INTERDEPENDENCIES BETWEEN AGRICULTURAL COMMODITY FUTURES PRICES ON THE LIFFE

P. J. DAWSON*
B. WHITE

Interdependencies between commodity prices can arise from the impact of changing macroeconomic variables, from complementarities or substitutabilities between commodities, or from common responses by speculators. Malliaris and Urrutia (1996) found significant linkages between rollover prices of six related agricultural commodities on the Chicago Board of Trade. This article examines interdependencies between futures prices for soft commodities traded on the London International Financial Futures Exchange (LIFFE), calculated using Clark indices. Results show that there are no interdependencies between any two prices; price discovery of one contract provides no information about others. © 2002 John Wiley & Sons, Inc. Jrl Fut Mark 22: 269–280, 2002

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*Correspondence author, Department of Agricultural Economics and Food Marketing, The University of Newcastle upon Tyne, Newcastle upon Tyne NE1 7RU, United Kingdom; e-mail: P.J.Dawson@ncl.ac.uk

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- *P. J. Dawson is a Reader in the Department of Agricultural Economics and Food Marketing at the University of Newcastle upon Tyne in the United Kingdom.*
- *B. White is a Senior Lecturer in the Department of Agricultural Economics and Food Marketing at the University of Newcastle upon Tyne in the United Kindom and an Adjunct Senior Lecturer in the Agricultural Resource Economics Group, University of Western Australia.*

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INTRODUCTION

Co-movement between futures prices exists when two or more prices move together in the long run. Its existence implies that the discovery of one price will provide valuable information about others. Therefore, evidence of interdependencies between futures prices is important to traders who often regard its existence as axiomatic. As Pindyck and Rotemberg (1990) noted:

> [Interdependencies between futures prices] . . . would be of little surprise to brokers, traders, and others who deal regularly in the futures and cash markets, many of whom have a common belief that commodity prices tend to move together. Analyses of futures and commodity markets issued by brokerage firms, or that appear on the financial pages of newspapers and magazines, refer to copper or oil or coffee prices rising because commodity prices in general are rising, as though increases in those prices are caused by or have the same causes as increases in wheat, cotton, and gold prices.

Commonality between nominal commodity futures (or cash) prices can arise for four reasons. First, macroeconomic variables, like aggregate demand, inflation, and interest rates, are common determinants of commodity prices. Second, correlations between prices of similar commodities may exist because of substitutabilities or complementarities in supply or demand. Third, correlations between the futures prices of all commodities may be caused by speculators' behavior, and Pindyck and Rotemberg (1990) provided two possible explanations. First, imperfect capital markets impose liquidity constraints on speculators who are long in several commodities so that a fall in one futures price will decrease others by reducing the demand for futures contracts. Second, *herd* behavior arises if a sufficiently large proportion of traders react to news in a similar way: traders are either bullish or bearish on *all* markets without economic reason, but because of market psychology, and this may be compounded by automated trading. Finally, in the context of the EU Common Agricultural Policy, commonality may arise when policies are coordinated: for instance, patterns of wheat and barley EU intervention prices are very similar.

Malliaris and Urrutia (1996), using daily Chicago Board of Trade (CBOT) data for 1981–1991, examined pair-wise linkages between futures prices of six related agricultural commodities—corn, wheat, oats, soybean, soybean meal, and soybean oil—where cross-price elasticities of supply and demand are non-zero. Using residual-based tests, cointegration was found in all cases. Malliaris and Urrutia concluded that significant interdependencies existed between pair-wise futures prices in the long run, thus, ". . . the price discovery function of a commodity futures contract signals valuable information that is relevant to other related commodity futures contracts."

This article complements and extends that of Malliaris and Urrutia. First, interdependencies between all major commodity futures contracts on the London International Financial Futures Exchange (LIFFE) are measured. The commodities are barley, cocoa, coffee, sugar, and wheat.[1] Barley and wheat are expected to be related closely because of their substitutability in the animal-feed market, whereas other pairs are not; intuitively, cross-price elasticities of supply and demand between barley and wheat are non-zero, while those between other commodity pairs tend towards zero.[2] Second, Clark (1973) prices rather than rollover prices are used. Third, interdependency is tested for using Johansen's (1988) procedure for cointegration; here, and in contrast to residual-based tests, causality is not assumed *ex ante*.

EMPIRICAL METHOD

Many economic time series are nonstationary, and OLS regressions between nonstationary data generally are spurious. The presence of unit roots in the autoregressive representation of a time series leads to nonstationarity; such series must be first-differenced to render them stationary or integrated. Where integrated series move together and their linear combination is stationary, the series are cointegrated and the problem of spurious regression does not arise. Cointegration implies the existence of a meaningful long-run equilibrium (Granger, 1988). Since a cointegrating relationship cannot exist between two variables that are integrated of a different order, the order of integration of the variables is tested for first. The subsequent test for cointegration can be interpreted here as a formal test of the long-run equilibrium relationship between any two prices.

The augmented Dickey–Fuller (ADF) test (Dickey & Fuller, 1981; Said & Dickey, 1984), both with and without a deterministic trend, was used to test for unit roots in each price series. The number of lags in the ADF equation was chosen to ensure that serial correlation up to fifth-order is absent using the Breusch–Godfrey statistic (Greene, 2000, p. 541).

[1]The LIFFE also trades a potato contract, but in recent years, volumes traded have been low.

[2]There also has been coordination of EU intervention price policies for wheat and barley. However, this price link through policy may not affect directly spot or futures prices because of the high quality standards required for grain sold into intervention.

When a series is subject to a deterministic trend and an exogenous permanent shock causes a structural break,[3] the ADF test tends to under-reject (Perron, 1989). Accordingly, the Zivot and Andrews (1992) test was used, which tests the null of a unit root against the alternative of a deterministic trend with a structural break, which is estimated endogenously.

If two price series were integrated of the same order, Johansen's (1988) procedure was used to test for cointegration in the vector autoregressive (VAR) model. The procedure is based on maximum-likelihood estimation of the associated vector error-correction model. For two (endogenous) prices P_{1t} and P_{2t}, this is:

$$\Delta \mathbf{x}_t = \boldsymbol{\mu} + \boldsymbol{\Gamma}_1 \Delta \mathbf{x}_{t-1} + \boldsymbol{\Gamma}_2 \Delta \mathbf{x}_{t-2} + \cdots + \boldsymbol{\Gamma}_{k+1} \Delta \mathbf{x}_{t-k+1} + \boldsymbol{\pi} \mathbf{x}_{t-k} + \boldsymbol{\varepsilon}_t \quad (1)$$

where $\mathbf{x}_t = [P_{1t}, P_{2t}]'$, which are $I(1)$, $\Delta \mathbf{x}_t = \mathbf{x}_t - \mathbf{x}_{t-1}$, $\boldsymbol{\mu}$ is a (2×1) vector of parameters, $\boldsymbol{\Gamma}_1, \ldots \boldsymbol{\Gamma}_{k+1}$ and $\boldsymbol{\pi}$ are (2×2) matrices of parameters, and $\boldsymbol{\varepsilon}_t$ is a (2×1) vector of white noise errors. Where $\boldsymbol{\pi}$ is of reduced rank, that is, $r = 1$, it can be decomposed into $\boldsymbol{\pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$ where $\boldsymbol{\alpha} = [\alpha_1, \alpha_2]'$ is the adjustment matrix and $\boldsymbol{\beta} = [\beta_1, \beta_2]'$ is the cointegrating vector. In this case, eq. (1) can be rewritten in full as:

$$\begin{bmatrix} \Delta C_t \\ \Delta M_t \end{bmatrix} = \begin{bmatrix} \mu_1 \\ \mu_2 \end{bmatrix} + \sum_{i=1}^{k-1} \begin{bmatrix} \Gamma_{i,11} & \Gamma_{i,12} \\ \Gamma_{i,21} & \Gamma_{i,22} \end{bmatrix} \begin{bmatrix} \Delta C_{t-i} \\ \Delta M_{t-i} \end{bmatrix} + \begin{bmatrix} \alpha_1 \\ \alpha_2 \end{bmatrix} [\beta_1 \quad \beta_2] \begin{bmatrix} C_{t-k} \\ M_{t-k} \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix} \quad (2)$$

The Granger representation theorem (Engle & Granger, 1987) shows that $\boldsymbol{\beta}'\mathbf{x}_t$ is stationary, implying that $\mathbf{x}_t$ is cointegrated with r distinct cointegrating vectors given by the columns of $\boldsymbol{\beta}$. The Johansen procedure estimates (1), and trace statistics are used to determine the rank of $\boldsymbol{\pi}$.

Harris (1995, p. 96) noted that there are three realistic models (denoted as Models 2–4) implicit in (1). Model 2 is where there are no linear trends in the levels of the $I(1)$ variables and the first-differenced series have a zero mean; here, the intercept is restricted to the cointegration space. Model 3 is where there are linear trends in the levels of the $I(1)$ variables and there is an intercept in the short-run model only. Model 4 is where the model does not account for any long-run linear growth and a linear trend is present in the cointegration vectors.[4] To test

[3]The frost damage to the Brazilian crop in September 1994 may have caused a structural break in the coffee price.

[4]Model 1 accounts for no intercepts or deterministic trends in the cointegrating space, which is unrealistic; Model 5 is appropriate if the data exhibit quadratic trends in level form, which is difficult to justify when the variables are in logarithms, as it implies an unlikely ever-increasing or decreasing growth.

between these models, the Pantula principle (Harris, 1995, p. 97) was used to test the joint hypothesis of both rank and the deterministic components (Johansen, 1992). From estimates of all three models, the testing procedure tests the nulls of no cointegration ($r = 0$) against the alternatives that $r = 1$ using trace statistics in sequence for Model 2–4; if not rejected, the procedure stops because there is no confirmation of cointegration. If the nulls are rejected, then those of $r \leq 1$ are tested against $r = 2$ again in sequence for Models 2–4; the procedure stops the first time the null is not rejected with the conclusion that there is one cointegrating vector.

DATA

Daily nominal close prices of futures contracts for barley, cocoa, coffee, sugar, and wheat on the LIFFE (source: LIFFE, 2000) are used from December 9, 1991 to April 3, 2000, giving 2084 observations. All contracts trade at the same time each day, with the exception of sugar, where the market closes $1^1/_2$ hours later. Some (common) daily prices are missing for all commodities because of public holidays; they are omitted.

A problem with such data is that for any day, a number of contracts are open and the choice arises of which price corresponds to that day. A simple and common solution is to select the price for the nearby contract to construct a rollover price series. In some instances, the rollover series produces a break at or about the time at which a contract expires. For example, Figure 1 shows the rollover price for wheat between May 9 and July 27, 1999; two contracts expired during this period, one on May 21 and the other on July 14, when the rollover price series shows a discontinuity caused by the switch from the July contract to the September contract.[5] The reason for the discontinuity here is that the July contract is based on prices for the stored crop from the 1998 harvest while the September contract reflects expectations of prices for the 1999 harvest. Discontinuity in rollover prices also are found in rollover series created by other standard procedures, such as volume crossover (where the rollover series switches to the contract with the highest traded volume), volume threshold (where the contract closest to expiry is used until its traded volume falls to a specified threshold level), and fixed date before expiry.

Clark (1973) proposed a solution to price selection that overcomes this problem. The Clark index generates a nominal futures

[5]Commodities on the LIFFE are traded for different months and discontinuities occur at different times.

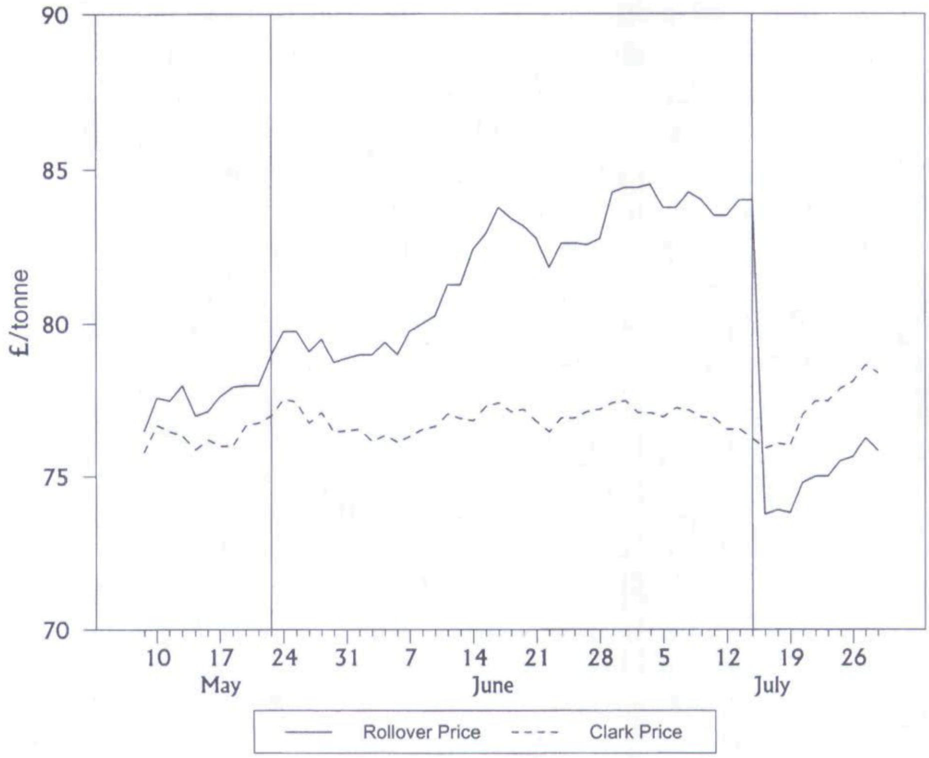

FIGURE 1
Wheat price, 1999.

contract that has an approximately constant time to expiry by weighting futures prices based on the average proportion of open interest for contracts at the range of times to expiry for the whole sample. The weights are:

$$w_\tau = \frac{\sum_t \left(z_{t\tau} / \sum_\tau z_{t\tau} \right)}{n_\tau} \tag{3}$$

where $z_{t\tau}$ is the open interest, t is time in days, τ is days to expiry, and n_τ is the number of contracts for each expiry time. For example, consider wheat, where the longest time to expiry is 420 days and the number of contracts in the sample for each expiry date average 33 (maximum: 53, minimum: 5). The number of contracts decline as expiry time increases. Figure 2 illustrates the Clark index for wheat, which shows a five-period *seasonality* due to the pattern of contract availability. Clark (1973) suggested that the index is smoothed to eliminate local maxima/minima using a simple moving average, also shown in Figure 2.

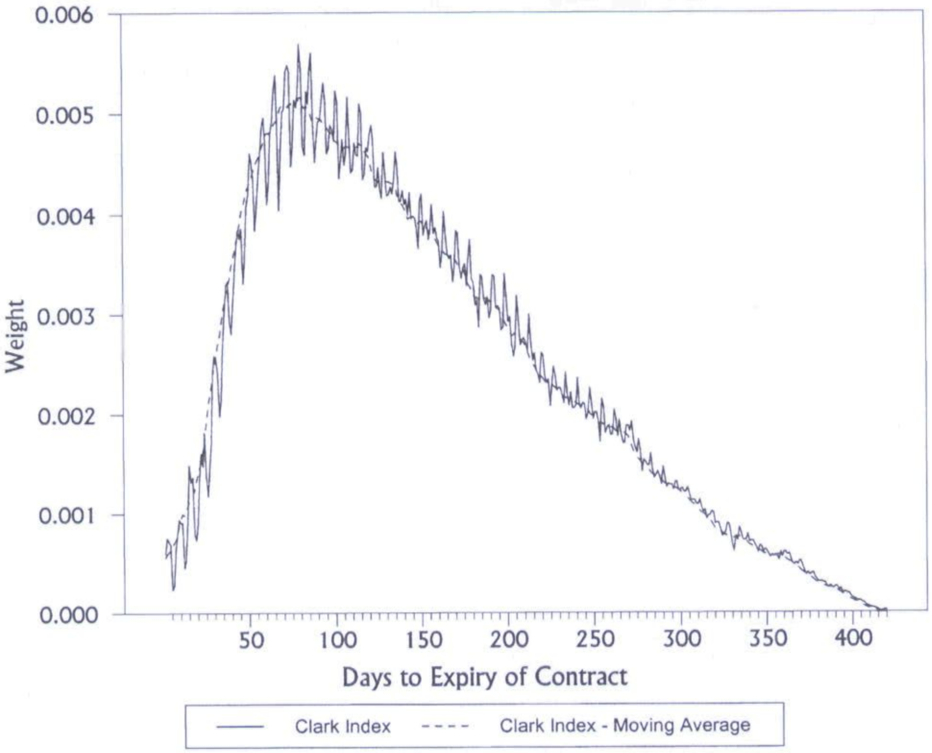

FIGURE 2
Clark index for wheat.

From the moving average of (3), the Clark price is:

$$P_t = \frac{\sum_{T_t} w_\tau P_t^\tau}{\sum_{T_t} w_\tau} \tag{4}$$

where P_t^τ is the futures contract settlement price traded on day t with an expiry time of τ. The denominator normalizes the weights where T_t is the subset of τ traded on a particular day. For the wheat data in Figure 1, the Clark price also is shown; observe its ability to eliminate the discontinuity on July 14, when a contract expires.

The Clark price data are presented in Figure 3 and described in the Appendix. Barley and wheat prices appear interdependent: the barley price ranges from £70–136/tonne, while the corresponding values for wheat are £70–138/tonne. Relationships between other prices appear weaker: cocoa varies between £560–1197/tonne and sugar between \$329–831/tonne. The coffee price is more volatile, ranging from \$689/tonne to a maximum of \$4052/tonne on September 21, 1994, when an apparent structural break occurs caused by severe frost damage to the Brazilian crop.

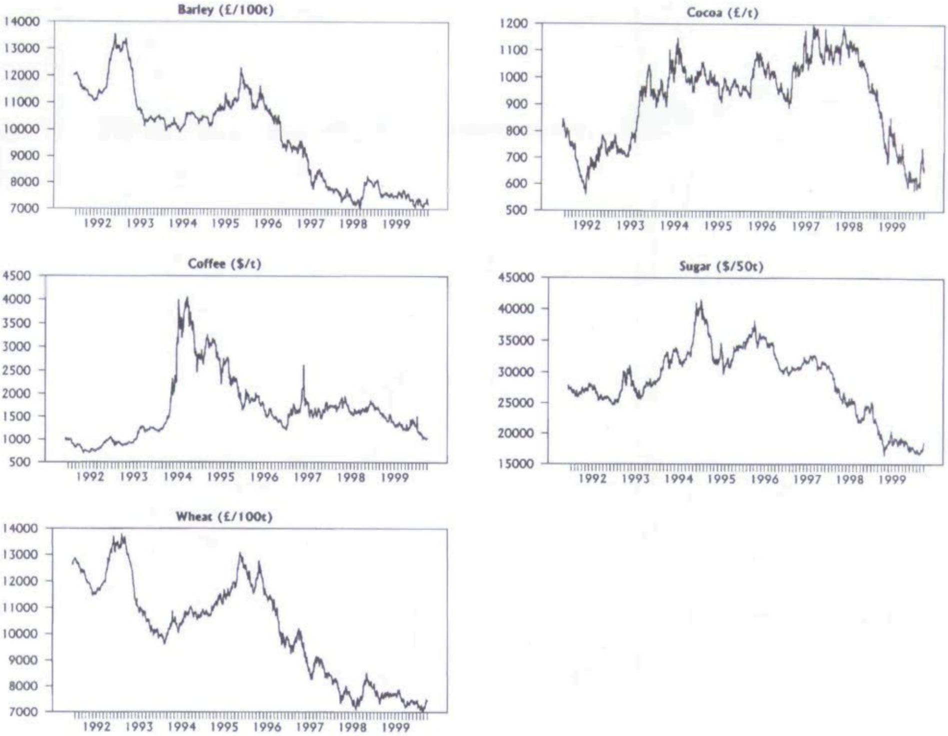

FIGURE 3
Clark price indices.

RESULTS

The time-series properties of the prices in logarithms are examined using the ADF test. The null that each series is $I(1)$, irrespective of whether a trend is included or not, is not rejected in all cases, as shown in Table I. The results of the Zivot and Andrews (Z&A) tests also are presented in Table I, and they substantiate these conclusions; no significant structural break is present in any series.[6] Accordingly, cointegrating relationships between all pair-wise prices are sought.

The first step of the Johansen procedure is to select the order of the VAR (k) for each pair-wise price relationship. The Schwarz Bayesian information criterion (Greene, 2000, p. 717) is used, with maximum $k = 10$, and the results are shown in Table II. Also shown in Table II are the trace statistics for each price relationship for Models 2–4. Since our sample size is large, the 99% confidence interval is used. The null of no cointegration

[6]Interestingly, there is no statistical evidence of a break in the coffee price in September 1994. However, two discrete increases appear to have occurred around this time, which the Zivot and Andrews test may not be able to identify because it tests only for a single break.

TABLE I
Unit Roots Tests (H_0: 1 unit root)

Price Variable	*ADF Test H_1: Deterministic Trend*	*ADF Test H_1: Stationarity*	*Z&A Test H_1: Deterministic Trend with Break*
Barley	−1.80	−0.43	−3.20
Cocoa	−0.74	−1.02	−3.40
Coffee	−1.12	−1.43	−3.22
Sugar	−1.05	−0.15	−2.90
Wheat	−1.72	−0.64	−2.88
Critical Values[a]	−3.41[b]	−2.86[b]	−4.80[b]
	−3.96[c]	−3.43[c]	−5.34[c]

Notes: [a]Critical values for ADF-test (Fuller, 1976, p. 373); [b]95% confidence level; and [c]99% confidence level.

TABLE II
Johansen Cointegration Tests: Trace Statistics

Relationship	*H_0 (H_1)*	*No. of Lags (k)*	*Model 2*	*Model 3*	*Model 4*
Barley–Cocoa	$r = 0$ ($r = 1$)	2	5.671	2.332	5.861
	$r \leq 1$ ($r = 2$)		2.182	0.139	2.152
Barley–Coffee	$r = 0$ ($r = 1$)	2	6.378	3.080	11.791
	$r \leq 1$ ($r = 2$)		1.052	0.009	3.062
Barley–Sugar	$r = 0$ ($r = 1$)	2	10.212	6.344	9.617
	$r \leq 1$ ($r = 2$)		3.200	0.443	3.331
Barley–Wheat	$r = 0$ ($r = 1$)	2	16.221	12.854	18.188
	$r \leq 1$ ($r = 2$)		3.491	0.186	3.066
Cocoa–Coffee	$r = 0$ ($r = 1$)	1	6.163	6.004	8.596
	$r \leq 1$ ($r = 2$)		0.732	0.660	3.244
Cocoa–Sugar	$r = 0$ ($r = 1$)	1	9.764	8.848	14.703
	$r \leq 1$ ($r = 2$)		0.961	0.333	3.833
Cocoa–Wheat	$r = 0$ ($r = 1$)	2	5.416	2.821	5.418
	$r \leq 1$ ($r = 2$)		2.314	0.504	2.309
Coffee–Sugar	$r = 0$ ($r = 1$)	1	4.552	3.752	9.863
	$r \leq 1$ ($r = 2$)		0.810	0.020	3.716
Coffee–Wheat	$r = 0$ ($r = 1$)	2	6.083	3.524	12.279
	$r \leq 1$ ($r = 2$)		0.883	0.015	3.364
Sugar–Wheat	$r = 0$ ($r = 1$)	2	9.796	6.544	9.178
	$r \leq 1$ ($r = 2$)		2.934	0.357	2.704
Critical Values[a]	$r = 0$ ($r = 1$)		19.96[b]	15.41[b]	25.32[b]
			24.60[c]	20.04[c]	30.45[c]
	$r \leq 1$ ($r = 2$)		9.24[b]	3.76[b]	12.25[b]
			12.97[c]	6.65[c]	16.26[c]

Notes: [a]Osterwald-Lenum (1992); [b]95% confidence level; and [c]99% confidence level.

cannot be rejected for each of Models 2–4 and for each price relationship: in each case, the trace statistic is well below its corresponding critical value, even at the 95% level. Thus, no long-run pair-wise relationship exists and agricultural commodity futures prices on the LIFFE are not interdependent; there is no price discovery between commodities.[7]

CONCLUSIONS

Malliaris and Urrutia (1996) examined the linkages between six closely related agricultural futures contracts—corn, wheat, oats, soybean, soybean meal, and soybean oil—on the CBOT. Using residual-based cointegration tests and rollover prices, significant long-run bivariate relationships exist. Thus, futures contracts are interdependent and the price discovery of one provides information about others.

In this study, interdependencies between the agricultural futures contracts of barley, cocoa, coffee, sugar, and wheat are examined on the LIFFE. Barley and wheat are substitutes in demand and supply and are expected to be related; other pair-wise combinations are expected to be unrelated because they are neither complements nor substitutes in either production or consumption. Long-run interdependence between prices are tested for using Johansen's cointegration procedure, which does not assume causality *ex ante*. Clark price indices are used, rather than rollover prices, for daily nominal closing prices of futures contracts for 1991–2000.

Inspection of Clark price indices for barley and wheat suggest that there is some evidence of interdependence; conversely, for other price pairs, there seems little evidence of interdependence. Further, coffee prices appear to have increased markedly because of frost damage to the Brazilian crop in September 1994, and linkages between coffee and other prices may not be expected.

For two prices to be cointegrated, they must be integrated of the same order and unit root tests imply that all prices are $I(1)$. Johansen's tests of cointegration between all pair-wise price combinations imply that no long-run relationships exist and that there is no interdependence between agricultural futures prices on the LIFFE, including between the intuitively related commodities of barley and wheat.

Excepting the barley–wheat relationship, there are two reasons why the results here differ from those of Malliaris and Urrutia (1996). First,

[7]An anonymous referee has suggested that testing for cointegration among a larger group of prices is a reasonable alternative strategy to that adopted here. Using all five prices and $k = 1$, the trace statistic for Model 2 and $r = 0$ is 62.64 (critical value at the 99% confidence interval is 83.93), and there are no long-run relationships between any of the five prices.

agricultural futures contracts for commodities whose cross-price elasticities of supply and demand are expected to be close to zero are examined here, whereas Malliaris and Urrutia examined closely related commodities. Second, futures contracts on a different exchange are examined. Two further reasons might be hypothesized. First, Clark price indices are used instead of rollover prices; however, no cointegrating relationships are found using the latter. Second, the Johansen procedure is a more powerful tool for testing for cointegration than residual-based tests using the errors from single equation models; however, using this latter method, cointegration also is rejected in all cases. For the barley–wheat relationship, the reasons why the results here differ from those of Malliaris and Urrutia (1996) are less clear. A possible reason is that while the contracts are for soft (feed) wheat and feed barley, their patterns of demand are sufficiently different to suggest that they are not close substitutes: wheat has a higher nutritional value in feed milling, but some animals, notably sheep, are restricted to a barley-based diet, and thus, the feed barley price can move independently of the wheat price if demand from sheep producers changes out of step with feed demand from other livestock sectors.

The results here have important implications for brokers, traders, and speculators at the LIFFE. In particular, the prices of agricultural futures contracts are independent, and the discovery of the price of one futures contract does not provide information about other contracts. This implies that traders assess the macroeconomic affects on commodity prices in relation to a commodity in isolation. Pindyck and Rotemberg's (1990) observation that brokers and traders believe that commodity prices move together is not supported by evidence from the LIFFE. Furthermore, there is no firm evidence of *herd* behavior or of liquidity constraints. These findings might be explained by the fact that traders on the LIFFE tend to specialize in a commodity with a relatively small number trading across all commodities.

APPENDIX

Description of Data ($n = 2084$)

Clark Index	*Units*	*Mean*	*Standard Error*	*Minimum*	*Maximum*
Barley	£/100 tonnes	9774	1730	7027	13562
Cocoa	£/tonne	917	158	560	1197
Coffee	$/tonne	1660	713	689	4052
Sugar	$/50 tonnes	28555	5480	16434	41528
Wheat	£/100 tonnes	10144	1876	7004	13793

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