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Price Discovery in Wheat Futures Markets

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ABSTRACT

This paper examines the price discovery function for three U.S. wheat futures markets: the Chicago Board of Trade, Kansas City Board of Trade, and Minneapolis Grain Exchange. The maintained hypothesis is that futures markets search more for information than cash markets to find an equilibrium price, thus greatly improving the price discovery function. The tests reveal the existence of one equilibrium price across the three futures markets in the long run, but no cointegration among prices in the three representative cash markets.

Key Words: error correction model, price discovery, wheat futures.

Price discovery is the process through which markets attempt to reach equilibrium prices (Schreiber and Schwartz). In a static sense, price discovery implies the existence of equilibrium prices. In a dynamic sense, the price discovery process describes how information is produced and transmitted across the markets.

Price discovery is a major function for commodity futures markets. Information on price discovery is essential because these markets are widely used by firms engaged in the production, marketing, and processing of commodities. Production and consumption decisions depend on efficient price signals from the markets. It is generally argued (Leuthold, Junkus, and Cordier, p.4; Peck, p.70) that price discovery in commodity futures markets is more efficient than that in cash markets. The literature, as reviewed in the next section, focuses on whether futures rather than cash markets are the primary source for price discovery. The issue relates to the temporal price relationship between cash and future markets. Specifically, for similar commodities, do futures prices dominantly lead the cash price or vice versa? The paradigms presented here are only concerned with a specific commodity that is traded in a single centralized cash and/or futures market.

Rather than focusing on the cash-futures price relationship, this study investigates the spatial nature and extent of the price discovery process when multiple futures markets exist for a homogenous or closely linked commodity (i.e., a futures-futures price relationship). This issue has not been addressed previously and may provide a completely new perspective and insight for exploring the price discovery function of futures markets. Specifically, the study examines whether prices in the multiple futures market are more likely to search out equilibrium prices (i.e., futures-to-futures price discovery) than prices in the multiple cash markets (i.e., cash-to-cash price discovery) when multiple cash markets for the commodity also exist. Furthermore, the study seeks to determine whether or not there is a

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dominant market that processes and transmits most of the relevant information to other markets when futures prices from different sites converge to equilibrium prices. If such a dominant market exists, what are the important determinants characterizing it? The hypotheses concerning these issues are further developed in the following sections.

The study employed cointegration analysis and an error correction model (ECM) developed by Engle and Granger, and Johansen, to explore the structure of U.S. wheat futures market price discovery. Wheat is a major agricultural commodity in the U.S. and is traded on many cash markets. Wheat futures contracts have been one of the most actively traded commodity futures in the country. The major exchanges trading wheat futures include the Chicago Board of Trade (CBT), the Kansas City Board of Trade (KCBT), and the Minneapolis Grain Exchange (MGE). The three wheat cash markets selected for this study are those that publish daily prices in the *Wall Street Journal.*

Cointegration Based Literature Review

Past studies on the price discovery function of futures markets have concentrated on the temporal price relationship between futures and cash prices. If the current futures price is an unbiased predictor of future cash prices, it can provide direct evidence in favor of price discovery occurring primarily in the futures market. This type of test is also called a *test of futures market efficiency* and is closely associated with the price discovery role. Researchers use cointegration techniques to investigate price discovery issues because cash and futures prices of most commodities are nonstationary. The traditional regression analysis regarding nonstationary prices may yield spurious results. Nonstationary prices can become arbitrarily large or small and there is no tendency for them to revert to their mean level. The unbiased predictor hypothesis immediately implies an equilibrium relationship between cash and futures prices, which may be captured by cointegrating vectors. Cointegration of cash and futures prices is thus necessary for price discovery (in a static sense). Several sources (Quan; Schwartz and Szakmary; Covey and Bessler; Karbuz and Jumah) reported cointegrated cash and futures prices for many commodities. However, the results were mixed for the leading role between the cash and future prices. Other sources (Baillie and Myers; Chowdhury; Bessler and Covey; Schroeder and Goodwin) reported no cointegration between cash and futures prices for many other commodities. Thus, the cointegration test results were not consistent and difficult to explain.

Brenner and Kroner were the first to note the possibility of misspecification in the bivariate analysis of cash and futures prices. They argued that the results of such cointegration tests depend entirely on the time series properties for the cost-of-carry, and that the frequent failure to find cointegrated cash and futures prices for commodities may be due to interest rates with a stochastic trend during the sample period. Theoretically, the bivariate analysis between cash and futures prices fails to account for the price differential (the cost-of-carry). However, if the cost-of-carry is a stationary variable, the cash and futures price still cannot drift apart, and the validity of inference from the bivariate analysis is not imperiled (Brenner and Kroner; Zapata and Fortenbery). In the case of stochastic cost-of-carry, the cointegration test should account for all the random elements of the cost-of-carry, particularly the interest rate. The interest rate is recognized as an important, and most likely nonstationary, part of the cost-of-carry. In this case, cash and futures prices could drift apart when considered by themselves, but they may be cointegrated when the stochastic cost-of-carry is also considered. Unfortunately, this argument was ignored by most studies except Zapata and Fortenbery. Their results supported Brenner and Kroner's argument. Hence, the results reporting no cointegration between cash and future prices in the past were incorrect and the associated conclusions on price discovery in commodity futures markets should be reexamined. In contrast, cointegration-based research on the spatial nature of price discovery in multiple future (or cash) markets does not run the risk of this kind of misspecification, simply because there is no difference in the time dimension involved for several spatial markets.

A review of the literature reveals that little work has been done to explore price discovery across spatially different futures markets for a commodity. However, a few cointegration-based studies were conducted on spatial price relationships across cash commodity markets. We may draw some inference on price discovery function in cash markets from these previous studies. The earlier works suffer some econometric shortcomings, such as ignoring non-stationary properties of the analyzed variables and inappropriate application of price difference modeling (Ardeni; Goodwin and Schroeder). According to Engle and Granger, prices for the same commodity in different spatial markets may be expected to move together, i.e., cointegrate, although they may individually wander extensively. If prices in different spatial markets are fully cointegrated with one price, this cointegration provides direct evidence for perfect price discovery in the long-run. Goodwin and Schroeder found varying degrees of price cointegration in various combinations of regional cattle markets. Jung and Doroodian's results supported the hypothesis of a single long run equilibrium price for four U.S. regional softwood lumber markets. Silvapulle and Jayasuriya demonstrated that the Philippines rice spatial markets were well integrated in the long run. Lutz *et al.* found mixed evidence for one equilibrium price across each pair of several spatially separated maize markets in Benin. These studies suggest that spatially isolated cash markets in some cases may function well in commodity price discovery.

Our study explores the price discovery function when multiple futures markets exist. It also improves the implementation of time series analysis. A multivariate Maximum Likelihood cointegration test was used, rather than the Engle-Granger two-step approach (Engle and Granger) or the bivariate Maximum Likelihood cointegration analysis used by many previous studies on spatial price relationship. Engle-Granger's two-step approach suffers several drawbacks, particularly finite sample estimation bias, as mentioned in Urbain (p.28). As noted by Lutz *et al.*, bivariate analysis can only capture the direct price link between a pair of markets. In the case of an indirect link through an intermediate market, or a chain of markets, more insight may be obtained though a multivariate analysis.

Hypothesis Development

Price discovery may occur in both cash markets and futures markets. The former aims to discover equilibrium current prices while the latter aims to find equilibrium forward prices. For storable commodities, the discovered current price and forward price are expected to be linked by cost-of-carry or approximately the cost of storage. With respect to wheat cash markets, Bessler and Fuller argued for the presence of cointegration between wheat prices in twelve U.S. hinterland markets in Kansas and Texas and the Houston port price. However, they examined prices of the same kind of wheat and the involved spatial markets were limited to a smaller region. In contrast, the price discovery for cash markets in this study was at the national level and across all three major kinds of wheat.

Fundamentally, commodity futures markets are expected to increase the information content of market prices (Peck). There are three reasons for this. First, transaction costs are typically lower in an active futures market than in a cash market, which provides greater incentive to search more for better information. Second, futures markets attract additional speculation and the added speculation is expected to improve the amount of information reflected in the current price. Third, in processing the information, speculators must take into account the responses of all participants to the prices implied by any single piece of information, thus improving the rationality of market prices. Hence, a corollary to this theory is that futures prices should be more cointegrated than the cash prices. This can be tested by comparing the numbers of cointegrating vectors in cointegration tests separately for three futures and cash markets. This is the first hypothesis we tested.

Some special considerations are worth noting in testing the above hypothesis. Three major kinds of wheat are traded on the U.S. cash markets. The Minneapolis (MPLS) market trades spring wheat, Kansas City (KC) market trades No. 2 hard wheat, and St. Louis (SL) market trades No. 2 soft red wheat. The wheat futures contracts traded on these three exchanges differ somewhat in the underlying cash wheat market. CBT futures contracts mainly involve delivery of soft red winter wheat, as well as hard red wheat and some kinds of spring wheat. KCBT contracts usually require hard red winter wheat, and MGE contracts only require No. 2 northern spring wheat. However, the price discovery of both cash and futures markets may be reasonably comparable for two reasons. First, Gray notes spring wheat, hard wheat, and soft red wheat are substitutable over a wide range and are blended with one another. For example, for most of the period since World War II the Chicago and Kansas City prices of cash and futures have been highly correlated because the export market makes little distinction between soft and hard wheat. Thus, the possible equilibrium prices (across cash markets or across futures markets) can be produced by processing similar information about demand and supply for all three kinds of wheat. Second, the underlying assets for the three futures markets are comparable to those for cash markets. Note that the kind of wheat traded in the SL, KC, and MPLS cash markets matches the most deliverable wheat at CBT, KCBT, and MGE, respectively.

The second hypothesis centers on price discovery dynamics of multiple futures markets. Several factors may affect the relative importance of the information role among these three futures exchanges.

One factor affecting the role of information among the three futures exchanges is a different degree of association of the underlying wheat cash varieties with the world market. Hard red winter wheat, mostly traded on KCBT, is exported far more often than all other kinds of wheat and accounts for a large proportion of U.S. wheat production and exports (Hinebaugh). Because exports are the most relevant factor in both the U.S. and world total wheat supply and demand evaluations, the KCBT futures market may be the one that is most closely linked with the fundamentals in wheat.

A second factor affecting the role of information among the three futures exchanges is that the speculation level for futures trading is different. Small traders are usually identified as the "speculative" element in markets. According to Hinebaugh, the long-term annual average for small-trader long positions at CBT has been around 45 percent of the total open interest while small-trader short positions average around 35 percent. At KCBT, small-trader long positions make up little more than 30 percent, and short positions account for less than 10 percent of the total open interest. Small traders at MGE generally have accounted for 25 percent of the total open interest. Though a proper amount of speculation helps price discovery, excessive speculation may cause the market to ignore long-run fundamental information. Based on these figures, one might suspect that the speculation at CBT is too high and the speculation at MGE is too low. Hinebaugh argued that a smaller share of speculative traders at KCBT contributed to the close connection of KCBT prices with the basic information on wheat demand and supply.

Different market size is the third factor affecting the role of information among the three futures exchanges. Market size indicates the market liquidity which is usually positively correlated with the information flow processed for each market. Recently, CBT traded nearly two times as many wheat futures contracts as KCBT. MGE is relatively small. These differences may affect the direction of price information flow across the markets. However, the effect of market size on the information role should be considered together with the speculation level, because only the number of truly informed traders in a market contributes to the price discovery. Though wheat futures prices on both CBT and KCBT are claimed to serve as world price references (Chalmin, p.145), we argue that KCBT is the one most likely to be a primary source in forming equilibrium wheat futures prices, if such an equilibrium price does exist. This is the second hypothesis that can be tested.

Description of the Data

The data for this study covered a three-year period, from January 1, 1993, to December 29, 1995, totaling 781 daily observations for each price time series. The selected period was after the passage of the 1990 farm bill, but before the passage of the 1996 farm bill. This helps guarantee that the cash prices observed are the most market-oriented ones available and that the possible cointegration relationship is not affected by major changes in government policy (Bessler and Peterson). A longer price series would force us to consider a possible structural break in cointegration because the farm bills are changed every four or five years.

The original daily close price information was collected for wheat futures traded on CBT, KCBT, and MGE. A nearby futures price time series was constructed to be used in this study as follows. First, we specified that a nearby futures contract is one with the nearest active trading delivery month to the day of trading (e.g., March 1993 is the nearby contract for February 3, 1993). The reason for using a nearby contract is that it is highly liquid and is the most active. Second, to address the common concern of abnormal volatility close to the expiration of a futures contract, we replaced the last seven observations for each nearby contract price with seven observations from the same dates for the successive nearby futures price. Combining the above adjusted nearby futures prices resulted in a nearby futures price time series[1].

[1] A reviewer raised the concern of possible periodic price jumps in the nearby futures price time series when a nearby contract rolls over to the next nearby contract. The existence of such a price jump could bias the cointegration analysis. We performed a recursive estimation of the cointegration model, as described in Hansen and Juselius (1995), to test the robustness of the reported cointegration analysis results. Using selected numbers of observations (150, 200, etc.) as a base period, the recursive estimation of a cointegration relationship across every data point during the sample period confirmed our results reported here. Particularly, the existence of two cointegrating vectors and the constancy of estimated beta coefficients were confirmed.

Cointegration and Static Price Discovery

In order to test whether cash or futures market price series are cointegrated, it is necessary to establish the integration of each individual series with the order of 1. Two standard procedures were applied. The null hypothesis for both procedures was that a unit root exists. If the test statistics are smaller than the corresponding critical values, the null hypothesis may be rejected. Equation 1 is the augmented Dickey-Fuller (ADF) regression, which may be written as:

$$(1) \qquad \Delta y_t = a_0 + a_1 y_{t-1} + \sum_{i=1}^{p} a_i \Delta y_{t-i} + u_t$$

This test assumes that there is at most one unit root and that the error term is a Gaussian white noise. The test statistics reported here are subject to the τ-test distribution (Dickey and Fuller, 1979, 1981).

The second test procedure was proposed by Phillips-Perron (PP) (Phillips and Perron). This test used a non-parametric correction for the serial corrections, thus allowing for weak dependence and heterogeneity in disturbances. The test was performed using the following regression:

$$(2) \qquad y_t = b_0 + b_1 y_{t-1} + \upsilon_t$$

where υ_t is white noise. The test statistics reported here are subject to the z-test distribution.

Table 1 reports results for I(1) versus I(0) (level prices), and I(2) versus I(1) (first price differences), applying the two tests. The null hypothesis in each test was that each of the cash and futures price series contained a unit root, and it should be rejected if the test statistics are smaller than the critical values. The optimal lags were determined by the AIC+2 rule; the lags were determined by the mini-

Table 1. Results of Unit Root Tests

	Without Trend		With Trend	
Market	ADF[a]	PP[b]	ADF[a]	PP[b]
Level Prices				
SL	−0.49	−2.28	−1.97	−8.01
KC	0.17	0.16	−1.91	−7.50
MPLS	−1.26	−5.71	−1.60	−7.48
CBT	−0.09	−0.41	−1.76	−6.68
KCBT	−0.34	−0.96	−2.24	−9.49
MGE	−0.70	−2.00	−2.33	−2.29
First Difference of Prices				
SL	−15.37	−873.22	−15.49	−872.09
KC	−16.84	−826.43	−16.94	−827.03
MPLS	−6.43	−489.05	−6.45	−488.09
CBT	−17.18	−771.45	−17.28	−772.15
KCBT	−16.05	−760.99	−16.10	−761.37
MGE	−13.54	−699.72	−13.57	−699.40

[a] A unit root test developed by Dickey and Fuller.
[b] A unit root test developed by Phillips and Perron.
Notes: The source of critical values is Davidson and Mackinnon (p. 708). The optimal lags are selected by applying the principle of AIC+2 (Pansula *et al.*, 1994). The critical values of the ADF unit root tests with constant and without trend are −3.43, −2.86 and −2.57, respectively, at 1, 5, and 10 percent level. The critical values of the ADF unit root tests with constant and with trend are −3.96, −3.41 and −3.13, respectively, at 1, 5, and 10 percent level. The critical values of the PP unit root tests with constant and without trend are −20.6, −14.1 and −11.2 at 1, 5, 10 percent level, respectively. The critical values of the PP unit root tests with constant and with trend are −29.4, −21.7, −18.2 at 1, 5, 10 percent level, respectively.

mum AIC (Akaike Information Criterion) plus two. Pansula *et al.* (1994) point out that AIC+2 rule corrects the problems of size distortion of ADF tests. Both tests included with and without trend. The results showed that there was one unit root in each wheat cash and futures price, but no unit root in their first differences, at the five percent significant level.

The confirmation that each series is I(1) allowed us to proceed to the multivariate cointegration test. Johansen and Juselius (1990) and Johansen (1991) developed a maximum likelihood (ML) estimator for a cointegrated system which allows for the presence of a linear trend (hereafter called the *Johansen test*). X_t denotes a vector which includes the k time series. If the elements in X_t are cointegrated, X_t can be expressed by a vector autoregressive model:

$$(3) \qquad X_t = \sum_{i=1}^{k} \Pi_i X_{t-i} + \mu + e_t \qquad (t = 1, \ldots, T);$$

This can be written more often as a reduced form ECM model:

$$(4) \qquad \Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu + e_t \qquad (t = 1, \ldots, T)$$

where $\Pi = \alpha\beta'$ and the rank of Π determines the number of cointegrating vectors. The Johansen test statistic of the null hypothesis is that there are at most r ($0 \leq r \leq k$) cointegrating vectors and thus (n−r) common stochastic trends. To test the order of at most r cointegrating vectors for a $k \times 1$ vector, the trace test statistics is calculated as follows:

$$(5) \qquad Trace = -T \sum_{i=r+1}^{k} \ln(1 - \lambda_i)$$

where T is the number of observations, and λ_i is the n-r smallest squared canonical correlation of X_{t-1} with respect to ΔX_t corrected for lagged differences (also called *eigenvalue*).

All the computations and hypothesis tests in this study were carried out with the special software CATS in RATS by Hansen and Juselius. The Johansen cointegration test was performed on both the three cash prices and the three futures prices. The optimal lags in the Johansen cointegration tests were determined by applying Likelihood Ratio (LR) tests to successive vector autoregressive regressions. We set the maximum lags at 12 and conducted LR tests from largest to smallest lags. The null hypothesis in each sequential LR test was that all the coefficients on the higher k versus $k - 1$ order lag are equal to zero. LR tests were conducted at the one percent level and the relevant critical value of χ^2 with a degree of freedom of 9 was 21.67. The LR tests suggested the optimal lag of $k = 12$ for level VAR in the cash prices, and the optimal lag of $k = 10$ for level VAR (and thus $k = 9$ for

Table 2. Johansen Trace Test Statistics for Wheat Cash Markets[a]

	Without Linear Trend			With Linear Trend		
HO[b]:	T[c]	C[d](5%)	C[d](10%)	T[c]	C[d](5%)	C[d](10%)
$r = 0$	26.43	35.07	32.10	24.79	29.51	26.79
$r \leq 1$	6.58	20.17	17.96	5.13	15.20	13.34
$r \leq 2$	1.37	9.09	7.56	0.09	3.96	2.82

[a] The critical values are from Tables A1 and A3 in Johansen and Juselius (1990).
[b] r is the number of cointegrating vectors.
[c] T is the trace test statistics.
[d] C is the trace test critical value.

ECM) in the futures prices. The larger lag times may be attributed to the difference in the underlying cash wheat prices for these markets.

The trace test results for cash prices are listed in Table 2. Following Johansen (1992), the table should be read from left to right and from top to bottom. The first hypothesis cannot be rejected because the calculated test statistics are smaller than the critical values. Obviously, the null hypothesis of $r = 0$, no cointegration among cash prices, could not be rejected even at a 10 percent significance level, either with or without a linear trend. Also, the results were robust against the alternative lags in the cointegration analysis. No cointegration among cash prices indicates that the price discovery function performed poorly for the U.S. wheat cash markets. Consequently, there is evidence that participants in cash markets have less incentive to search for long-run price information, which may be caused by higher transaction costs, less liquidity, etc. Crain and Lee explained that the cash wheat market is for immediate delivery and thus suppliers and buyers may not have time to respond to price information.

The trace test results for futures prices are listed in Table 3. We find in Table 3 that the hypothesis of $r = 0$ and $r \leq 1$ can be rejected in both cases with or without a linear trend at the five percent level (except the hypothesis of $r \leq 1$ is rejected at 10 percent level in the case of no linear trend). However, $r \leq 2$ cannot be rejected in the case of no linear trend. Thus, we conclude that two cointegrating vectors and no linear trend exist. This implies a single common trend driving the three future prices.

We may further investigate the price information mechanism among the cointegrated futures prices. Further tests showed that constants were not significant in the cointegrating vectors. For the convenience of expression later, we can model the movement of three futures market prices—CBT, KCBT and MGE—in the following ECM (Engle and Granger, 1987):

$$\begin{bmatrix} \Delta CBT_t \\ \Delta KCBT_t \\ \Delta MGE_t \end{bmatrix} = \sum_{i=1}^{9} \Gamma_i \begin{bmatrix} \Delta CBT_{t-i} \\ \Delta KCBT_{t-i} \\ \Delta MGE_{t-i} \end{bmatrix} + \begin{bmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \\ \alpha_{31} & \alpha_{32} \end{bmatrix} \times \begin{bmatrix} \beta_{11} & \beta_{12} & \beta_{13} \\ \beta_{21} & \beta_{22} & \beta_{23} \end{bmatrix} \begin{bmatrix} CBT_{t-1} \\ KCBT_{t-1} \\ MGE_{t-1} \end{bmatrix} + \begin{bmatrix} \epsilon_{1t} \\ \epsilon_{2t} \\ \epsilon_{3t} \end{bmatrix} \tag{6}$$

where Γ_i ($i = 1, \ldots, 9$) is a 3×3 matrix and called the *short-run dynamics*; α_{ij} ($i = 1, 2, 3$, $j = 1, 2$) consists of the loading matrix that is a measure of the average speed of convergence towards the long run equilibrium; β_{ji} ($j = 1, 2$, $i = 1, 2, 3$) consists of coefficients in two cointegrating vectors; Δ is the first difference operator; and ϵ_t is a white noise. The two cointegrating vectors are β_{11}CBT + β_{12}KCBT + β_{13}MGE = 0, and β_{21}CBT + β_{22}KCBT + β_{23}MGE = 0. We reported the unrestricted estimates of α_{ji} and β_{ji} in Table 4 (Panel A). The LaGrange multiplier test and Ljung-Box Q test statistics ensure that the error terms of the ML estimators are not autocorrelated.

We further examined whether one equilib-

Table 3. Johansen Trace Test Statistics for Wheat Futures Markets[a]

	Without Linear Trend			With Linear Trend		
HO[b]:	T[c]	C[d](5%)	C[d](10%)	T[c]	C[d](5%)	C[d](10%)
$r = 0$	48.31	35.07	32.10	47.05	29.51	26.79
$r \leq 1$	18.99	20.17	17.96	18.10	15.20	13.34
$r \leq 2$	1.53	9.09	7.56	0.64	3.96	2.82

[a] The critical values are from Tables A1 and A3 in Johansen and Juselius (1990).
[b] r is the number of cointegrating vectors.
[c] T is the trace test statistics.
[d] C is the trace test critical value.

rium price prevailed in the three markets in the long run. This can be tested by the joint hypothesis (H_1) of $\beta_{11} = -\beta_{12}$, $\beta_{13} = 0$, $\beta_{21} = 0$, $\beta_{22} = -\beta_{23}$. The LR test statistics in Table 4 (Panel A) show that the hypothesis cannot be rejected at a 5 percent level because the calculated statistic (6.66) is smaller than the appropriate critical value 11.14. The acceptance of this hypothesis yields the restricted estimates of the cointegrating vectors CBT − KCBT = 0 and KCBT − MGE = 0, which implies simultaneously equal prices between CBT and KCBT, and KCBT and MGE in the long-run equilibrium.

Table 4. ECM Estimation Results and Hypothesis Testing: β and α

Panel A. Cointegrating vectors β matrix

Unrestricted estimates of β matrix (normalized by CBT)

CBT −1. 003KCBT + 0. 046 MGE = 0
CBT −3. 854KCBT + 3. 279 MGE = 0

H_1: Price Equality ($\beta_{11} = -\beta_{12}$, $\beta_{13} = 0$, $\beta_{21} = 0$, $\beta_{22} = -\beta_{23}$)
LR test (χ^2 (4)) = 6. 66

Restricted cointegrating vectors β^*
CBT − KCBT = 0
KCBT − MGE = 0

Panel B: Loading Matrix α

Unrestricted element estimates of α (based on the unrestricted β normalized by CBT)[a]

α_{11}	−0.086 (0.019)	α_{12}	0.004 (0.006)
α_{21}	−0.043 (0.019)	α_{22}	0.008 (0.006)
α_{31}	−0.041 (0.019)	α_{32}	−0.01 (0.006)

H_2 Weak Exogeneity of *i*th market price, $\alpha_{i1} = \alpha_{i2} = 0$, with no restriction on β matrix
(a) $\alpha_{11} = \alpha_{12} = 0$, $\chi^2(2)$ = 21. 47
(b) $\alpha_{21} = \alpha_{22} = 0$, $\chi^2(2)$ = 6. 86
(c) $\alpha_{31} = \alpha_{32} = 0$, $\chi^2(2)$ = 7. 48

[a] Standard errors are in parenthesis.

Combining all the findings above, there is evidence that the same equilibrium price held on the three futures markets in the long run, thus indicating perfect price discovery in the U.S. wheat futures markets. In contrast, there was no cointegration among prices across the three cash markets, indicating that price discovery for the cash markets functioned poorly.

Price Discovery Dynamic Structure

Our tests also explored price discovery dynamics among the three U.S. wheat futures markets. The Granger causality tests are normally applied to explore the price dynamic transmission mechanism. However, unlike the standard VAR we had to examine two parts, Γ_i ($i = 1, \ldots, 9$) and α_{ij} ($i = 1, 2, 3, j = 1, 2$), to draw inference on the Granger causality. First, we identified the structure of the adjustment coefficients matrix, α_{ij} ($i = 1, 2, 3, j = 1, 2$). The unrestricted estimates of α_{ij} are presented in Table 4 (Panel B). The corresponding t-ratios are listed in the parenthesis. We examined whether or not the i^{th} market price was forced to respond to a temporary deviation from the long-run equilibrium relationships. If $\alpha_{i1} = \alpha_{i2} = 0$ ($i = 1, 2, 3$), we say the i^{th} market price is weakly exogenous, and thus the deviation from the long-run equilibrium relationship does not affect the i^{th} market price formation. The results of hypothesis testing (H_2) reported in Table 4 (Panel B) show that at the five percent significance level, we failed to reject the weak exogeneity of $\alpha_{i1} = \alpha_{i2} = 0$ for $i = 2$ but not for $i = 1, 3$. The appropriate critical value is 7.38 at the five percent level, and only the calculated statistic for $\alpha_{i1} = \alpha_{i2} = 0$ ($i = 2$) is smaller than the critical value. The results suggest that only KCBT price changes did not adjust to deviation from long-run equilibrium prices, or more precisely, that KCBT prices were weakly exogenous for the price changes in CBT and MGE at the five percent level.

We also conducted the similar hypothesis testing for H_2 with the restricted β matrix identified previously (the result is not reported here but available on request). In this case, we still failed to reject weak exogeneity of KCBT, but also did not reject weak exogeneity of MGE at the five percent level. Overall, the weak exogeneity of MGE is not supported as strongly as that of KCBT by the data. Also, the weak exogeneity may happen for different reasons. For KCBT, the weak exogeneity may be an indicator of its role in driving the single common trend in the three U.S. wheat futures markets, and thus its prices do not respond to price changes in other markets in the long-run. For MGE, the possible weak exogeneity may simply imply its ignorance of deviation from equilibrium price relationships in other two markets and thus much of the long-run information because MGE is a relatively small and highly professional market (Hinebaugh). The judgement is further supported by the following forecast error variance decomposition tests (it is not necessary to impose an exogenous structure). The simultaneous imposition of two weak exogeneity restrictions was also rejected. Thus, in the following analysis of short-run dynamics, only the results based on weakly exogenous KCBT prices are reported.

The short-run dynamics, Γ_i, may offer new insights for short-run price information transmission. We ran an ML estimation based on cointegrating vectors restricted with long-run price equality and a weak exogenous KCBT price. To save space, we only reported significant coefficient estimates entering the three price changes equations in Table 5. We found

Table 5. ECM Estimation Results of Short-run Dynamics Γi with Restricted β^* and α^{*a}

Exogenous Variables	Dependent Variables ΔCBT_t	$\Delta KCBT_t$	ΔMGE_t
ΔKC_{t-1}	-0.140^b (0.069)	—	—
ΔMGE_{t-1}	—	0.142 (0.063)	0.149 (0.054)
ΔCBT_{t-3}	-0.145 (0.064)	—	-0.153 (0.065)
ΔMGE_{t-3}	0.145 (0.053)	0.147 (0.054)	—
ΔCBT_{t-4}	0.150 (0.064)	0.168 (0.066)	0.194 (0.065)
$\Delta KCBT_{t-4}$	-0.169 (0.069)	-0.208 (0.070)	-0.169 (0.070)
$\Delta KCBT_{t-5}$	-0.178 (0.069)	—	—
ΔCBT_{t-9}	—	0.158 (0.066)	0.154 (0.066)
$\Delta KCBT_{t-9}$	—	-0.161 (0.070)	-0.160 (0.070)

[a] See Equation (6). The β^* matrix is imposed with price equality restrictions, and the α^* matrix is imposed with weak exogeneity of KCBT price change.
[b] Only the coefficient estimates significant at 5 percent are reported here. The corresponding standard errors are indicated in parenthesis.

that bi-directional price change transmission existed between CBT, KCBT, and MGE, i.e., price changes in each of these three markets can Granger-cause price changes in another market in the short-run. This is not surprising because each of the three markets regularly trades a different kind of wheat, and information that is specifically related to a particular kind of wheat in the short run may originate from each market and then be transmitted to other markets. Note that the price change in KCBT always contributes to the reduction of price fluctuation in the other two markets. In contrast, the price change in MGE always contributes to the increase of price fluctuation in the other two markets. We also found that the influence of the MGE price change is limited to a shorter time horizon, as short as three days, while KCBT and CBT play an important role in price formation for a longer time horizon. These findings were consistent with the above projected hypothesis

Table 6. Forecast Error Variance Decomposition, When Ordered as KCBT, MGE, and CBT, Number of Lags Equals 10

Days	KCBT	CBT	MGE
Forecast variance of KCBT (percent) explained by ---- shock to prices at selected exchanges -----			
1	100.00	0.00	0.00
5	97.68	0.75	1.57
10	95.56	0.58	3.85
20	93.48	1.42	5.10
30	90.29	4.06	5.65
40	87.05	6.85	6.09
50	84.23	9.28	6.49
Forecast variance of CBT (percent) explained by ---- shock to prices at selected exchanges -----			
1	67.58	31.15	1.27
5	68.28	28.09	3.63
10	66.19	28.95	4.86
20	74.53	20.59	4.88
30	80.77	14.20	5.03
40	82.45	12.29	5.31
50	81.96	12.38	5.67
Forecast variance of MGE (percent) explained by ---- shock to prices at selected exchanges -----			
1	52.16	0.00	47.84
5	55.58	0.92	43.50
10	57.53	0.47	41.99
20	60.54	3.23	36.23
30	62.69	6.73	30.57
40	63.64	9.80	26.56
50	64.04	12.18	23.78

Table 7. Forecast Error Variance Decomposition, When Ordered as CBT, KCBT, and MGE, Number of Lags Equals 10

Days	KCBT	CBT	MGE
Forecast variance of KCBT (percent) explained by ---- shock to prices at selected exchanges -----			
1	32.42	67.58	0.00
5	35.86	62.69	1.30
10	33.45	62.39	4.16
20	35.46	58.74	5.80
30	41.27	51.55	7.18
40	45.99	45.65	8.36
50	49.47	41.16	9.37
Forecast variance of CBT (percent) explained by ---- shock to prices at selected exchanges -----			
1	0.00	100.00	0.00
5	0.14	98.95	0.91
10	0.12	98.49	1.39
20	4.09	93.95	1.95
30	14.04	82.69	3.28
40	23.95	71.29	4.76
50	31.78	62.06	6.16
Forecast variance of MGE (percent) explained by ---- shock to prices at selected exchanges -----			
1	8.93	45.10	45.97
5	14.58	41.81	43.61
10	16.93	41.08	41.99
20	23.20	38.45	38.35
30	31.46	34.22	33.60
40	37.79	31.47	30.74
50	42.36	29.18	28.46

based on trading microstructure (e.g., market size, speculation share) in these markets.

Forecast error variance decomposition in a level vector autoregression (VAR) model also can be employed as a supplementary test to explore the price dynamics. These tests may give a clearer picture of the force that drives the single common trend in U.S. wheat futures markets. The level VAR does not impose the long-run cointegration relationship which was identified in the previous section, but it does allow for these restrictions to be satisfied asymptotically (Otto and Voss). Bessler mentioned that ordering of the vector should follow the decreasing exogeneity. The shocks are Choleski factored ones.

Forecast error variance decomposition results with the selected lags of 10 are reported in Table 6. The ordering was KCBT, MGE, and CBT, arranged by weak exogeneity test results. KCBT seemed exogenous for CBT and MGE, because about 85 percent of the variance errors for the KCBT prices can be explained by its own shock even after 50 days. MGE's variance error may be largely explained by KCBT, and partly by itself, but not by CBT. The variance error of the CBT price can be explained mostly by the CBT price shock, but little by the MGE price shock. This is particularly true for longer horizons. To ensure that the above results were not sensitive to the choice of how the variables were ordered, Table 7 presents the forecast error variance decomposition results for the alternative ordering of CBT, KCBT, and MGE, according to the total trading volume. The results in Table 7 support our argument derived from Table 6 that the KCBT was more exogenous than either CBT or MGE, and that the KCBT price shock significantly explains the price movements in the other two markets in the long run. Hence, allowing for both ECM Granger causality tests and forecast error variance decomposition results, we conclude that KCBT leads the single price common trend in U.S. wheat future markets.

Conclusions

In this study, we investigated price discovery of U.S. wheat futures and cash markets separately. We found evidence for the argument that futures markets have improved the price discovery function. The three U.S. wheat futures markets were driven by an equilibrium price in the long-run, but no equilibrium relationship of prices across wheat cash markets exists. Results of this study showed that the futures markets provided informed prices that cannot be embodied in the cash markets. Work was also done to examine the dynamic-price discovery mechanism in wheat futures markets. The prices of KCBT were found to drive the price changes in both CBT and MGE in the long run. In the short run, KCBT and CBT contributed more to the price information transmission for a longer time while MGE was limited to a shorter time horizon. These findings were explainable by the market microstructure of the three futures markets, including the role of underlying cash wheat, market size, and speculation level.

The findings of this study may improve our understanding of relative pricing efficiency on futures markets for storable agricultural commodities. This additional knowledge of relative information roles among multiple futures markets and the associated market microstructure determinants helps traders better track forward price signals and improve their decisions for futures trading. The study suggests that traders should pay more attention to the wheat futures price movements from the Kansas City Board of Trade, even though their underlying assets may be various kinds of wheat. The findings of this study also question the need of having three different futures markets for wheat. In the long run, only the KCBT market significantly affects the wheat futures price discovery, suggesting some redundancy in having three wheat futures markets.

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