# Price Discovery and Cointegration for Live Hogs

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## INTRODUCTION

An often-cited, major role of futures markets is their contribution to price discovery [Working (1960, 1970); Garbade and Silber (1983)]. Price discovery in futures markets refers to the use of futures prices in determining cash market prices. That is, the futures price may serve as the market's expectation of a subsequent delivery period cash price [Gardner (1976)]. The share of price discovery originating in the futures market and the degree to which futures price changes are reflected in the cash market and vice versa have important implications for hedgers and arbitragers who use these markets. The objectives of this article are to determine the short-term price leadership roles and the long-term efficiency of the cash to futures price relationship for live hogs.

The theory of intertemporal relationships between cash and futures prices of continuously storable commodities is well-developed [Working (1948, 1949); Brennan (1958)]. The concepts associated with intertemporal cash and futures price relationships for nonstorable commodities (e.g., live hogs) is less developed. Empirical evidence suggests cash and nearby futures prices in livestock markets are strongly related [Purcell et al. (1979); Leuthold (1979); Naik and Leuthold (1988); Oellermann and Farris (1985); Oellermann et al. (1989); Koontz, et al. (1990)]. Kendall (1981) shows that positive correlations can arise among intertemporal prices of a nonstorable commodity due to the interdependencies of current and future supplies. Thus, though perhaps not temporally linked as strongly as storable commodities, nonstorable commodities do possess qualities which link prices across time.

In this article the dynamic regression methodology developed by Garbade and Silber (1983) is utilized to examine the short-run price discovery role of the live hog cash and futures markets. To examine the longer-run stability of the live hog futures to cash market prices relationship, the relatively new concept of cointegration [Granger (1986); Engle and Granger (1987); Engle and Yoo (1987)] is used.

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## DYNAMIC PRICE DISCOVERY

The Garbade and Silber (1983) model used to investigate short-run price discovery is designed for continuously storable commodities. Because live hogs are essentially nonstorable, adjustments must be made to remove storage costs from the model. The Garbade and Silber (1983) model without the storage cost adjustment is employed for nonstorables by Oellermann et al. (1989) in investigating feeder cattle price discovery. The model (with storage costs removed) can be specified as follows:

$$\begin{bmatrix} C_t \\ F_t \end{bmatrix} = \begin{bmatrix} \alpha_c \\ \alpha_f \end{bmatrix} + \begin{bmatrix} 1 - \beta_c & \beta_c \\ \beta_f & 1 - \beta_f \end{bmatrix} \begin{bmatrix} C_{t-1} \\ F_{t-1} \end{bmatrix} + \begin{bmatrix} e_t \\ e_t \end{bmatrix} \tag{1}$$

where $t$ refers to day, and $C$ and $F$ are the logarithms of the cash and futures prices, respectively. The coefficients $\beta_c$ and $\beta_f$ reflect the impact of the previous day's price in one market on the other market's price. In this light it is expected that $\beta_c \geq 0$ and $\beta_f \geq 0$. The constant terms $\alpha_c$ and $\alpha_f$ reflect any trends in the price series.

The ratio $\beta_c/(\beta_c + \beta_f)$ provides an indication of the level of price discovery occurring in each market. If the ratio is equal to one, implying that $\beta_f = 0$, then the cash price follows the futures price, and price discovery originates in the futures market. In this case, the cash market is referred to as a "pure satellite" of the futures market [Garbade and Silber (1983)]. If $\beta_c = 0$, the ratio equals zero, which implies that price discovery originates in the cash market and the futures price adjusts toward the cash price. Values of the ratio between zero and one imply mutual adjustment and feedback of the two markets' prices to each other.

Equation (1) can be rearranged into an easily estimated structure as follows:

$$\begin{bmatrix} C_t - C_{t-1} \\ F_t - F_{t-1} \end{bmatrix} = \begin{bmatrix} \alpha_c \\ \alpha_f \end{bmatrix} + \begin{bmatrix} \beta_c \\ -\beta_f \end{bmatrix} [F_{t-1} - C_{t-1}] + \begin{bmatrix} e_{tc} \\ e_{tf} \end{bmatrix} \tag{2}$$

This equation can be estimated using ordinary least squares regression.

Equation (1) can also be used to examine the rate of convergence or persistence of the cash-futures price difference (basis) from one day to the next. Solving eq. (1) for today's basis as a function of the previous day's basis yields

$$F_t - C_t = \alpha + \delta(F_{t-1} - C_{t-1}) + e_t \tag{3}$$

where $\alpha = \alpha_f - \alpha_c$, $\delta = 1 - \beta_f - \beta_c$, and $e_t = e_{tc} - e_{tf}$. If $\delta$ is near 1, this would suggest that the basis is fairly stable from day to day and that the cash and futures prices do not tend to converge rapidly. If $\delta$ is small, relatively little of the previous day's basis is present in the current day's basis. A large value for $\delta$ would suggest that relatively large cash to futures price differences can occur before arbitrage is undertaken to bring about convergence.

## COINTEGRATION OF CASH AND FUTURES MARKETS

The concept of cointegration posits that two efficient markets for the same asset whose prices are each nonstationary by themselves should have an equilibrium relationship which is stationary. Although significant short-run deviations may be observed, economic forces should prohibit persistent long-run deviations from equilibrium. Alternatively, prices from two efficient markets for different assets cannot be cointegrated (Granger, 1986). That is, an asset's price change should be unpredictable assuming the price incorporates all available information.

Analysis of market integration often relies on the premise that price changes in one market are reflected by equilibrating changes in the other market. Such a relation between the cash and futures markets can be written as

$$C_t - a - bF_t = u_t \tag{4}$$

where $C$ and $F$ are as defined previously, $a$ and $b$ are parameters, and $u$ is the proportional deviation from price parity. If $C$ and $F$ are each nonstationary (and require a single difference to make them stationary), but their linear combination ($u$) is stationary, the two markets' prices are cointegrated of order (1, 1) with a cointegration parameter of $b$.[1] This suggests a natural procedure for evaluating cash and futures price linkages. If the price series are cointegrated, assuming they represent the same asset, then these markets are operating efficiently [Hakkio and Rush (1989)]. If the price series are not cointegrated, then either they are not operating efficiently relative to each other (at least one market is inefficient) or they represent different assets.

Perfectly integrated markets, where price changes in one market are fully reflected by proportionally equal price changes in another market, require that the cointegrating regression slope parameter ($b$) have a value of one. However, because $C$ and $F$ are nonstationary, conventional statistical procedures cannot be used to provide reliable hypothesis tests regarding the value of $b$. In particular, parameter estimates of the cointegrating regression are consistent but the estimated standard errors are not consistent, thereby making explicit hypothesis testing of the cointegration parameter impossible [see Elam and Dixon, (1988); Engle and Granger (1987) for discussions regarding this issue].

Two steps are used to evaluate the cointegrating properties of a concurrent pair of nonstationary economic time series. The first stage involves estimating the cointegrating regression using OLS, which is then used to calculate estimates of the residual errors, $\hat{u}_t$, where

$$\hat{u}_t = C_t - \hat{a} - \hat{b}F_t \tag{5}$$

Upon obtaining estimates of the first stage residual errors, Engle and Granger (1987) propose seven tests for cointegration. Each test has as its null hypothesis the case of no cointegration. Rejections of the tests lend support for cointegration among the markets. The test for cointegration recommended by Engle and Granger (1987) utilizes a Dickey-Fuller (1979, 1981) type regression to determine whether the autoregressive parameter for the estimated residuals from the cointegrating regression is significantly different from one. If there is a unit root, then the two series are not cointegrated. The augmented Dickey-Fuller test depends on estimates of

$$\Delta\hat{u}_t = -\phi\hat{u}_{t-1} + \theta_1\Delta\hat{u}_{t-1} + \ldots + \theta_p\Delta\hat{u}_{t-p} + \varepsilon_t \tag{6}$$

where $\hat{u}_t$ is the first stage estimate of the residual from the cointegrating regression and $\Delta$ implies the first difference. The lagged differences are included to ensure that the second stage residuals of the augmented Dickey-Fuller regression, $\varepsilon_t$, are serially uncorrelated. A test statistic is constructed from the ratio of the estimated $\phi$ to its standard error (a "$t$-ratio"). The null hypothesis of no cointegration is re-

[1]In cointegration notation, two series are said to be cointegrated of order $(d, b)$ if the individual series are integrated of order $(d)$ and their linear combination is integrated of order $(d-b)$ [Engle and Granger (1987)].

jected for values of $\phi$ which are significantly different from zero. Engle and Yoo (1987) report critical values for the augmented Dickey-Fuller cointegration test statistic for sample sizes of 50, 100, and 200 observations. In the ensuing analysis, the lag order ($p$) is selected by the minimum value of Akaike's (1969) final prediction error (FPE).

## RESULTS AND DISCUSSION

The cash price data used in this study are daily slaughter hog prices from the Omaha market (midpoint of the daily range).[2] The Omaha market is selected as the cash price location because it is fairly liquid and a par delivery point for the Chicago Mercantile Exchange's (CME) live hog futures contract. Daily closing futures prices from the CME's live hog contract are used also. The data period selected for this analysis includes the years 1975 through 1989. Critical values for the cointegration test statistics are not standard and are estimated using Monte Carlo techniques by Engle and Yoo (1987) for data sets containing as many as 200 observations. To comply with the number of observations corresponding to the critical values of the test statistic, the empirical analyses are conducted on each of the 15 years separately. The cointegration tests use the first 200 daily price observations of each year (thereby not including roughly the last 50 trading days). The dynamic Garbade-Silber price discovery tests are conducted using all of the daily price data for each year.

The analysis is performed using the nearby futures contract prices. The nearby contract period is defined as the 16th day of the previous contract month through the 15th day of the contract expiration month. For the Garbade-Silber analysis a continuous price series is created by defining the change in the nearby futures price up through the 15th day of the expiration month and switching to the next contract the following trading day. All price differences are calculated using the same contract month for each price.

### Price Discovery Results

The parameter estimates from the Garbade-Silber model presented in eq. (2) are reported in Table I. Over all 15 years the estimate of $\beta_c$ is positive and significantly different from zero at the 5% level, providing strong evidence that the futures price leads the cash price in reflecting new information. The magnitudes of $\beta_c$ vary from year to year, although no general trend is detected over the time period.

Estimates of $\beta_f$ are generally not significantly different from zero. Thus, the live hog futures price appears to discover price independently of the Omaha cash hog price with little feedback from the cash price. Once again, no definitive trend in the degree of cash price feedback is detected in the estimates of $\beta_f$. This result contrasts with that of Oellermann et al. (1989) that the feeder cattle futures market discovers more new information before the cash market during the 1979–1982 period, but this declines during the 1983–1986 period. This also contrasts with the results offered by Koontz et al. (1990) that the live cattle futures price discovery role has increased in recent years.

With a few exceptions, the ratio of $\beta_c/(\beta_c + \beta_f)$ generally has a value of at least 0.60 or greater. Thus, at least 60% of the market information is generally discovered

[2]Cash prices for different weight ranges of hogs over the period are used because of changing price reporting practices by the Agricultural Marketing Service (AMS) and to reflect changes in weights of the largest volume of hogs marketed. The weight ranges used are as follows: 1975–1978 200-220#; 1979–1984 200-230#; 1985–1987 210-240#; and 1988–1989 230-240#.

first in the live hog futures market and then transferred to the cash market. Garbade and Silber's (1983) estimates of this ratio range from 0.75 to 0.85 for corn, wheat, and orange juice, and are 0.54 for oats and copper. Oellermann et al. (1989) find estimates of 0.76 to 1.00 for feeder cattle. During five of the 15 years (1975, 1978, 1980, 1981, and 1983) the Omaha cash hog price appeared to be discovering the majority of new information more rapidly than the live hog futures market (i.e., the ratio was less than 0.50). In four of these 5 years (1978 perhaps as the exception) cash hog prices exhibit relatively large swings during the year. For example, in 1975 the Omaha cash hog prices range from \$37/cwt to nearly \$65/cwt, a greater than 75% change in price. Similarly, during 1980 and 1983 the high to low range of hog cash prices exceed \$22/cwt.[3] Thus, during these years the new information entering the market may well originate in the cash market as the futures market apparently does not anticipate the magnitude of price changes as rapidly as they occurred in the cash market.

The value of $\delta$ remains fairly stable over the 15 years ranging from 0.72 to 1.00 (Table I). These estimates are similar to those of Garbade and Silber (1983), which range from 0.84 to 0.97 for wheat, corn, oats, orange juice, and copper, and to those of Oellermann et al. (1989) of 0.95 to 0.97 for feeder cattle. They contrast somewhat with Garbade and Silber's (1983) estimates for gold and silver of 0.59 and 0.56, respectively.

One can infer from the estimates of $\delta$ that the previous day's live hog cash to futures price basis persists and is similar to the current day's basis. Thus, in the very short run, from day to day, little basis fluctuations occur and the cash and futures prices do not necessarily converge rapidly. This suggests that there would generally be little gain for a hedger to speculate on (day to day) basis on the nearby contract. Likewise, it suggests that very short-term basis risk for hogs in Omaha, as measured by daily fluctuations in basis, is not large on average. However, this does not necessarily imply that longer-term basis risk is small. This issue is addressed below.

### Cointegration Results

Prior to conducting the cointegration tests, one must confirm that the individual cash and futures price series are nonstationary. This is verified using the Dickey-Fuller (1979, 1981) unit root test. During all 15 years the hypothesis of nonstationarity for both the cash and futures price levels is not rejected. Dickey-Fuller (1979, 1981) unit root tests on first differences of the cash and futures price series result in strong rejections of the presence of a unit root across each of the 15 years. This provides evidence that the first-differenced cash and futures price series are stationary.

The results of the cointegration tests between live hog cash and futures prices are reported in Table II. Because the one-way causality from the cash to the futures market is frequent, though not perpetual, the cointegrating regression [eq. (4)] is estimated twice, once with the case price as the dependent variable and the futures price as the independent variable and a second time with the dependence reversed. As shown in Table II, the results of the two orderings are qualitatively identical.

[3]This is not meant to imply that during all periods of large price moves in the hog market the cash will lead the futures price discovery. A recent example is 1986 in which 81% of the new information contained in prices originated in the futures market, in spite of a \$25/cwt cash hog price swing during the year.

**Table I**
**PARAMETER ESTIMATES FOR GARBADE AND SILBER TEST FOR PRICE DISCOVERY IN LIVE HOG CASH AND FUTURES MARKETS, 1975–1989[a]**

| | Parameter | | | |
|---|---|---|---|---|
| **Year** | $B_c$ | $B_f$ | $Bc/(B_c + B_f)$ | $\delta$ |
| 1975 | 0.058** | 0.137** | 0.30 | 0.81 |
| | (2.78) | (4.34) | | |
| 1976 | 0.121** | 0.067 | 0.64 | 0.81 |
| | (5.54) | (0.64) | | |
| 1977 | 0.074** | 0.021 | 0.78 | 0.91 |
| | (4.93) | (0.46) | | |
| 1978 | 0.074** | 0.115** | 0.39 | 0.81 |
| | (3.75) | (1.98) | | |
| 1979 | 0.066** | 0.020 | 0.77 | 0.91 |
| | (3.08) | (0.30) | | |
| 1980 | 0.033** | 0.083 | 0.28 | 0.88 |
| | (1.83) | (1.21) | | |
| 1981 | 0.068** | 0.213** | 0.24 | 0.72 |
| | (3.47) | (3.03) | | |
| 1982 | 0.165** | −0.177** | 1.00[b] | 1.00 |
| | (5.53) | (−2.06) | | |
| 1983 | 0.085** | 0.135* | 0.39 | 0.78 |
| | (3.78) | (1.77) | | |
| 1984 | 0.095** | 0.051 | 0.65 | 0.85 |
| | (4.52) | (1.15) | | |
| 1985 | 0.057** | 0.031 | 0.65 | 0.91 |
| | (3.18) | (0.82) | | |
| 1986 | 0.137** | 0.032 | 0.81 | 0.83 |
| | (4.93) | (0.42) | | |
| 1987 | 0.086** | −0.034 | 1.00 | 0.95 |
| | (4.48) | (−0.89) | | |
| 1988 | 0.069** | −0.91** | 1.00 | 1.00 |
| | (3.87) | (−2.23) | | |
| 1989 | 0.049** | 0.020 | 0.71 | 0.93 |
| | (3.26) | (0.94) | | |

[a]Numbers in parentheses are *t*-statistics. Single and double asterisks indicate parameter is significantly different from zero at the 0.10 and 0.05 levels, respectively, using a one-tailed test.
[b]Following Garbade and Silber this ratio is constrained between 0 and 1 in that $B_c$ and $B_f$ are both assumed positive.

Essentially, with the exception of 1980 (and perhaps 1986), the daily cash and live hog futures markets were not cointegrated during the 1975–1989 period. This is consistent with the results of Bessler's (1990) tests for cointegration in live cattle cash and futures prices which suggest "... at best, marginal support for the cointegration hypothesis" (p. 11). Given these results, one must question the implications. That is, are the markets inefficient? Or do they represent different assets? As noted earlier, efficient markets for the same asset are cointegrated, whereas, efficient markets for different assets cannot be cointegrated [Granger (1986)].

Table II
AUGMENTED DICKEY-FULLER COINTEGRATION
TEST STATISTICS FOR LIVE HOG CASH AND FUTURE PRICES, 1975–1989[a]

| | Cash = $a + b$ Futures | | Futures = $a + b$ Cash | |
|---|---|---|---|---|
| Year | Augmented Dickey-Fuller Test Statistic | Lags[b] | Augmented Dickey-Fuller Test Statistic | Lags[b] |
| 1975 | 2.92 | 2 | 2.95 | 2 |
| 1976 | 1.24 | 4 | 1.76 | 4 |
| 1977 | 1.88 | 2 | 2.09 | 2 |
| 1978 | 2.02 | 2 | 2.03 | 2 |
| 1979 | 2.35 | 2 | 2.46 | 2 |
| 1980 | 4.19** | 0 | 4.33** | 0 |
| 1981 | 1.69 | 4 | 2.02 | 4 |
| 1982 | 2.41 | 3 | 2.28 | 3 |
| 1983 | 2.08 | 2 | 2.35 | 2 |
| 1984 | 1.80 | 4 | 2.18 | 4 |
| 1985 | 2.04 | 2 | 2.00 | 2 |
| 1986 | 3.17* | 2 | 3.26** | 2 |
| 1987 | 2.23 | 2 | 2.32 | 4 |
| 1988 | 1.72 | 4 | 1.48 | 2 |
| 1989 | 1.02 | 4 | 2.21 | 0 |

[a]Single and double asterisks indicate statistical significance at the 0.10 and 0.05 levels, respectively. A significant test statistic implies rejection of the null hypothesis of lack of cointegration.
[b]Selected using the FPE criterion of Akaike (1969).

To make conclusions regarding market efficiency one must decide whether cash and nearby futures markets for live hogs represent the same asset. Indeed cash and futures price during the delivery period (approximately the first three weeks of the contract month) on the futures contract arguably represent the same asset. However, the farther away the delivery date, the less tie one expects between the cash and futures prices. Thus, during some of the periods of analysis the cash price is being compared to the nearby futures price which does not expire for as many as seven to eight weeks. In these cases the futures price clearly does not represent the same asset in terms of delivery time as the cash market. Large price swings can, and often do occur in live hog markets within fairly short periods of time. For example, over the 1975–1989 period the average absolute value of the daily price change in the Omaha cash hog market was approximately $0.70/cwt.

To formally test the hypothesis that time to delivery has an impact on the cointegration relationship between cash and futures prices, the squared residuals ($\hat{u}_t^2$) of the cointegrating regressions are plotted and regressed against approximate weeks ($WKS_t$) to futures contract expiration. Figure 1 shows the plots of the squared cointegration residuals against weeks to futures contract expiration for 1989 (the residual patterns for the remaining 14 years are similar to those of 1989). The plots show that, although some of the cointegration residuals are near zero for each week, there is a clear divergence between the live hog cash and futures prices as weeks to futures contract expiration increase. Through examination of Figure 1 it is evident that a regression of the squared cointegration residuals on the weeks to ex-

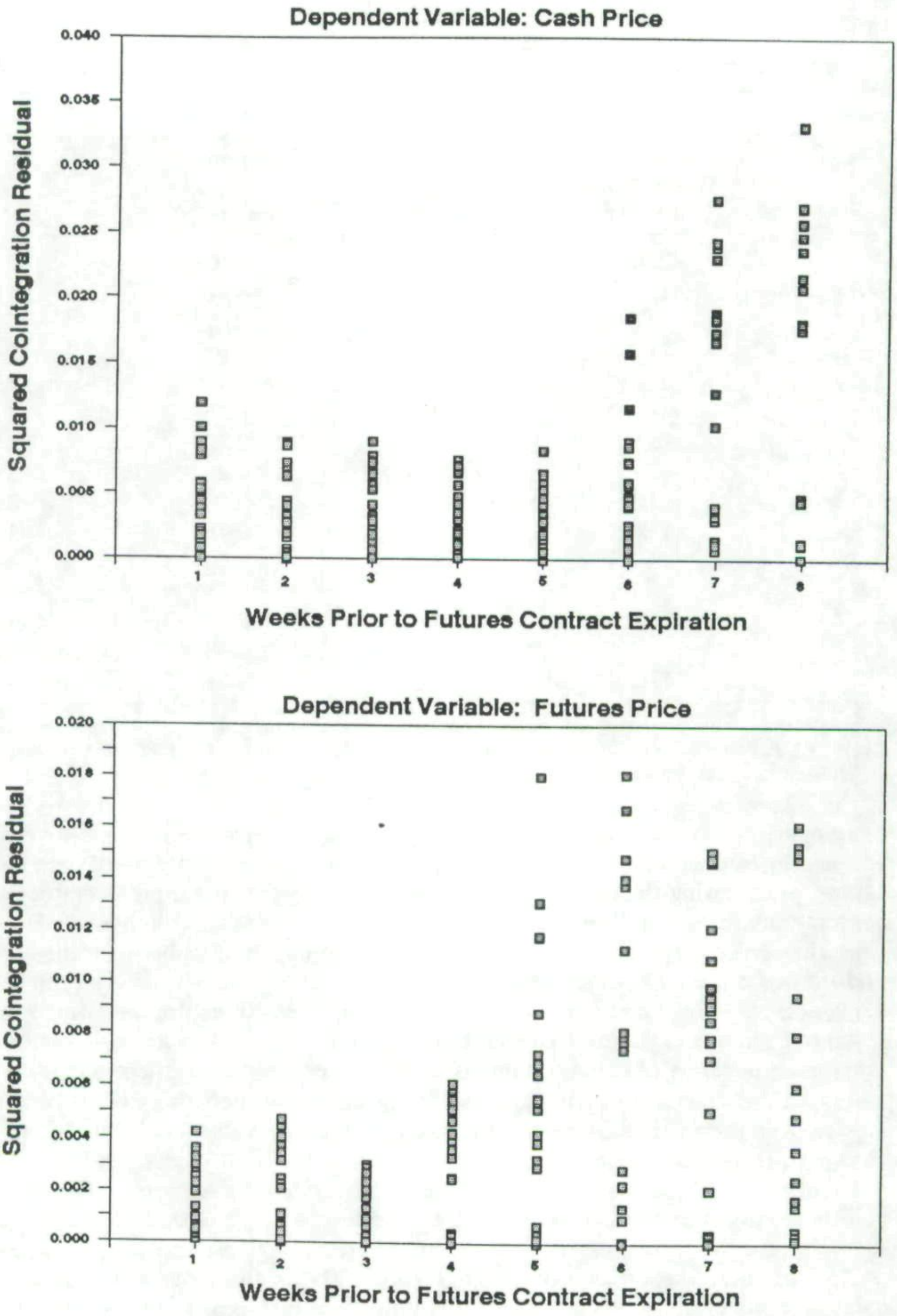


**Figure 1**
Plots of squared cointegration residuals against weeks to futures contract expiration, 1989.

piration would likely suffer from heteroscedasticity (nonconstant residual variance) resulting in inefficient parameter estimates. Thus, these regression equations are estimated using both ordinary least squares and weighted least squares. The resulting regressions from these two approaches are very similar. For 1989 the following weighted least squares regressions are obtained.

Dependent variable = cash price:

$$\hat{u}_t^2 = \underset{(-2.639)^*}{-0.00385} + \underset{(7.977)^*}{0.00205}\, WKS_t \quad R^2 = 0.24 \tag{7}$$

Dependent variable = futures price:

$$\hat{u}_t^2 = \underset{(-0.813)}{-0.00078} + \underset{(5.884)^*}{0.00099}\, WKS_t \quad R^2 = 0.20 \tag{8}$$

$t$-statistics are reported in parentheses, and * indicates that the respective coefficient is statistically different from zero at the 0.01 level. Similar regressions of the squared cointegrating residuals on weeks to expiration for the remaining 14 years confirm the same implications. In every case, the coefficient on *WKS* is positive and highly significant, indicating that weeks to contract expiration have a positive and significant correlation with the cointegrating residuals. Thus, as expected, the farther from contract expiration, the more the cash and futures prices tend to diverge from each other.

The effect of weeks to futures contract expiration on cointegration between hog cash and futures prices is further investigated by including the weeks to contract expiration (*WKS*) as an exogenous variable in the cointegrating regression. Equation (4) is thus modified as follows:

$$C_t - a' - b'F_t - c'WKS_t = u'_t \tag{9}$$

Equation (9) represents the first stage of the cointegration test with weeks to futures contract expiration accounted for.[4] Cointegration tests on the residuals of estimates of eq. (9) are reported in Table III. These tests yield similar results to those presented in Table II (i.e., those in which *WKS* is not included as a regressor). The only different result from including *WKS* as a regressor in the cointegration test is that in 1986 the cash and futures prices are not cointegrated as they are in the models that exclude *WKS*. The remaining conclusions are identical for each of the models.

Finally, in light of the presence of heteroskedasticity across weeks to contract expiration among the residuals of all of the ordinary least squares cointegrating regressions estimated, they are all reestimated using weighted least squares. The weighted least squares results are similar and the implications identical to those estimated using OLS and are therefore not reported.

[4]In the absence of either a continuous trend in prices or of a systematic bias in the cash or futures prices, one would expect that in the long run the estimate of $c'$ in eq. (9) would not be significantly different from zero. However, in any given year having a trend in prices, the estimate of $c'$ could be significantly different from zero. For example, in an upward trending market the cash price may increase over time adjusting upward to the levels the futures price has already attained and accurately predicted. In this case, $c'$ would represent the trend in prices over the period. Over the 1975–1989 period, of the 30 separate equations estimated with *WKS* as a regressor (15 with the cash price as the dependent variable and 15 having the futures price as the dependent variable), 14 had $c'$ estimates significantly different from zero. The estimates of $c'$ does not have a systematic sign indicating that both upward and downward trends are present over the periods investigated.

**Table III**
**AUGMENTED DICKEY-FULLER COINTEGRATION TEST STATISTICS FOR LIVE HOG CASH AND FUTURES PRICES AFTER ACCOUNTING FOR WEEKS TO CONTRACT EXPIRATION IN THE COINTEGRATING REGRESSIONS, 1975–1989[a]**

| | Cash = $a' + b'$ Futures + $c'$ *WKS* | | Futures = $a' + b'$ Cash + $c'$ *WKS* | |
|---|---|---|---|---|
| | Augmented Dickey-Fuller Test Statistic | Lags[b] | Augmented Dickey-Fuller Test Statistic | Lags[b] |
| 1975 | 2.92 | 2 | 2.93 | 2 |
| 1976 | 1.48 | 4 | 1.85 | 4 |
| 1977 | 1.87 | 3 | 1.97 | 3 |
| 1978 | 2.06 | 3 | 1.87 | 3 |
| 1979 | 2.28 | 2 | 2.43 | 2 |
| 1980 | 4.23* | 0 | 4.35* | 0 |
| 1981 | 1.54 | 4 | 1.83 | 4 |
| 1982 | 2.41 | 3 | 2.27 | 3 |
| 1983 | 2.10 | 2 | 2.29 | 2 |
| 1984 | 1.95 | 4 | 2.15 | 4 |
| 1985 | 2.05 | 2 | 1.98 | 2 |
| 1986 | 2.95 | 2 | 2.98 | 2 |
| 1987 | 1.72 | 3 | 2.04 | 2 |
| 1988 | 1.72 | 4 | 1.43 | 2 |
| 1989 | 1.23 | 4 | 2.24 | 0 |

[a]Asterisks indicate statistical significance at the 0.05 level. A significant test statistic implies rejection of the null hypothesis of lack of cointegration.
[b]Selected using the FPE criterion of Akaike (1969).

These results suggest that long-term basis risk (basis fluctuations) is present between the Omaha cash and CME futures live hog prices. This is especially so during the nondelivery periods, when the cash and futures prices are valuing different commodities. Although it is also evident from Figure 1 and the above regressions that in the longer-term (over the period from 8 weeks to 1 week prior to contract expiration) the hog cash and futures prices do tend to converge.

## CONCLUDING COMMENTS

This article examines short- and long-run price relationships between Omaha cash and CME futures daily prices for live hogs. Price discovery generally originates in the futures market with an average of roughly 65% of new information being passed from the futures to the cash market. Thus, the Omaha cash market tends to be a satellite to the CME live hog futures contract. Little short-run feedback is generally present from the cash market to the futures market. However, at times, especially during large price moves that are not necessarily anticipated in the futures market, the cash market price discovery dominates (or at least relies less on) the futures market.

The very short-term (day to day) basis for hogs is fairly stable. On average, approximately 85% of yesterday's nearby basis persists today. Futures and cash prices do not converge rapidly to each other and little can generally be gained by speculating on basis from day to day.

In spite of causality from the futures to the cash market, the two price series still operate somewhat independently and the long-term basis is generally nonstationary. Today's cash price is likely responding to the futures price in addition to variations in short-term demand for hogs by packers and varying deliveries from day to day. The futures price, on the other hand, is responding to information expected to influence subsequent cash market prices (in this study at times as far as seven to eight week hence).[5] The farther from delivery, the more the futures and cash prices diverge from each other. Thus, a lack of cointegration is not surprising between cash and futures prices. In addition, a lack of cointegration does not preclude causality between the cash and futures prices. Granger (1986) states that if two variables are cointegrated "there must be Granger causality in at least one direction, as one variable can help forecast the other" (p. 218). The converse of this does not necessarily hold. That is, the futures price can Granger cause the cash price in that knowledge of the futures price could help one forecast future values of the cash price, yet the two could have sufficient stochastic properties independent of each other such that the series are not cointegrated. Hedgers liquidating cash and futures positions prior to the contract delivery month face relatively larger basis risk than those liquidating positions in the contract month and as the delivery period approaches the cash and futures prices tend to converge.

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[5]Perhaps further developments in time series methods and empirical designs will allow future research to focus cash and futures price cointegration testing on solely the contract delivery period. To maintain a consistent, continuous data series this study is unable to do this.

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