION

Market efficiency implies that the current futures price, F_t, of a commodity futures contract expiring in $t+1$ should equal the commodity spot price expected to prevail in $t+1$, i.e. $F_t = E(S_{t+1}/I_t)$. If market participants can use additional information to predict S_{t+1} then they will profit by buying or selling futures contracts whenever $F_t \neq E(S_{t+1}/I_t)$ until equality is re-established. Hence, efficiency implies F_t is the best forecast of S_{t+1}, and that F_t incorporates all relevant information including past spot and futures prices.

To test this hypothesis, various versions of

$$S_{t+1} = c_0 + c_1 F_t + u_{t+1} \tag{1}$$

have been estimated. The expected spot price replaces the actual spot price plus an error by assuming rational expectations. The null hypothesis is $c_0 = 0$ and $c_1 = 1$. Examples of studies on this relationship are Gray and Tomek (1970), Kofi (1973), Leuthold (1974) and Kahl and Tomek (1986).

The hypothesis that the futures price is an unbiased predictor of the spot price is a joint hypothesis that markets are efficient and there are no risk premia. The latter assumes that market participants are risk neutral, an assumption that is neither theoretically defensible nor empirically plausible. For example, if risk-averse producers demand futures contracts to hedge their output, a risk premium would be created that biases futures prices away from expected spot prices, hence $c_0 \neq 0$ (Keynes, 1930). Moreover, Danthine (1978) showed that intertemporal optimization by risk-averse producers implies that the risk premium created by their hedging demand causes both $c_0 \neq 0$ and $c_1 \neq 1$ even in efficient markets.[1] Furthermore, because there is evidence they exist, tests of efficiency should not depend on the absence of risk premia (e.g., Beck, 1993, Kolb, 1992, Park, 1985).

If estimates of Equation 1 reject the null hypothesis of unbiasedness, is it because markets are inefficient or because there is a risk premium? In this paper, I show that the cointegration techniques developed by Engle and Granger (1987) can test for market efficiency while allowing for the possible existence of risk premia. A rejection of the null hypothesis here implies rejection of efficiency, conditional on the assumed form of the risk premium. Moreover, since many commodity price series are nonstationary so hypothesis tests on estimates from Equation 1 are misleading, cointegration techniques are appropriate in any case (Shen and Wang, 1990). Other transformations of Equation 1 can achieve stationarity but they lead to very different results on tests of market efficiency and unbiasedness. Which transformation is correct? Cointegration theory answers this question by making explicit the assumptions behind each specification.

Cointegration techniques have been applied by Hakkio and Rush (1989), Baillie and Bollerslev (1989), Baillie (1989),

[1] Asset equilibrium models such as the capital asset pricing model (CAPM) and the arbitrage pricing theory (APT) also indicate risk premiums exist but these models are not well supported by the data (Baxter, Conine and Tamarkin, 1985; Dusak, 1973; Ehrhardt, Jordan and Walkling, 1987).

and Barnhart and Szakmary (1991) to test unbiasedness in foreign exchange markets and by Chowdhury (1991) to test unbiasedness in the copper, tin, lead and zinc markets at the London Metal Exchange. In this paper, market efficiency, rather than unbiasedness, is tested in five US commodity futures markets. In Section II, two tests of efficiency based on cointegration theory are described. They are then compared to those based on other transformations of Equation 1 appearing elsewhere in the literature. Section III describes the data and reports results on the two efficiency tests for five markets. Tests of unbiasedness are also computed and compared to those from alternative specifications. Section IV presents conclusions.

II. METHODOLOGY

The first test determines whether the spot and futures price series are cointegrated. The series are cointegrated if $u_{t+1} = S_{t+1} - c_0 - c_1 F_t$ is a stationary series, which implies S_{t+1} and F_t do not drift too far from each other although both are non-stationary. Market efficiency implies cointegration because the same factors that determine the future spot price are reflected in the current futures price, so the two should not drift apart.[2] This test permits a risk premium since c_0 can be nonzero and c_1 is not required to equal one. However, it does not rule out residual serial correlation. Residual correlation implies that spot prices rely on past spot prices in addition to current futures prices thus violating market efficiency. The next test investigates this possibility.

Cointegrated series can be rewritten in an error correction representation described in Granger (1986). The transformed series are now stationary so coefficient estimates are asymptotically normally distributed and hypothesis testing is valid.[3] Thus

$$\Delta S_{t+1} = a - \rho u_t + b\Delta F_t + \sum_{i=1}^{m} \beta_i \Delta S_{t+1-i} + \sum_{i=1}^{n} \gamma_i \Delta F_{t-i} \quad (2)$$

where ε_{t+1} is a stationary, possibly serially correlated series with zero mean.

Cointegration implies $\rho > 0$ because spot price changes respond to deviations from the long run equilibrium Equation 1. Market efficiency implies the additional restrictions: $\rho = 1$, $\rho c_1 = b \neq 0$ and $\beta_i = \gamma_i = 0$. The current futures price coefficient b is non-zero because all new information concerning future spot price changes is immediately reflected in a change in the current futures price. The coefficients of lagged spot and futures price changes, β_i and γ_i, are zero because past information is already completely incorporated in the current futures price. To illustrate why ρ must equal one and $\rho c_1 = b$, Equation 2 is rewritten below by substituting $(S_{t+1} - c_0 - c_1 F_t)$ for u_{t+1} and omitting the lagged terms for the sake of clarity:

$$S_{t+1} = (1-\rho)S_t + bF_t + (\rho c_1 - b)F_{t-1} + \rho c_0 + \varepsilon_{t+1} \quad (3)$$

Substituting S_t and solving Equation 3 backwards yields

$$S_{t+1} = bF_t + [b(1-\rho) + (\rho c_1 - b)]\,[F_{t-1} + (1-\rho)F_{t-2} + (1-\rho)^2 F_{t-3} + \ldots] + \rho c_0[1 + (1-\rho) + \ldots] + \varepsilon_{t+1} + (1-\rho)\varepsilon_t + (1-\rho)^2 \varepsilon_{t-1} + \ldots \quad (4)$$

If the restrictions do not hold, then past spot and futures prices contribute information useful for predicting S_{t+1}, therefore all available information is not fully reflected in the current futures price, F_t, so the futures market is inefficient. For the same reason, market efficiency also implies ε_{t+1} is serially uncorrelated. In this test, more restrictions are tested than in the first, therefore it is a stronger test of efficiency. Notice that this test does not impose the assumption $c_0 = 0$ and $c_1 = 1$, thus allowing the presence of a risk premium.

Other stationarity-inducing transformations of Equation 1 turn out to be special cases of Equation 2 which impose additional constraints. For example, stationarity in Equation 1 can be achieved by differencing both S_{t+1} and F_t (e.g., Ma, 1989, p. 396):

$$\Delta S_{t+1} = A + B\Delta F_t \quad (5)$$

In terms of Equation 2, the constraints $\rho = 0$ and $\beta_i = \gamma_i = 0$ are imposed. Equation 5 is misspecified because $\rho > 0$ when S_{t+1} and F_t are non-stationary cointegrated variables, thus tests of the unbiasedness restrictions, $B = 1$ and $A = 0$, are likely to be misleading.

Another variation of Equation 1 regresses spot price changes on the basis (e.g., Fama and French, 1987, p. 63):

$$\Delta S_{t+1} = A + B(F_t - S_t) \quad (6)$$

In terms of Equation 2, the constraints $\rho = \rho c_1 = B \neq 0$ as well as $\beta_i = \gamma_i = 0$ are imposed. This specification is consistent

[2] Although Brenner and Kroner (1992) have argued that arbitrage between storage and futures contracts prevents spot and futures prices from being cointegrated, this is only true of F_t and S_t (the basis), not F_t and S_{t+1} (the risk premium). The basis includes a rate of return to storage at least equal to the market interest rate; research has indicated the interest rate is non-stationary. However, the risk premium reflects only the return to risk-bearing, not the market rate itself, and there is no reason to expect the risk premium to be non-stationary. In fact, tests conducted for this study indicated that the risk premium is stationary (see footnote 7). Authors who have investigated the basis in commodity markets are Baillie and Myers (1991), Bessler and Covey (1991) and Schroeder and Goodwin (1991).

[3] Coefficient estimates from regressions based on Equation 1 are super consistent but not asymptotically normally distributed (Stock, 1987). Coefficient variance estimates from Equation 1 could be adjusted using the method described in West (1988) if the price series followed processes with drifts. There is no theoretical reason for believing that they do, hence this procedure was not used.

with the cointegration theory because $\rho > 0$. However, the constraints imposed on Equation 6 may not hold and they are inconsistent with tests of market efficiency. For example, the first constraint implies $c_1 = 1$ and the second, $c_0 = 0$. Earlier it was pointed out that these conditions do not necessarily hold even if a market is efficient, the absence of a risk premium is also necessary. Tests of $B = 1$ and $A = 0$ actually test the unbiasedness hypothesis but impose $\beta_i = \gamma_i = 0$. Equation 2 is preferred for testing both market efficiency and unbiasedness because all restrictions, including $\beta_i = \gamma_i = 0$, are tested.

III. DATA AND ESTIMATION

Data from seven commodities markets were obtained from the Center for the Study of Futures Markets, Columbia University: live hogs, frozen orange juice concentrate, soybeans, live cattle, cocoa, copper and corn.[4] Since S_{t+1} is the spot price at contract expiration, the frequency of observations is limited to the number of contracts offered per year. Furthermore, the futures price must be chosen at a forecast horizon less than or equal to the observation interval to avoid introducing residual correlation by overlapping observation intervals (Granger and Newbold, 1977, p. 115). Six contracts spaced two months apart are available for hogs, orange juice, soybeans and live cattle.[5] By choosing the futures price eight weeks prior to expiration, all six expiration dates can be pooled. However, for a twenty-four week forecast horizon only two expiration dates can be pooled, hence six contracts yield three subsets of data. Five contracts per year were pooled to obtain cocoa, copper and corn series for the eight week horizon regressions and one subset of data for each were used to estimate the twenty-four week regressions.[6] Because risk premia may be more likely to exist at longer horizons (Beck, 1993) both the eight week and twenty-four week regressions were estimated to examine their effects on tests of efficiency.

The spot price series is actually the futures price on the day of contract expiration. Theoretically, the two prices are the same at expiration since arbitrage will drive them together. In practice, the futures price is often used because spot price data is not generally available for the same grade of commodity delivered at the same time and location as that specified in the futures contract (e.g., Gray and Tomek, 1970; Fama and French, 1987). The use of futures price data avoids biases introduced by inaccurate spot price data.

For cointegration, non-stationary series must be integrated of the same order, i.e., they must have the same number of unit roots. Furthermore, they must be difference stationary rather than trend stationary, i.e., they must have stochastic rather than deterministic trends.[7] The augmented Dickey–Fuller (1981) test could be used to determine the number of unit roots but it is based on the initial assumption of a single unit root. Dickey and Pantula (1987) showed that this test can yield incorrect conclusions if more than one root actually exists. Instead, they suggest sequentially testing for three unit roots, two unit roots, then one unit root. This procedure involves regressing $\Delta^3 x_t$ on $\Delta^2 x_{t-1}$, then on $\Delta^2 x_{t-1}$ and Δx_{t-1}, then on $\Delta^2 x_{t-1}$, Δx_{t-1} and x_{t-1} where Δ^n indicates the degree of differencing. The t-statistic of the most recently added variable is checked for significance against the table reported in Fuller (1976) or Dickey, Bell and Miller (1986). In Table 1, the Dickey–Pantula statistics establish that the price series have, at most, a single unit root.

The augmented Dickey–Fuller test was used to determine whether the series were difference stationary or trend stationary (see Table 2).[8] The results indicated the cattle, copper, cocoa and corn spot and eight week futures price series are difference stationary and the hog and soybean series are trend stationary. Because cointegration techniques apply only to series that are difference stationary, hogs and soybeans were dropped from further testing. The orange juice spot price series narrowly rejected the hypothesis of difference stationarity and the eight-week futures series narrowly failed to reject it so orange juice was retained as a marginal case. Subsets of spot and futures data for the twenty-four week horizon appeared to be difference stationary except one orange juice subset. However, tests on the eight week data provide the best conclusions regarding the stationarity of both eight and twenty-four week price series because they have more observations.

[4] The live hogs, frozen orange juice, soybeans, and live cattle series cover 1966–1987, 1967–1987, 1973–1987, and 1967–1987 respectively. The cocoa, copper, and corn series cover 1966–1986.

[5] The contracts are the February, April, June, August, October and December contracts for hogs and cattle and the January, March, May, July, September and November contracts for soybeans and orange juice.

[6] The contracts are the March, May, July, September and December contracts.

[7] Both price series are in levels rather than transformed to logarithms as is common in the financial futures literatures (e.g. Hakkio and Rush, 1989). Logarithmic transformation when $c_0 = 0$ and $c_1 = 1$ leaves Equation 1 unaffected but this is not true if $c_0 \neq 0$ and $c_1 \neq 1$ as assumed here. Transformations of the dependent variable into a rate of return or a premium, S_{t+1}/F_t or $S_{t+1} - F_t$ (see Dusak, 1973 and Kahl and Tomek, 1986 respectively), are stationary in both the eight and twenty-four week horizons according to tests performed by the author which are available upon request. Thus, cointegration techniques could not be applied. However, this transformation leaves F_t, a non-stationary series, on the right-hand side assuming $c_1 \neq 1$, so hypothesis tests on ordinary least squares estimates would be invalid.

[8] The order of the estimated autoregression for the augmented Dickey–Fuller test was determined using the technique described in Campbell and Perron (1991, p. 155). According to Lagrange multiplier tests, there was no significant residual serial correlation except marginally significant fourth order correlation in the cattle futures series. The augmented Dickey–Fuller test is valid when residuals are uncorrelated although heteroskedastic (Phillips and Perron, 1988, p. 339).

Table 1. *Dickey–Pantula tests*

$$\Delta^3 x_t = \alpha_0 + \alpha_1 \Delta^2 x_{t-1} + \alpha_2 \Delta x_{t-1} + \alpha_3 x_{t-1}$$

Commodity	Price Series	Contracts Pooled	Forecast Horizon	DP_1 a_1	DP_2 a_2	DP_3 a_3
Cattle	Spot	all	0 weeks	−20.06*	−11.80*	−1.68
Orange Juice	Spot	all	0 weeks	−21.57*	−8.24*	−1.72
Hogs	Spot	all	0 weeks	−19.16*	−10.78*	−1.55
Corn	Spot	all	0 weeks	−18.92*	−8.60*	−1.90
Copper	Spot	all	0 weeks	−18.62*	−8.23*	−2.45
Cocoa	Spot	all	0 weeks	−18.46*	−8.27*	−1.74
Soybeans	Spot	all	0 weeks	−16.56*	−10.24*	−2.38
Cattle	Futures	all	8 weeks	−21.45*	−12.20*	−1.68
Orange Juice	Futures	all	8 weeks	−22.49*	−8.65*	−1.49
Hogs	Futures	all	8 weeks	−17.70*	−9.76*	−1.79
Corn	Futures	all	8 weeks	−17.57*	−8.22*	−1.74
Copper	Futures	all	8 weeks	−19.61*	−8.95*	−2.33
Cocoa	Futures	all	8 weeks	−20.92*	−8.20*	−1.91
Soybeans	Futures	all	8 weeks	−15.72*	−9.64*	−2.30
Cattle	Spot	02, 08	0 weeks	−17.90*	−4.43*	−1.54
Cattle	Spot	04, 10	0 weeks	−18.62*	−4.95*	−1.66
Cattle	Spot	06, 12	0 weeks	−14.18*	−5.37*	−1.85
Orange Juice	Spot	01, 07	0 weeks	−9.68*	−6.22*	−1.11
Orange Juice	Spot	03, 09	0 weeks	−9.24*	−6.19*	−1.18
Orange Juice	Spot	05, 11	0 weeks	−11.59*	−6.27*	−1.29
Hogs	Spot	02, 08	0 weeks	−11.70*	−5.82*	−1.64
Hogs	Spot	04, 10	0 weeks	−8.77*	−5.55*	−1.57
Hogs	Spot	06, 12	0 weeks	−10.43*	−6.10*	−1.60
Corn	Spot	03, 09	0 weeks	−11.04*	−4.32*	−1.93
Copper	Spot	03, 09	0 weeks	−11.57*	−4.81*	−2.32
Cocoa	Spot	03, 09	0 weeks	−8.95*	−3.97*	−1.60
Soybeans	Spot	01, 07	0 weeks	−8.35*	−4.41*	−2.12
Soybeans	Spot	03, 09	0 weeks	−8.79*	−4.36*	−1.32
Soybeans	Spot	05, 11	0 weeks	−9.41*	−4.82*	−2.39
Cattle	Futures	02, 08	24 weeks	−11.75*	−4.54*	−1.49
Cattle	Futures	04, 10	24 weeks	−15.73*	−4.30*	−1.60
Cattle	Futures	06, 12	24 weeks	−9.93*	−4.18*	−1.92
Orange Juice	Futures	01, 07	24 weeks	−13.47*	−5.42*	−1.63
Orange Juice	Futures	03, 09	24 weeks	−11.60*	−6.04*	−1.40
Orange Juice	Futures	05, 11	24 weeks	−12.10*	−5.88*	−1.32
Hogs	Futures	02, 08	24 weeks	−10.28*	−5.34*	−1.11
Hogs	Futures	04, 10	24 weeks	−11.71*	−5.66*	−1.41
Hogs	Futures	06, 12	24 weeks	−9.84*	−6.64*	−1.37
Corn	Futures	03, 09	24 weeks	−13.89*	−4.26*	−1.39
Copper	Futures	03, 09	24 weeks	−12.93*	−4.78*	−2.49
Cocoa	Futures	03, 09	24 weeks	−8.68*	−3.66*	−1.59
Soybean	Futures	01, 07	24 weeks	−10.46*	−4.27*	−2.07
Soybean	Futures	03, 09	24 weeks	−9.79*	−5.06*	−1.75
Soybean	Futures	05, 11	24 weeks	−10.15*	−5.06*	−1.94

* The null hypothesis was rejected at the 95% level; the confidence level is according to Dickey, Bell and Miller (1986, p. 17).
Note: DP_i tests the null hypotheses $\alpha_i = 0$ with α_{i+1} constrained to zero.

Tests for cointegration between S_{t+1} and F_t are reported in Table 3 for Equation 1 and with the two variables reversed because the cointegrating technique does not specify which should be the left-hand variable. The augmented Dickey–Fuller statistic (ADF) on the residuals from these two regressions determines whether the residuals are stationary. All eight week and most twenty-four week cointegrating regressions pass this test. The exceptions are the twenty-four week cocoa regression and two twenty-four week cattle regressions. With these exceptions, the results

Table 2. *Test of the order of integration*

$$\Delta x_t = a + Bt + (\rho - 1)x_{t-1} + \sum_{i=1}^{m} c_i \Delta x_{t-1}$$

Commodity	Price Series	Contracts Pooled	Forecast Horizon	ADF[a]	m
Cattle	Spot	all	0 weeks	2.14	2
Orange Juice	Spot	all	0 weeks	6.72*	0
Hogs	Spot	all	0 weeks	10.26*	0
Corn	Spot	all	0 weeks	2.74	0
Copper	Spot	all	0 weeks	5.78	0
Cocoa	Spot	all	0 weeks	2.24	0
Soybeans	Spot	all	0 weeks	6.99*	0
Cattle	Futures	all	8 weeks	2.15	2
Orange Juice	Futures	all	8 weeks	5.60	0
Hogs	Futures	all	8 weeks	7.48*	0
Corn	Futures	all	8 weeks	1.58	0
Copper	Futures	all	8 weeks	6.40	0
Cocoa	Futures	all	8 weeks	2.57	0
Soybeans	Futures	all	8 weeks	7.17*	0
Cattle	Spot	02, 08	0 weeks	3.42	2
Cattle	Spot	04, 10	0 weeks	2.32	1
Cattle	Spot	06, 12	0 weeks	4.61	0
Hogs	Spot	02, 08	0 weeks	5.38	0
Hogs	Spot	04, 10	0 weeks	4.11	0
Hogs	Spot	06, 12	0 weeks	4.83	0
Orange Juice	Spot	01, 07	0 weeks	5.34	0
Orange Juice	Spot	03, 09	0 weeks	4.24	0
Orange Juice	Spot	05, 11	0 weeks	7.98*	0
Corn	Spot	03, 09	0 weeks	0.60	0
Copper	Spot	03, 09	0 weeks	0.37	0
Cocoa	Spot	03, 09	0 weeks	3.42	0
Soybeans	Spot	01, 07	0 weeks	1.69	3
Soybeans	Spot	03, 09	0 weeks	6.38	0
Soybeans	Spot	05, 11	0 weeks	5.91	0
Cattle	Futures	02, 08	24 weeks	2.45	0
Cattle	Futures	04, 10	24 weeks	3.18	2
Cattle	Futures	06, 12	24 weeks	3.62	1
Orange Juice	Futures	01, 07	24 weeks	1.73	3
Orange Juice	Futures	03, 09	24 weeks	4.88	0
Orange Juice	Futures	05, 11	24 weeks	7.19*	0
Hogs	Futures	02, 08	24 weeks	3.39	0
Hogs	Futures	04, 10	24 weeks	5.05	0
Hogs	Futures	06, 12	24 weeks	3.90	0
Corn	Futures	03, 09	24 weeks	0.84	0
Copper	Futures	03, 09	24 weeks	0.19	0
Cocoa	Futures	03, 09	24 weeks	3.53	0
Soybean	Futures	01, 07	24 weeks	3.52	0
Soybean	Futures	03, 09	24 weeks	7.68*	0
Soybean	Futures	05, 11	24 weeks	5.63	0

[a] The augmented Dickey–Fuller statistic tests the joint restriction $(\rho-1)=B=0$.
* The null hypothesis of a unit root was rejected at the 95% confidence level; the confidence level is according to Dickey and Fuller (1981), p. 1063, Table VI.

show the series are cointegrated and market efficiency cannot be rejected.

Estimates of the error correction model Equation 2 appear in Table 4. The model was estimated with zero to six

Table 3. *Cointegration tests*

Commodity	Contracts Pooled	ADF	Number of observations
$S_{t+1} = c_0 + c_1 F_t + u_{t+1}$			
Eight Week Horizon			
Cattle	all	9.77*	124
Orange Juice	all	6.05*	118
Corn	all	8.42*	104
Copper	all	7.35*	105
Cocoa	all	4.37*	105
Twenty-four Week Horizon			
Cattle	02, 08	4.77*	36
Cattle	04, 10	4.37*	37
Cattle	06, 12	4.71*	40
Orange Juice	01, 07	5.26*	37
Orange Juice	03, 09	4.53*	35
Orange Juice	05, 11	5.60*	38
Corn	03, 09	5.62*	36
Copper	03, 09	5.32*	36
Cocoa	03, 09	3.11	36
$F_t = k_0 + k_1 S_{t+1} + \tilde{u}_{t+1}$			
Eight Week Horizon			
Cattle	all	9.64*	124
Orange Juice	all	5.76*	118
Corn	all	8.15*	104
Copper	all	9.11*	105
Cocoa	all	4.29*	105
Twenty-four Week Horizon			
Cattle	02, 08	3.77	36
Cattle	04, 10	3.90	37
Cattle	06, 12	4.45*	40
Orange Juice	01, 07	5.59*	37
Orange Juice	03, 09	4.49*	35
Orange Juice	05, 11	5.06*	38
Corn	03, 09	5.80*	36
Copper	03, 09	4.52	36
Cocoa	03, 09	2.95	36

* The null hypothesis of no cointegration was rejected at the 95% confidence level. Confidence levels are from Phillips and Ouliaris (1990), p. 190.

lags of ΔS_t and ΔF_t; lags with significant coefficients were retained (see Engle and Granger, 1987). The lagged residual term, u_t, was recovered from the cointegrating regression Equation 1 to estimate Equation 2. Residual serial correlation was not detected by Lagrange multiplier statistics which were computed for up to sixth order residual serial correlation. Heteroskedasticity was found in the cattle, copper and cocoa eight week regressions and two orange juice twenty-four week regressions so these t and Wald statistics were computed with heteroskedastic-consistent standard errors (White, 1980).

The coefficient of u_t is significantly different from zero in nearly all cases, which is consistent with the previous result that spot and futures prices are cointegrated. The Wald statistic was computed for the additional restrictions im-

Table 4. *Estimated error correction model*

$$\Delta S_{t+1}=\alpha_0+\alpha_1 u_t+\alpha_2\Delta F_t+\sum_{i=1}^{m}\beta_i\Delta S_{t+1-i}+\sum_{i=1}^{n}\gamma_i\Delta F_{t-i}$$

Commodity	Pooled contracts	Estimated Model	R^2	DW	SER	Wald[a]
Cattle	all	$\Delta S_{t+1}=0.09-0.68u_t+0.70\Delta F_t$ (0.28) (−5.01) (4.70)	0.17	2.03	4.05	5.74
Orange Juice	all	$\Delta S_{t+1}=0.01-0.86u_t+0.86\Delta F_t+0.23\Delta F_{t-1}$ (0.01) (−5.35) (4.77) (2.25)	0.21	1.99	11.89	6.57
Corn	all	$\Delta S_{t+1}=-0.03-1.06u_t+1.22\Delta F_t$ (−0.01) (−7.14) (7.57)	0.38	1.97	26.17	3.04
Copper	all	$\Delta S_{t+1}=-0.14-1.11u_t+0.28\Delta F_t+0.79\Delta S_t+0.34\Delta S_{t-1}-0.33\Delta F_{t-1}$ (−0.19) (−4.65) (1.52) (4.33) (2.09) (−2.67)	0.24	1.95	7.75	22.90*
Cocoa	all	$\Delta S_{t+1}=-2.17-1.06u_t+0.95\Delta F_t+0.21\Delta F_{t-2}$ (−0.08) (−5.16) (4.62) (1.56)	0.31	2.04	262.50	4.27
Twenty-Four Week Horizon						
Cattle	02, 08	$\Delta S_{t+1}=0.33-0.79u_t+0.70\Delta F_t$ (0.39) (−3.11) (2.58)	0.20	2.05	5.23	0.76
Cattle	04, 10	$\Delta S_{t+1}=0.65-0.74u_t+0.32\Delta F_t$ (0.65) (−3.02) (1.15)	0.24	1.92	6.09	5.13
Cattle	06, 12	$\Delta S_{t+1}=0.61-0.56u_t+0.42\Delta F_t$ (0.64) (−2.88) (0.68)	0.18	2.03	5.87	5.57
Orange Juice	01, 07	$\Delta S_{t+1}=0.90-0.45u_t+0.56\Delta F_t-0.32F_{t-2}$ (0.37) (−1.50) (1.96) (−3.26)	0.31	1.87	15.36	15.13*
Orange Juice	03, 09	$\Delta S_{t+1}=0.64-0.63u_t+0.75\Delta F_t-0.30\Delta F_{t-1}$ (0.21) (−1.85) (2.42) (−1.96)	0.24	1.97	17.40	5.49
Orange Juice	05, 11	$\Delta S_{t+1}=0.05-1.02u_t+1.09\Delta F_t-0.23\Delta F_{t-1}$ (0.01) (−4.72) (2.46) (−1.70)	0.34	2.01	24.20	3.74
Corn	03, 09	$\Delta S_{t+1}=0.49-0.88u_t-0.07\Delta F_t+0.94\Delta S_t$ (−0.07) (−2.01) (−0.24) (2.33)	0.17	2.00	38.43	11.86*
Copper	03, 09	$\Delta S_{t+1}=1.19-1.49u_t+1.19\Delta F_t+0.43\Delta S_t$ (0.51) (−4.65) (3.58) (2.21)	0.41	1.97	13.77	5.92
Cocoa	03, 09	$\Delta S_{t+1}=-18.19-0.85u_t+1.64\Delta F_t-0.21\Delta S_{t-3}$ (−0.24) (−2.45) (3.52) (−1.41)	0.38	1.89	407.47	18.96*

Note: Numbers in parentheses are t-statistics where H_0: $\alpha_0=\alpha_1=\alpha_2=\beta_i=\gamma_i=0$. Wald and t-statistics are corrected for heteroskedasticity where necessary.

[a]The joint restrictions are $\alpha_1=1$, $\alpha_2=c_1$ and $\beta_i=\gamma_i=0$. This statistic has a chi-square distribution with p degrees of freedom where $p=2+$the number of lags of ΔF_t and ΔS_{t+1}. Parameter c_1 is the estimated futures price coefficient from the cointegrating regression (Equation 1).

*Significant at the 95% confidence level.

posed by market efficiency, i.e. $\alpha_1=-1$, $\alpha_2=c_1$ and $\beta_i=\gamma_i=0$. The efficient market hypothesis was rejected in the eight week copper regression but not in the eight week cattle, orange juice, corn and cocoa regressions. Tests for autoregressive conditional heteroskedasticity (ARCH) revealed a first order ARCH process in the eight week cattle residuals. Since ARCH processes imply that residual variances are serially correlated, there is a (nonlinear) dependence of spot price changes on previous spot and futures price information, violating market efficiency. In the twenty-four week regressions, market efficiency was rejected in the corn and cocoa regressions and one orange juice regression.

There is at least one regression in every commodity market that rejects a test of market efficiency, however efficiency is not rejected by every regression for any market. Thus, conclusions on market efficiency are mixed. Rejection

Table 5. *A comparison of hypothesis tests using various specifications: Wald statistics*

Commodity	Forecast Horizon	Contracts Pooled	Error Correction Model: Market Efficiency Hypothesis I	Error Correction Model: Unbiasedness Hypothesis II	Levels Regression: Unbiasedness Hypothesis III	Differenced Regression: Unbiasedness Hypothesis IV	Basis Regression: Unbiasedness Hypothesis V
Cattle	8 weeks	all	5.74	6.06*	4.86	45.67*	10.43
Orange Juice	8 weeks	all	6.57	6.60	2.15	66.80*	2.48
Corn	8 weeks	all	3.04	2.63	1.09	24.48*	1.26
Copper	8 weeks	all	22.90*	36.35*	12.39*	74.25*	4.19
Cocoa	8 weeks	all	4.27	3.71	2.50	32.38*	3.11
Cattle	24 weeks	02, 08	0.76	1.17	5.77	24.94*	5.38
Cattle	24 weeks	04, 10	5.13	7.30	4.71	43.64*	4.79
Cattle	24 weeks	06, 12	5.57	6.34*	3.87	19.34*	9.27
Orange Juice	24 weeks	01, 07	15.13*	13.18*	3.06	22.32*	2.30
Orange Juice	24 weeks	03, 09	5.49	4.50	3.66	21.31*	1.05
Orange Juice	24 weeks	05, 11	3.74	4.25	2.03	10.82*	1.04
Corn	24 weeks	03, 09	11.86*	15.58*	4.71	56.89*	9.92*
Copper	24 weeks	03, 09	5.92	7.95	9.36*	26.89*	0.14
Cocoa	24 weeks	03, 09	18.96*	18.32*	2.39	3.61	4.46

* Indicates significance at the 95% level.
Note: The estimated equations and null hypotheses are:

I. $\Delta S_{t+1} = a_0 + a_1 u_t + a_2 \Delta F_t + \sum_{i=1}^{m} \beta_i \Delta S_{t+1-i} + \sum_{i=1}^{m} \gamma_i \Delta F_{t-i}$ and H_0: $a_1 = -1$, $a_2 = c_1$, $\gamma_i = \beta_i = 0$. Parameter c_1 is from Equation 1.

II. $\Delta S_{t+1} = a_0 + a_1 u_t + a_2 \Delta F_t + \sum_{i=1}^{m} \beta_i \Delta S_{t+i-i} + \sum_{i=1}^{m} \gamma_i \Delta F_{t-1}$ and H_0: $a_1 = -1$, $a_2 = 1$, $\gamma_i = \beta_i = 0$.

III. $S_{t+1} = c_0 + c_1 F_t$ and H_0: $c_0 = 0$ and $c_1 = 1$.

IV. $\Delta S_{t+1} = A + B \Delta F_t$ and H_0: $A = 0$, $B = 1$.

V. $\Delta S_{t+1} = A + B(F_t - S_t)$ and H_0: $A = 0$, $B = 1$.

of the hypothesis appears to depend on the forecast horizon, and it is not consistently rejected for any commodity or either forecast horizon.

Table 5 compares Wald statistics on tests of the unbiasedness hypothesis with specifications (1), (2), (5) and (6) described in Section II. Tests of the less restrictive market efficiency hypothesis $\rho c_1 = b$, $\rho = 1$ and $\beta_i = \gamma_i = 0$ are also included from Table 4. Hypotheses tests on Equation 1, although invalid with non-stationary data, are reported here to facilitate comparisons between the procedure described here and previous research.

Table 5 shows that the levels regression (Equation 1) rarely rejects unbiasedness. This is not surprising since, when non-stationarity implied by trends in the price series is ignored, regression estimates will overstate the closeness of their relationship (Plosser and Schwert, 1978). The misspecified Equation 5 nearly always rejects unbiasedness. The conclusions on unbiasedness in the error correction model (Equation 2) (Column II) and the basis regression (Equation 6) (Column V) are usually the same. They differ when the imposed constraints, $\beta_i = \gamma_i = 0$, in the basis regression are violated. This occurs in the eight week copper regression and one twenty-four week orange juice regression (see Table 4). Moreover, two cases where lagged coefficients are significant in Table 4, the eight week orange juice and twenty-four week copper regression, easily pass the unbiasedness test with Equation 6 in Table 5 (Column V) but only marginally pass it with Equation 2 (Column II). Differences in results for the eight week cattle and twenty-four week cocoa regressions cannot be attributed to significant lags. Instead, they may be due to a violation of the constraint $c_1 = 1$ imposed in Equation 6.

A comparison of tests of unbiasedness and market efficiency (Columns I and II) shows only one case in which unbiasedness was rejected when efficiency was not: the eight week cattle regression. This implies that inefficiency, not risk premia, play a dominant role in rejections of unbiasedness.

IV. CONCLUSIONS

The objective of this paper was to use cointegration techniques to test market efficiency without relying on the assumption that risk premia do not exist. The results are conditional on the assumed form of the risk premium. In particular, the risk premium is assumed to be constant or dependent on variables uncorrelated with past spot or futures prices. Violations of these assumptions would cause the efficient markets hypothesis to be rejected even if markets are efficient.

The results indicate that all five markets are sometimes inefficient but no market rejected efficiency all the time. This explains previous research which has produced inconclusive and conflicting results. For example, Gross (1988) and MacDonald and Taylor (1988) found the copper market efficient whereas Chowdhury (1991) did not. Bessler and Covey (1991) and Hudson, Leuthold and Sarassoro (1987) found mixed evidence that the cattle market is efficient.

Comparisons between the error correction model and previous models show that the error correction model is the most general specification; this explains the differences in test results obtained with various models. Hence, results here show instances where the error correction model rejects unbiasedness when other models fail to reject it. Lastly, the results indicate that cases where the error correction model rejects unbiasedness are those where efficiency is also rejected, implying that inefficiency rather than the presence of risk premia causes rejection of unbiasedness in commodities futures prices.

REFERENCES

Baillie, R. (1989) Econometric tests of rationality and market efficiency, *Econometric Reviews*, **8**, 151–86.

Baillie, R. and Bollerslev, T. (1989) Common stochastic trends in a system of exchange rates, *Journal of Finance*, **44**, 167–81.

Baillie, R. and Myers, R. (1991) Bivariate Garch estimation of the optimal commodities futures hedge, *Journal of Applied Econometrics*, **6**, 109–24.

Barnhart, S. and Szakmary, A. (1991) Testing the unbiased forward rate hypothesis: evidence on unit roots, cointegration and stochastic coefficients, *Journal of Financial and Quantitative Analysis*, **26**, 245–67.

Baxter, J., Conine, T. and Tamarkin, M. (1985) On commodity market risk premiums: additional evidence, *Journal of Futures Markets*, **5**, 121–25.

Beck, S. (1993) Tests of the intertemporal hedging model, *Journal of Futures Markets*, **13**, 223–36.

Bessler, D. and Covey, T. (1991) Cointegration: some results on US cattle prices, *Journal of Futures Markets*, **11**, 461–74.

Brenner, R. and Kroner, K. (1992) *Arbitrage and Cointegration*, working paper, University of Arizona, Departments of Finance and Economics.

Campbell, J. and Perron, D. (1991) Pitfalls and opportunities: what macroeconomists should know about unit roots, in *NBER Macroeconomics Annual*, MIT Press, Cambridge, MA.

Chowdhury, A. (1991) Futures market efficiency: evidence from cointegration tests, *Journal of Futures Markets*, **11**, 577–89.

Danthine, J. (1978) Information, futures prices and stabilizing speculation, *Journal of Economic Theory*, **17**, 79–98.

Dickey, D., Bell, W. and Miller, R. (1986) Unit-roots in time series models: tests and implications, *The American Statistician*, **40**, 12–26.

Dickey, D. and Fuller, W. (1981) Likelihood ratio statistics for autoregressive time series with a unit root, *Econometrica*, **49**, 1057–72.

Dickey, D. and Pantula, S. (1987) Determining the order of differencing in autogressive processes, *Journal of Business and Economic Statistics*, **5**, 455–61.

Dusak, K. (1973) Futures trading and investor returns: an investigation of commodity market risk premiums, *Journal of Political Economy*, **81**, 1387–1406.

Ehrhardt, M., Jordan, J. and Walkling, R. (1987) An application of arbitrage pricing theory to futures markets: tests of normal backwardation, *Journal of Futures Markets*, **7**, 21–34.

Engle, R. and Granger, C..(1987) Cointegration and error correction: representation, estimation and testing, *Econometrica*, **55**, 251–76.

Fama, E. and French, K. (1987) Commodity futures prices: some evidence on forecast power, premiums and the theory of storage, *Journal of Business*, **60**, 55–73.

Fuller, W.A. (1976) *Introduction to Statistical Time Series*, John Wiley, New York.

Granger, C.W.J. (1986) Developments in the study of cointegrated economic variables, *Oxford Bulletin of Economics and Statistics*, **48**, 213–28.

Granger, C. and Newbold, P. (1977) *Forecasting Economic Time Series*, Academic Press, New York.

Gray, R. and Tomek, W. (1970) Temporal relationships among prices on commodities futures markets: their allocative and stabilizing role, reprinted in *Selected Writings on Futures Markets, Vol. 2*, ed. A. Peck, (1977), Chicago Board of Trade, Chicago.

Gross, M. (1988) A semi-strong test of the efficiency of the aluminum and copper markets at the LME, *Journal of Futures Markets*, **8**, 67–77.

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.

Hudson, M., Leuthold, R. and Sarassoro, G. (1987) Commodities futures price changes: recent evidence for wheat, soybeans and live cattle, *Journal of Futures Markets*, **7**, 287–301.

Kahl, K. and Tomek, W. (1986) Forward-pricing models for futures markets: some statistical and interpretive issues, *Food Research Institute Studies*, **20**, 71–85.

Keynes, J.M. (1930) *A Treatise on Money, Vol. 2*, MacMillan, London.

Kofi, T. (1973) A framework for comparing the efficiency of futures markets, *American Journal of Agricultural Economics*, **55**, 584–94.

Kolb, R. (1992) Is normal backwardation normal?, *Journal of Futures Markets*, **12**, 75–91.

Leuthold, R. (1974) The price performance in the futures market of a nonstorable commodity: live beef cattle, reprinted in *Selected Writings on Futures Markets, Vol. 2*, ed. A. Peck, (1977), Chicago Board of Trade.

Ma, C. (1989) Forecasting efficiency in energy futures markets, *Journal of Futures Markets*, **9**, 393–419.

MacDonald, R. and Taylor, M. (1988) Testing rational expectations and efficiency in the London Metal Exchange, *Oxford Bulletin of Economics and Statistics*, **50**, 41–52.

Park, H. (1985) Re-examination of normal backwardation hypothesis in futures markets, *Journal of Futures Markets*, **5**, 505–15.

Phillips, P. and Ouliaris, S. (1990) Asymptotic properties of residual based tests for cointegration, *Econometrica*, **58**, 165–93.

Phillips, P. and Perron, P. (1988) Testing for a unit root in time series regression, *Biometrika*, **75**, 335–46.

Plosser, C. and Schwert, G. (1978) Money, income and sunspots: measuring economic relationships and the effects of differencing, *Journal of Monetary Economics*, **4**, 637–60.

Schroeder, T. and Goodwin, B. (1991) Price discovery and cointegration for live hogs, *Journal of Futures Markets*, **11**, 685–96.

Shen, C. and Wang, L. (1990) Examining the validity of a test of futures market efficiency: a comment, *Journal of Futures Markets*, **10**, 195–96.

Stock, J. (1987) Asymptotic properties of least squares estimators of cointegrating vectors, *Econometrica*, **55**, 1035–56.

West, K. (1988) Asymptotic normality, when regressors have unit root, *Econometrica*, **56**, 1397–1417.

White, H. (1980) A heteroskedasticity-consistent covariance matrix estimator and a direct test for heteroskedasticity, *Econometrica*, **48**, 817–38.