Feeder Cattle Cash Settlement: Hedging Risk Reduction or Illusion?

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INTRODUCTION

One of the primary economic functions of the futures market is to provide a mechanism by which producers can transfer cash price risk to others. By hedging, market participants seek to exchange cash price risk for a more stable basis risk. From 1980 to 1986 excessive basis risk is believed to have reduced the effectiveness of the feeder cattle contract as a hedging mechanism [CME (1985A)].

The Chicago Mercantile Exchange (CME) initiated a mandatory cash settlement system for the feeder cattle contract in September, 1986. The move from physical delivery to cash settlement was designed to:

> (1) eliminate the uncertainties and disputes associated with the grading of CME feeder deliveries; (2) eliminate the risk of a long receiving delivery at an inconvenient delivery location and/or receiving an undesirable type of feeder; (3) eliminate the costs incurred in making or taking delivery on the contract; (4) eliminate the need for periodic contract amendments regarding discounts for non-par grades, weights, and locations . . . ; and (5) reduce the basis variability for users of the contract (CME, 1985B, p. 9).

In sum, cash settlement was expected to improve contract performance by reducing hedgers' risks during the expiration months.[1] However, from a skeptical perspective, one could argue that a settlement price derived by averaging cash prices across time and

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[1]Paul (1985, 1987); Kahl, Hudson, and Ward (1989); and others point out that for cash settlement to be effective, the settlement price must be free of exploitation and accurately depict cash market prices. To improve pricing accuracy, the settlement price must be determined by cash market demand and supply conditions, representing current cash market commodity value. This lessens arbitrage opportunities, and chances for misrepresentation.

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space could create the illusion that maturity basis variability is reduced; when, in fact, actual hedging conditions for individual producers have remained largely unchanged.[2]

The purpose of this article is to ascertain how cash settlement issues have influenced hedging conditions at the individual market level and for the feeder cattle industry in general. Cash settlement is estimated to be responsible for a 7.25% decline in hedging risk for the feeder cattle industry; yet, the individual market results tend to be ambiguous. The overall conclusions, however, are not completely negative. While not statistically linked to the introduction of cash settlement, it is found that maturity basis risk has declined slightly at most individual markets. This, combined with mean bases converging distinctly closer to zero at expiration, which appears to be clearly attributable to the presence of cash settlement, implies that the settlement change enhances contract performance in general.

BACKGROUND

Preliminary evidence on feeder cattle cash settlement is mixed. Elam (1988) developed a simulation model to measure the estimated change in hedging risk due to cash settlement for Arkansas producers. He found that hedging risk (as measured by estimated hedge ratios) decreased with cash settlement, especially for heavier weight Arkansas feeders. Following Elam's procedures, Schroeder and Mintert (1988) used data from four cash markets (Amarillo, Dodge City, Kansas City, and Illinois Direct) and found results similar to Elam's. Caution may be in order when generalizing these results, however. Elam and Schroeder et al. used U.S. feeder steer prices (USFSP) to proxy for actual post futures prices. Kenyon et al. (1991) found, using actual post futures prices in a basis investigation concentrating on Virginia feeder cattle, that the introduction of cash settlement did *not* significantly affect basis variability. All studies to date used limited postcash settlement data points and analyzed a small number of markets. As Kenyon et al. put it, their "result need[s] to be tested and evaluated in other markets using actual cash and futures prices under cash settlement as more data are generated over time" (p. 104). The current study uses more powerful tests and expands the sample in both size and length: 27 individual markets are analyzed over a period of 54 contracts (3.5 years of data pre and post the settlement change).

When cash settlement replaced physical delivery for feeder cattle futures contracts, there was not only a change in the settlement procedures but also changes in contract specifications which could have affected basis levels and basis patterns.[3] In this article, mean basis levels and hedging risk are analyzed for 27 precash settlement contracts (April, 1983 to August, 1986) and 27 postcontracts (September, 1986 to November, 1989).[4] The analysis is conducted for 600–700 pound heifers and steers across 27 selected U.S. price reporting stations, located in 20 different states.

[2]Kenyon et al. (1991) found anecdotal evidence in support of the skeptical point of view. But it should be noted that while some academicians and practitioners voiced this hypothesis, no documented form of this hypothesis is found. Also note that the same arguments for cash settlement's effect on maturity basis risk can be made to deferred basis risk if futures prices are "tracking" the index. For a discussion on the relationship between the settlement index price and deferred and nearby futures prices, see Rowsell and Purcell (1990).

[3]Contract specification changes include deleting par delivery points, altering muscles score requirements, adjustments to weight specifications, position limit increases, and settlement day changes. See Rich (1988) for a detailed discussion of each of these changes.

[4]There are eight trading months for the feeder cattle futures contract: January, March, April, May, August, September, October, and November.

Feeder cattle contracts outstanding at contract expiration are now settled at the Cattle-Fax USFSP. This is a weighted-average 600–800 pound feeder steer price taken from prices in 27 states. Natural constraints (e.g., geographic location, sex, etc.) often require producers to hedge feeder cattle via cross hedges. "Feeder cattle that do not meet the contract weight, sex, or grade specifications are different commodities than those prescribed by the feeder cattle futures contract specifications" [Schroeder and Mintert (1988, p. 317)]. The most correct test of cash settlement should focus on the results from animals that most closely conform to contract specifications. However, it is still of interest to see how cash settlement affects cross hedges.

Kenyon et al. concluded that maturity basis variability has not been reduced for Virginian hedgers. Because their finding differs markedly from that of Elam and Schroeder et al. one wonders whether differentiating those markets in the USFSP formula from those not included in the settlement derivation process can explain the difference. Therefore, this study includes four markets that are *not* part of the USFSP formula: Illinois Direct, IL; Virginia Auctions, VA; S. St. Paul Terminal, MN; and South Carolina Auctions, SC. Basis behavior in these lightly traded, cross hedging markets, are assessed relative to those included in the price formula.

Four main methods are used to assess the impact of cash settlement on the feeder cattle contract. First, parametric tests are performed on weekly delivery week mean basis levels and basis variability pre- and postcash settlement. Weekly data, however, introduce two potential, not necessarily mutually exclusive, biases: (1) hedging risk estimates may be downward biased to the extent that averaging prices across time can gloss over short-run price volatility; and (2) most feeder cattle cash markets hold their "major" auction(s) one (or two) day(s) per week and use an alternative day (usually the day after) as a "clean-up" day. "Clean-up" calves usually sell at a substantial discount because they are much less likely to meet contract specifications. Therefore, for the second test, "clean up" prices are omitted and the initial parametric tests are conducted on daily data from two of the largest markets: Oklahoma City, OK and Dodge City, KS.

The third method of analysis, which deviates from previous research, uses parametric tests that are less sensitive to the normality assumption (of the first test) and nonparametric tests on delivery week mean basis levels and basis variability pre- and postcash settlement. Preliminary distributional tests suggest that the assumption of basis being normally distributed for many of the individual markets is questionable. Fourth, to separate cash settlement's bearing on basis moments from that of other economic variables, a regression analysis is conducted.

DATA AND METHODOLOGY

Data

Ideally, to obtain the most accurate interpretation of the impact of cash settlement on the entire feeder cattle industry, all of the cash markets that report prices should be analyzed. The CME reports 43 feeder cattle price reporting stations. However, several of the markets are not included in the sample size due to an insufficient amount of available data. Any market missing more than five consecutive weekly prices is eliminated from the analysis.

Most of the 43 feeder cattle price reporting stations report prices in 100 pound increments for 200–800 pound feeders (i.e., 200–300, 300–400, etc.). Furthermore, the reporting stations subdivide these prices in each 100 pound range into steers and heifers. Because most of the lighter weight groups (e.g., 200–500 pounds) for both sexes are thinly traded, they are disregarded in this analysis. Of the remaining weight groups (600–700 and 700–800 pounds), only the 600–700 pound range is investigated because this weight range contains the fewest missing prices. Given the above considerations, the sample size for this analysis includes 27 cash market terminals located in 20 different states (see Table I).

All of the weekly and daily cash and futures price data used in this study are from the CME Research Department. Weekly cash price i represents the average cash price, during the week that the futures contract expires, for the ith reporting station. Likewise, the weekly futures price represents the average futures price during the contract expiration week. The actual final futures price at expiration is used rather than the USFSP price.

Table I
FEEDER CATTLE CASH MARKETS INCLUDED IN SAMPLE

Reporting Station	Location
Alabama Auctions	Montgomery, AL
Amarillo Auctions	Amarillo, TX
Arkansas Auctions	Little Rock, AR
Clovis Auctions	Clovis, NM
Colorado Auctions	Greeley, CO
Georgia Auctions	Thomasville, GA
Illinois Direct	Springfield, IL
Iowa Auctions	Des Moines, IA
Iowa–S. Minnesota Direct	Des Moines, IA
Kansas City Terminal	Kansas City, MO
Kentucky Direct	Louisville, KY
Lexington Auctions	Lexington, KY
Louisville Auctions	Louisville, KY
McKinley–Winter Terminal	Dodge City, KS
N. San Joaquin Direct	Stockton, CA
Oklahoma City Auctions	Oklahoma City, OK
S. St. Joseph Terminal	S. St. Joseph, MO
S. St. Paul Terminal	S. St. Paul, MN
Shasta Auctions	Shasta, CA
Sioux City Terminal	Sioux City, IA
Sioux Falls Terminal	Sioux Falls, SD
South Carolina Auctions	Columbia, SC
Springfield Auctions	Springfield, MO
Tennessee Auctions	Nashville, TN
Virginia Auctions	Richmond, VA
Washington–Oregon Direct	Moses Lake, WA
West Fargo Terminal	West Fargo, ND

Delivery Week Parametric Tests Using Weekly Data

Risk Test. Hedging risk is proxied in this article by the sample standard deviation of the basis (in dollars per hundredweight).[5] Basis (B) is calculated as the cash (C) price minus the futures (F) price, or:

$$B = C - F \tag{1}$$

Using delivery week cash market data, expiration basis variability is calculated for each of the 27 price reporting stations. Variability is calculated only for the delivery weeks in which cash prices are reported; if there are no missing prices, there are 27 BEFORE observations and 27 AFTER for each sex at each of the 27 individual markets. Under physical delivery (BEFORE), contracts expired on the 20th of each delivery month. Following the change in specifications (AFTER), contracts cease trading on the last Thursday of the contract month.

Assuming cash settlement has the desired impact in reducing maturity hedging risk, the hypothesis of no change in the variance is tested:

$$H_o\colon \sigma^2_{\text{AFTER}} = \sigma^2_{\text{BEFORE}} \text{ against } H_1\colon \sigma^2_{\text{AFTER}} < \sigma^2_{\text{BEFORE}} \tag{2}$$

This hypothesis will be tested on 52 pairs of variances.[6] The test statistic is an F-test.

Mean Basis Tests. The successfulness of hedging with the futures market is determined primarily by the predictability of basis movement. Nevertheless, contract specification changes certainly have altered mean basis levels. Therefore, it is helpful to know at what mean level basis patterns are fluctuating around. The mean delivery week basis for each market is computed as a linear average ($\Sigma_i x_i/n$) of the delivery week bases for that time period (BEFORE or AFTER).

When cash settlement was initiated, Cohen and Gorham (1985) argued that it was unclear as to whether the settlement change would increase or decrease mean basis levels. They reasoned that mean futures prices could increase relative to cash prices (i.e., reduce mean basis levels) because prices from various markets would be averaged rather than representing the cheapest to deliver of the physical commodity. On the other hand, "looser quality specifications [under cash settlement] and the inclusion of unadjusted prices for cheaper feeder trade areas" could reduce mean future prices and raise mean basis levels (p. 323). Their empirical work revealed that if cash settlement had been in place from 1980 to 1984, estimated lower maturity futures prices would have resulted in roughly a \$4.00/cwt. increase in mean basis levels. Therefore, the following hypothesis

[5]It was suggested that examining the basis in the form of (1) could be inadequate do to stationarity problems. For example, if cash price and futures price behavior were modeled as geometric Brownian motion: $dC/C = \alpha_c dt + \sigma_c dz_c$, $dF/F = \alpha_f dt + \sigma_f dz_f$; the new prices would be: $C + dC = C(\alpha_c dt + \sigma_c dz_c) + C$, $F + dF = F(\alpha_f dt + \sigma_f dz_f) + F$, and the basis would be:

$$\text{BASIS} = (C - F) + (C - F)[(\alpha_c - \alpha_f)dt + \sigma_c dz_c - \sigma_f dz_f]$$

where dz_i is a Wiener process, and α_i and σ_i are, respectively, the instantaneous expected return and variance. That is, basis could be nonstationary as its mean and variance would change with C and F. The nonstationarity could, then, make the mean and variance estimated (listed in Table II) unreliable. This concern was tested by examining the daily Oklahoma City basis data. The results were virtually identical (to those listed in Table III) whether price relatives or raw prices were used. Stationarity doesn't appear to be an issue.

[6]There are 27 price reporting stations, each containing two categories: 600–700 pound steers and 600–700 pound heifers. However, only steer data are available from the Georgia Auctions and Louisville Auctions. Therefore, there are 52 pairs of data.

is tested on mean basis levels:

$$H_o\colon \mu_{\text{AFTER}} = \mu_{\text{BEFORE}} \qquad H_1\colon \mu_{\text{AFTER}} > \mu_{\text{BEFORE}} \tag{3}$$

This hypothesis is tested on 52 pairs of means. The test statistic is a Studentized or approximate t test depending on whether variability has significantly changed [i.e., the results from testing (2)].[7]

Delivery Week Parametric Tests Using Daily Data

Ideally, to get the most accurate assessment of the impact of cash settlement, daily data should be used. Daily data allow short-run basis risk to be analyzed in more detail. Unfortunately, there are only three markets from which daily cash price steer information could be obtained: Amarillo, TX; Oklahoma City, OK; and Dodge City, KS. These are, however, the three highest volume markets. Therefore, their results should serve as reliable proxies of the industry.[8]

Hypotheses (2) and (3) are tested using daily data. However, to eliminate any potential biases, basis is calculated only on "major" trading days during the final week of trading. The "major" auctions are: Amarillo each Tuesday, Dodge City each Wednesday, and Oklahoma City each Monday and Tuesday.[9]

Delivery Week Levene and Nonparametric Tests

Weekly steer basis data during delivery week from ten of the largest price reporting stations are tested for Normality using the Shapiro–Wilk statistic [see *SAS* (1990)]. In a (somewhat) random subsample, Normality is rejected in 70% of the tests—due, generally, to "heavy" tails.[10] Therefore, it is questionable whether the F-ratios and t tests, as described above (and reported in Table II) are appropriate. To address this concern, hypotheses (2) and (3) are retested (using weekly data) and using the modified Levene test and distribution-free tests.

It is well known that the F test is only valid for Normally distributed data. The Levene statistic can be used in lieu of the F statistic to test the null hypothesis on equal variances when the assumption of Normality is questionable. In fact, Hsieh and Miller (1990) assert that "the modified Levene statistic [the Levene statistic calculated using a trimmed mean] is *not* sensitive to departures from Normality" (p. 7) (brackets and underline were added). Following Hsieh and Miller (and the references cited within) (2) is retested using a 5% trimmed mean in the modified Levene test (see bottom of Table IV for computation of the Levene statistic).

[7]The BEFORE and AFTER population variances are unknown but are assumed to be equal or unequal depending on the results from testing hypothesis (2). When the null hypothesis in (2) is accepted, an exact Studentized t test is used to test the difference between the two means. When the variances are unknown and assumed to be unequal [i.e., (2) is rejected] no exact test exists; an approximate t test (sometimes called the Welch test) must be used. (The test is approximate due to rounding in the calculation of the degrees of freedom.) See Ostle (1975, pp. 118–125) for a discussion of each method.

[8]For some price reporting stations, daily and weekly data could be identical if there is only one market day per week. No attempt is made to identify those cases.

[9]Only two markets are used in the final analysis. Amarillo data is discarded due to changes in the "major" trading day during the sample period.

[10]The sampling procedure is not completely random because the three largest markets are intentionally included in the subsample. Of the three largest markets, only Oklahoma City tests Normal. Due to the sample size, Normality is tested by using the Shapiro–Wilk W-statistic. For Oklahoma City, Amarillo, and Dodge City, Prob $< W$ was $0.92, 0.21, 0.002$, respectively.

Table II
DELIVERY WEEK BASIS STANDARD DEVIATIONS AND MEANS, PRE- AND POSTCASH SETTLEMENT FOR 600–700 POUND FEEDER CATTLE, 1983–1989

	Heifers			Steers		
Reporting Station	**BEFORE**	**AFTER**	***p*-value**	**BEFORE**	**AFTER**	***p*-value**
	(Standard deviations are on the top line) (mean basis levels are on the second line)					
Amarillo Auctions	1.77	1.41	0.1318	1.30	1.52	0.2122
	−8.45	−4.89	0.0000	−2.33	0.81	0.0000
McKinley–Winter	1.71	1.71	0.4995	1.46	1.75	0.1772
Terminal	−7.28	−2.58	0.0000	−1.50	3.20	0.0000
Oklahoma City Auctions	1.51	1.31	0.2392	1.69	1.33	0.1153
	−7.89	−2.72	0.0000	−1.01	4.25	0.0000
Alabama Auctions	2.42	2.00	0.1809	1.94	1.95	0.4933
	−15.20	−9.64	0.0000	−7.11	−2.85	0.0000
Arkansas Auctions	2.11	1.97	0.3684	1.72	1.92	0.2893
	−11.18	−6.48	0.0000	−3.53	0.26	0.0000
Clovis Auctions	1.67	1.56	0.3781	1.89	2.12	0.2920
	−10.76	−6.81	0.0000	−3.23	0.36	0.0000
Colorado Auctions	1.58	1.05	0.0244[a]	1.98	1.65	0.1875
	−6.65	−2.38	0.0001	−1.02	3.33	0.0000
Georgia Auctions				2.54	2.36	0.3533
				−6.73	−2.28	0.0000
Illinois Direct	2.54	2.02	0.1457	2.86	2.82	0.4767
	−7.60	−0.92	0.0000	−0.92	4.34	0.0000
Iowa Auctions	2.60	1.88	0.0683[b]	2.26	1.93	0.2231
	−8.37	−3.71	0.0001	−1.73	0.94	0.0000
Iowa–S. Minnesota Direct	2.22	1.68	0.0975[b]	2.26	2.25	0.4996
	−7.36	−2.50	0.0001	−0.30	3.79	0.0000
Kansas City Terminal	2.33	1.61	0.0890[b]	2.40	2.20	0.3895
	−7.95	−3.45	0.0001	−0.67	2.94	0.0000
Kentucky Direct	2.87	1.94	0.0256[a]	2.10	2.03	0.4259
	−13.94	−8.03	0.0001	−4.98	−1.74	0.0000
Lexington Auctions	2.91	2.60	0.3184	2.43	2.53	0.4109
	−12.32	−6.90	0.0000	−4.27	0.21	0.0000
Louisville Auctions				2.44	2.35	0.4262
				−4.90	0.15	0.0000
N. San Joaquin Direct	1.77	2.03	0.2543	2.18	2.68	0.1552
	−12.03	−5.13	0.0000	−4.29	0.45	0.0000
S. St. Joseph Terminal	2.24	1.24	0.0029[a]	2.31	1.67	0.0570[b]
	−8.16	−3.01	0.0001	−1.06	3.68	0.0001
S. St. Paul Terminal	2.53	1.87	0.0705[b]	2.81	1.95	0.0359[a]
	−10.60	−6.57	0.0001	−3.26	0.87	0.0001

(continued)

Table II (continued)

Reporting Station	Heifers			Steers		
	BEFORE	AFTER	p-value	BEFORE	AFTER	p-value
	(Standard deviations are on the top line) (mean basis levels are on the second line)					
Shasta Auctions	2.07	3.96	0.0010[a]	2.29	2.95	0.1045
	−11.42	−3.90	0.0001	−3.56	1.93	0.0000
Sioux City Terminal	2.33	1.91	0.1620	1.96	1.94	0.4732
	−7.62	−2.52	0.0000	−1.49	4.32	0.0000
Sioux Falls Terminal	1.89	2.09	0.3060	2.03	2.17	0.3696
	−6.96	−1.24	0.0000	−0.31	5.21	0.0000
South Carolina Auctions	3.97	2.34	0.0041[a]	2.61	2.30	0.2740
	−17.08	−11.30	0.0001	−8.76	−5.22	0.0000
Springfield Auctions	2.09	1.89	0.3095	2.33	2.74	0.2050
	−10.31	−4.19	0.0000	−2.68	0.02	0.0003
Tennessee Auctions	2.78	1.95	0.0407[a]	2.20	2.58	0.2078
	−14.45	−9.11	0.0001	−6.21	−1.83	0.0000
Virginia Auctions	4.39	2.67	0.0092[a]	4.05	3.20	0.1266
	−14.99	−8.91	0.0001	−3.25	0.24	0.0012
Washington–Oregon Direct	2.82	2.20	0.1240	3.18	2.96	0.3710
	−10.24	−5.08	0.0000	−3.10	0.44	0.0002
West Fargo Terminal	2.03	1.40	0.0344[a]	1.99	1.98	0.4880
	−9.18	−3.46	0.0001	−2.67	2.22	0.0000

The variances are tested with a one-tailed right tailed F-test. The p-values for the means are based on a t test or approximate t test (see discussion in text).

[a]Significant at the 5-% level.

[b]Significant at the 10-% level.

Note: All the means are highly significant, and have not been marked.

"The F statistic rejects the null hypothesis of equal variances too frequently, if the data have heavier tails than the Normal distribution" [Hsieh and Miller (1990, p. 7)]. Thus, based on the findings from the random sample, this study expects to find fewer significant occasions using the Levene test.

There is not a "Levene type" statistic that can be used to retest (3). Therefore, hypotheses (2) and (3) are retested using nonparametric tests. Hypothesis (2) is retested two ways: (1) on each sex for each of the 27 markets, and (2) a cross-sectional composite test to compare all the BEFORE standard deviations with all the AFTER. First, steers and heifers at each market are tested individually as follows [see Ohlson and Penman (1985)]. The mean basis level of the relevant time period is subtracted from each individual basis observation. This quantity is then squared and proxied as basis variance.[11] There are 27 squared mean adjusted bases proxing for delivery time variance for each time period

[11]This follows because the expected value of basis squared is a very close approximation to basis variance if the mean basis levels for each time period are similar. To the extent that mean basis levels may have increased vis-a-via the settlement change, this test may be biased in favor of acceptance of the alternative hypothesis in (2).

for each sex (if there are no missing observations at a market). The two time periods for a market and sex may then be compared [i.e., hypothesis (2) retested] by calculating a Mann–Whitney Z-score [Newbold (1984)].[12] Intuitively, this test is comparing the BEFORE standard deviation distribution to the AFTER to see if the central location of the BEFORE is higher than that of the AFTER. Following this, the traditional Mann–Whitney U-test is used as the cross-sectional variance test. That is, all of the original BEFORE standard deviations calculated for hypothesis (2) are compared to all of the AFTER using the Mann–Whitney U-test.

There is not a direct nonparametric test corresponding to hypothesis (3). However, the traditional Mann–Whitney U-test can be employed again to test the cross-sectional change of central location in mean basis levels.

Note that seasonality in basis volatility can bias the parametric and nonparametric tests. This could be relevant since Elam (1988) detected seasonality in feeder cattle basis risk. He found basis risk for Arkansas producers to be higher in the spring delivery months than the fall months. However, Schroeder and Mintert (1988), using very similar techniques to Elam, did not find the spring seasonality component in their study. This study does not find any sign of seasonality at the three largest markets: Amarillo, Dodge City, and Oklahoma City.[13] (Also see footnote 5). Hence, constant seasonality is assumed in the previously mentioned parametric and nonparametric tests.[14]

Regression Model

For the weekly and daily, parametric and nonparametric, basis variability tests discussed above, it is implicitly assumed that any difference between the two time periods is attributable solely to the introduction of cash settlement. To isolate the impact of cash settlement on basis risk, and to gain greater insight into hedging risk across space and between sex, a regression model is developed and estimated.

Using ordinary least squares (OLS), the following cross-sectional model for delivery week standard deviations is estimated:

$$LOG(STD) = f(SEX, CS, L_i) \tag{4}$$

where:
$LOG(STD)$ = the logarithmic transformation of delivery week basis standard deviation;
$SEX = 0$ if heifer, 1 if steer;
CS = cash settlement, 0 if BEFORE, 1 if AFTER;
L_i = a binary dummy variable for location. The sample for the regression consists of 23 price reporting stations. $L1$ represents the first market, $L2$ the second, and so on up to $L23$. Market $L23$ (Washington–Oregon Direct) is used as the base in the analysis. The results are presented in Table II.

[12]The Mann–Whitney test is preferred over the binomial test (used by Ohlson and Penman, 1985) because it is a more powerful test. Further, note that the sample periods are of equal length but not of matched contract months. (There are three November BEFORE contracts and four AFTER contracts.) Hence, any seasonality would result in the Mann–Whitney test being more appropriate than a Wilcoxon test.

[13]Seasonal volatility is tested by regressing the squared mean adjusted basis on 0–1 contract month dummy variables. (November is used as the base month.) None of the t-ratios for any of the contract months at any of the three markets are found to be significant. These results are available upon request.

[14]Kenyon et al. (1987) found significant seasonal volatility in grain futures but concluded "that assuming a constant variance for cattle and hogs is a reasonable assumption."

The four markets which are not part of the settlement index formula are not included in this analysis because geographic differences can not be separated from settlement index formulation differences.[15] Thus, there are a total of 44 risk observations; one pre- and one postcash settlement for each sex of the 23 remaining price reporting stations (also see footnote 6). Assuming cash settlement has lessened uncertainty and eliminated delivery costs, cash settlement is hypothesized to have a negative sign. Delivery week basis risk should be lower under cash settlement. The dummy variable for sex is expected to have a positive sign following Kenyon et al. (1991) and Elam (1988) who both found a greater reduction in basis volatility for heifers than steers after the introduction of cash settlement. The dummy variables for geographic location proxy whether basis risk differs between markets.

The dependent variable is transformed to logarithms for two main reasons. First, the underlying Normality assumption of OLS is violated before the transformation. The null hypothesis that the log transformation is Normal cannot be rejected because the alternative (not Normal) is extremely insignificant (p-value = 0.9802, the smallest possible significance level at which the null can be rejected). Second, the semilogarithmic eq. (4) allows a more precise economic interpretation of the dummy variable coefficients.[16]

RESULTS

Table II presents the BEFORE and AFTER standard deviations and means during the delivery weeks by sex for each of the 27 markets analyzed. The p-values are also listed (see footnote 7). As expected, maturity basis risk generally decreases following the introduction of cash settlement. Hedging risk is reduced for 71% (37/52) of the comparisons.[17] However, only 8 of these 37 F-statistics are significant at the 95% confidence level, and only 13 are significant at the 90% confidence level. Further, of these 13 statistically significant findings, only 2 are steers. None of the largest three markets have a significant decrease in volatility and, in fact, steer basis variability actually increases at Amarillo and Dodge City (McKinley–Winter Terminal). Four of the 13 significant decreases occur in those markets *not* included in the USFSP, a disproportionate amount. That is, cash settlement may provide more benefits in cross-hedging situations. This may have been what Kenyon et al. detected. Finally, of the 15 cases where variability increases, the change is significant in one, the Shasta Auctions. Four of these 15 increases were in California. A possible explanation is that spatial differences and the bulkiness of the commodity make arbitrage difficult and expensive which results in a marketing region which may not be directly related to the principal feeder cattle production area.

More consistent with previous research, mean basis levels are found to have increased (futures price decreases relative to cash) and all are extremely significant. Most, but not

[15]Any attempt to separate index formulation differences from location differences with the use of dummy variables would result in a singular matrix.

[16]All of the weekly and daily cash and futures price data used in this study are from the CME Research Department.

[17]The individual F statistics listed in Table II test the hypothesis of equal variances for each sex of each market. However, a procedure developed by Fisher [see Bickel and Doksum (1977), (p. 188)] can be used to aggregate the p-values from the F-tests to gain insight into how cash settlement affects the feeder cattle industry. Using the statistic, $G = -2\Sigma_i \log \alpha_i$ (where α_i denotes the ith p-value, i runs from 1 to 37 (the cases where variance decreases), and G follows a chi-square distribution with (2)(37) degrees of freedom) it is easy to show that the combined evidence is significant. That is, for those observations where variance decreases, most are not individually significant but when this evidence is combined it is significant. This may be used as evidence in support of the illusion hypothesis discussed in the introduction.

all, of these means are nearer zero after cash settlement. Thus, based on this first test, most of the individual markets show a modest reduction in hedging risk (for the post period) but this is not generally linked to the structural change. (Also see footnote 17). Strong evidence is found that mean basis levels have risen—an apparently direct result of the settlement change.

Table III presents BEFORE and AFTER daily steer basis variability and mean basis levels during the delivery weeks for Dodge City and Oklahoma City. The results in Table III are very similar to those in Table II, except that the decrease in basis variability in Oklahoma City is now significant at the 5% level.[18] Although this is a limited comparison, the change in the Oklahoma City result does suggest that averaging prices across time and including "clean-up" prices can slightly distort actual hedging conditions. Overall, however, the picture remains the same. That is, daily data may more accurately reflect real basis conditions, yet these data are generally unavailable.

The modified Levene statistics are listed in Table IV. As expected, heavier tails than the Normal distribution result in the Levene test detecting fewer significant occasions than the F test. The null hypothesis (2) is rejected at the 95% level only six times. Of the eight significant occasions in Table II, six of these are again found significant in Table IV. Thus, the results between Tables II and IV display some sensitive to the Normality assumption.

The Mann–Whitney Z-scores for each sex for each of the 27 markets are listed in Table V. Recall, the squared mean adjusted basis observations are used as an approximation to basis variance in the ranking process. This test examines the null hypothesis that the location of the distribution of basis variance at each market is invariant to cash settlement's presence. In this respect, this test may be considered the nonparametric equivalent to the F-test discussed above. However, a word of caution is in order. Similar to how heavy tails can bias the F statistic in favor of rejecting the null

Table III
DAILY BASIS STANDARD DEVIATIONS AND MEANS DURING DELIVERY WEEK, PRE- AND POSTCASH SETTLEMENT FOR 600–700 POUND FEEDER CATTLE, 1983–1989

	Steers		
Reporting Station	**BEFORE**	**AFTER**	**p-value**
	(Standard deviations are on the top line) (Mean basis levels are on the second line)		
McKinley–Winter Terminal	1.69	1.81	0.3703
	−1.51	3.24	0.0000
Oklahoma City Auctions	1.82	1.27	0.0065[a]
	−0.89	4.15	0.0001[a]

[a]Significant at the 5% level.

[18]As noted before, Amarillo data are excluded because of a change in the "major" trading day, which is unknown. However, visual inspection of the data allows for an approximation of when the change took place. Based on this supposed change date, the following standard deviations are calculated: BEFORE = 1.41, AFTER = 1.56. The following means are calculated: BEFORE = −2.30, AFTER = 0.87. These results are similar to those in Table II.

Table IV
MODIFIED LEVENE STATISTICS[a] ON DELIVERY WEEK BASIS DATA FOR 600–700 POUND FEEDER CATTLE, 1983–1989

Reporting Station	Heifers Modified Levene Statistic	Steers Modified Levene Statistic
Amarillo Auctions	0.597	0.670
McKinley–Winter Terminal	0.001	1.401
Oklahoma City Auctions	1.286	2.170
Alabama Auctions	1.801	0.020
Arkansas Auctions	0.123	0.811
Clovis Auctions	0.468	0.349
Colorado Auctions	4.836[b]	1.651
Georgia Auctions		0.059
Illinois Direct	2.089	0.106
Iowa Auctions	1.235	0.299
Iowa–S. Minn. Direct	1.954	0.003
Kansas City Terminal	3.247	0.497
Kentucky Direct	3.788	0.223
Lexington Auctions	1.001	0.454
Louisville Auctions		0.374
N. San Joaquin Direct	0.173	2.430
S. St. Joseph Terminal	9.490[b]	3.757
S. St. Paul Terminal	6.753[b]	3.817
Shasta Auctions	3.871	2.163
Sioux City Terminal	0.848	0.272
Sioux Falls Terminal	0.596	0.515
South Carolina Auctions	6.942[b]	1.216
Springfield Auctions	0.678	2.281
Tennessee Auctions	3.158	0.001
Virginia Auctions	4.995[b]	1.189
Washington–Oregon Direct	2.417	0.033
West Fargo Terminal	6.379[b]	0.018

[a]The Levene Statistic is calculated as follows. Using Hsieh and Miller's (1990) notation, let i denote the group ($i = 1..G$), j represents an observation (e.g., x_{ij} would be the jth observation in group i), and n depicts the number of observations in group i. Then,

$$z_{ij} = |x_{ij} - x_i^*|, \text{ where } x_i^* = \Sigma_j x_{ij}/n_i$$

$$W = [\Sigma_i n_i(z_i - z_{..})^2/(G - 1)]/\left[\Sigma_i \Sigma_j (z_{ij} - z_i)^2/\Sigma_i(n_i - 1)\right]$$

where $z_i = \Sigma_j z_{ij}/n_i$ and $z_{..} = \Sigma_i \Sigma_j z_{ij}/\Sigma_k n_k$

The null hypothesis is: $\sigma_1 = \sigma_2 \ldots = \sigma_G$. If x_{ij} is independently and identically distributed, the Levene statistic W is asymptotically a $F(G - 1, \Sigma_i(n_i - 1))$. The modified Levene statistic is calculated by replacing x_i^* with a 5% trimmed mean.

[b]Significant at the 5% level.

hypothesis of (2), changes in mean basis levels (between the two time periods) could introduce a similar bias in this test (see footnote 11).

Table V shows that the null hypothesis (2) is rejected at the 95% level 26 out of 52 times. Considering that this is a much weaker test than the Levene test, the additional significant observations are not surprising. Nevertheless, there are at least three points to note from Table IV: (i) the results are broadly consistent with all previous results; (ii) all eight Z-scores for markets not included in the USFSP reject the null hypothesis, adding to the implication that cash settlement has a strong impact on those markets not in the index[19]; and (iii) the cross-sectional Mann–Whitney Z-score is found to be -2.415. This suggests that the null hypothesis, the "industry" distribution of basis variance BEFORE does not change during AFTER, can be rejected against the one-sided alternative at all levels higher than 79%. This reinforces the conclusion that the settlement change appears to have a favorable influence on basis behavior for the feeder cattle industry, but the individual market results are mixed.

Table VI presents the regression results from the semilogarithmic eq. (4). The Durbin–Watson statistic (1.92) for the model suggests no first order autocorrelation; and, the overall F-statistic (5.03) and adjusted R-square (0.53) reveal, respectively, a significant regression; and, that a fairly high proportion of the cross-sectional variation in basis risk can be explained by this model.

The sign of the cash settlement coefficient is negative and statistically significant, implying that the introduction of cash settlement appears to reduce maturity hedging risk at the industry level. The cross-sectional average BEFORE and AFTER standard deviations are 2.149 and 2.022, respectively, suggesting a decrease of 5.93%. However, following Halvorsen and Palmquist (1980), with a semilogarithmic equation the antilog of the cash settlement coefficient can be calculated and subtracted from one to find the percentage change in standard deviation attributable to the structural change. That is, the relative change (decrease) in hedging risk due to cash settlement is estimated to be 7.25% $(e^{-.075313} - 1)$.

All of the geographic location dummy variables are negative, and 18 of the 22 are statistically significant. The most highly significant coefficients are for markets located in the concentrated areas of feeder cattle production. That is, the standard deviations for basis are lower in the high volume markets.

Sex has a positive sign as hypothesized and is significant at the 0.063 level. Although consistent with previous research, this result seems somewhat paradoxical. The feeder cattle contract is written for steers, not heifers. Therefore, a priori, one would expect heifer basis variability (via cross hedging) to have a smaller reaction to cash settlement than the steer basis. Yet, the results suggest a larger reduction in heifer basis variability than for steers. Kenyon et al. (1991) argued that the "heifer finding" may be a result of other economic variables (e.g., increased demand for breeding stock) and not just the influence of cash settlement. To explore this "heifer finding" in more detail, an additional test is performed. (This test is similar to a Chow multistep procedure.) To extract from locational differences, $log(STD)$ is regressed in the first pass on just the 22 location dummy variables. Then, the first pass residuals are regressed in the second pass on cash settlement, sex, and an interactive cash settlement-sex dummy variable [i.e., $(CS)\,(SEX)$]. The differential intercept and slope coefficients of the second pass are found to be

[19]These results are not consistent with those reported by Kenyon et al. (1991).

Table V
MANN-WHITNEY Z-SCORES ON DELIVERY WEEK BASIS STANDARD DEVIATIONS FOR 600-700 POUND FEEDER CATTLE, 1983-1989

Reporting Station	Heifers Z-Score	Heifers p-value	Steers Z-Score	Steers p-value
Amarillo Auctions	−1.92[a]	0.0274	−0.16	0.4364
McKinley–Winter Terminal	−0.21	0.4168	−0.91	0.1814
Oklahoma City Auctions	−1.92[a]	0.0274	−1.39[b]	0.0823
Alabama Auctions	−1.46[b]	0.0721	−0.48	0.3156
Arkansas Auctions	−0.29	0.3859	−0.75	0.2266
Clovis Auctions	−0.16	0.4364	−1.67[a]	0.0475
Colorado Auctions	−2.25[a]	0.0122	−1.66[a]	0.0485
Georgia Auctions			−0.71	0.2389
Illinois Direct	−3.33[a]	0.0004	−1.77[a]	0.0384
Iowa Auctions	−3.24[a]	0.0006	−1.60[b]	0.0548
Iowa–S. Minn. Direct	−3.76[a]	0.0001	−2.07[a]	0.0192
Kansas City Terminal	−9.92[a]	0.0000	−9.08[a]	0.0000
Kentucky Direct	−1.55[b]	0.0606	−0.75	0.2266
Lexington Auctions	−5.63[a]	0.0000	−3.34[a]	0.0004
Louisville Auctions			−1.80[a]	0.0359
N. San Joaquin Direct	−0.75	0.2266	−1.04	0.1492
S. St. Joseph Terminal	−4.30[a]	0.0000	−3.03[a]	0.0012
S. St. Paul Terminal	−2.79[a]	0.0026	−1.97[a]	0.0244
Shasta Auctions	−0.70	0.2326	−0.75	0.2266
Sioux City Terminal	−1.33[b]	0.0918	−0.48	0.3156
Sioux Falls Terminal	−0.74	0.2296	−0.10	0.4960
South Carolina Auctions	−4.23[a]	0.0000	−3.15[a]	0.0008
Springfield Auctions	−1.30[b]	0.0968	−1.91[a]	0.0281
Tennessee Auctions	−1.21	0.1131	−1.21	0.1341
Virginia Auctions	−3.11[a]	0.0009	−1.80[a]	0.0359
Washington–Oregon Direct	−3.01[a]	0.0013	−0.79	0.2148
West Fargo Terminal	−3.15[a]	0.0008	−0.20	0.4207

[a]Significant at the 5% level.
[b]Significant at the 10% level.

statistically significant at the 5% level. Thus, there may be a gradual increased demand for breeding stock over this period, but there is also strong evidence that suggests a direct shift at the settlement change point in the basis risk for heifers relative to steers. The cause for this is not clear.

IMPLICATIONS AND CONCLUSIONS

Feeder cattle in the U.S. are widely dissimilar in terms of grade, weight, type, breed, and geographic location, making price ranges wide and arbitrage difficult. This set of characteristics is believed to have manifested itself in the 1980s through excessive basis risk, which was claimed to have reduced the effectiveness of the feeder cattle contract

Table VI
RESULTS OF REGRESSING DELIVERY WEEK
BASIS STANDARD DEVIATIONS ON INDEPENDENT VARIABLES, 1983–1989

Variable	Coefficient	Variable	Coefficient
Intercept	1.03[a] (0.08)	Kentucky Direct	−0.23[a] (0.11)
Cash settlement	−0.08[a] (0.03)	Lexington Auctions	−0.06 (0.11)
Sex	0.05[b] (0.03)	Louisville Auctions	−0.17 (0.14)
Alabama Auctions	−0.29[a] (0.11)	McKinley–Winter Terminal	−0.51[a] (0.11)
Amarillo Auctions	−0.62[a] (0.11)	N. San Joaquin Direct	−0.26[a] (0.11)
Arkansas Auctions	−0.36[a] (0.11)	Oklahoma City Auctions	−0.65[a] (0.11)
Clovis Auctions	−0.43[a] (0.11)	S. St. Joseph Terminal	−0.42[a] (0.11)
Colorado Auctions	−0.59[a] (0.11)	Shasta Auctions	−0.01 (0.11)
Georgia Auctions	−0.15 (0.14)	Sioux City Terminal	−0.31[a] (0.11)
Iowa Auctions	−0.25[a] (0.11)	Sioux Falls Terminal	−0.30[a] (0.11)
Iowa–S. Minn. Direct	−0.28[a] (0.11)	Springfield Auctions	−0.21[b] (0.11)
Kansas City Terminal	−0.27[a] (0.11)	Tennessee Auctions	−0.16 (0.11)
		West Fargo Terminal	−0.41[a] (0.11)

The base market is Washington–Oregon Direct. Standard errors are in parentheses.
[a]Significant at the 5% level.
[b]Significant at the 10% level.

as a hedging mechanism.[20] Consequently, a principal motivation for the CME to change the feeder cattle futures contract to cash settlement was to reduce this basis risk.

[20]See CME (1985A, 1985B) and Paul (1985). It is well known that from the beginning, the feeder cattle contract delivery terms were less than perfect. Pricing disputes over grading inaccuracies were common. The pricing controversies may have originated from one, or any combination, of the following: a lack of market information, an insufficient number of cash market sales, geographically dispersed terminal markets, animal quality differences, ineffective grades, changing demand and supply situations, intermarket price differentials, high delivery costs, etc. These factors, individually or grouped, all stem from the fact that U.S. feeder cattle are basically a nonstandardized commodity. Grading and location disputes continually arose over appropriate premiums and discounts for par and nonpar grades, weights, and locations during delivery. It is not clear that the shift to cash settlement can be directly attributed to any one specific factor. Nevertheless, it is clear that between 1979 and 1984, average month-end open interest decreased 80%. (*Monthly Commitment of Traders*)

This study investigates of feeder cattle basis behavior pre- and postcash settlement. It goes beyond previous studies in that it examines many more markets and price reporting stations (27 in total), including some in the USFSP index and some that are not. It also analyzes more contracts pre- and postcash settlement than earlier studies, and tests for changes in both mean and variance. Finally, this study investigates the possibility that usual statistical tests are inappropriate due to the lack of Normality in data distributions. The study therefore includes distribution-free nonparametric tests.

Using weekly average prices during delivery week, weak evidence supports the hypothesis for basis risk reduction. Only 13 out of 52 variance comparisons are significantly smaller after cash settlement. One variance change is statistically larger. However, a sample of tests on these data reveals the majority are not Normally distributed. Retesting the same data using a modified Levene test and distribution-free tests reveal the results are sensitive to the Normality assumption, but broadly consistent with the conclusion that cash settlement has mixed effects on basis variability at the individual markets and favorable effects at the industry level.

The change in basis variability is tested in two of the largest markets, Oklahoma City and Dodge City, utilizing daily data during delivery week. This analysis has the advantage of omitting the "clean-up" sales on the day following the major auction. The variability increases for Dodge City, but not significantly, after cash settlement. The variability is significantly reduced after cash settlement at Oklahoma City. This limited sample makes it difficult to draw distinct conclusions about the influence of "clean-up" sales on basis variability. But, based on these results, it seems safe to say that daily data are preferred to weekly when testing volatility changes. The same general picture emerges, however, with both data sets.

A regression model on weekly data confirms the industry level reduction in basis variability after cash settlement; but, in percentage terms, this reduction is modest. Anecdotal evidence suggests that there is an initial improvement in basis behavior, but that this change is not sustained; and, in fact, worsens during more recent observations. The initial basis reduction may be the result of the aforementioned dissimilarities and high delivery costs which create wide price ranges and which make it difficult to arbitrage among markets. It is possible that with the passage of time settlement prices are doing a poorer job of depicting cash market values thereby augmenting less information flow among markets due to cash settlement and hindering market performance in some individual markets, especially those geographically distant from the central U.S. This phenomenon of an initial improvement of the basis followed by more basis variability is not addressed in this study.

The most interesting aspect of this investigation is the significant increase in the mean of the basis after cash settlement in all 52 cases. In 83% of them, the mean basis is closer to zero than previously. Obviously, futures prices have decreased. Cash settlement should have negligible influence on cash price levels. Apparently, contract specification changes have an influence on the feeder cattle now acceptable under the contract. Cheaper feeder cattle are now being included. Also, the USFSP index is based on the unweighted mid-point of two price ranges, 600–700 pound and 700–800 pound feeder steers from each reporting station. This study analyzes only 600–700 pound animals because of so many missing data points for the 700–800 pound group. When there are no animals in this latter group, a price reporting station presents only the mid-point of the lighter group, which may affect the USFSP index and the basis.

Nevertheless, mean bases closer to zero imply cash and futures are closing together at maturity more tightly than previously. If this average basis that is nearer zero also has a smaller variance, the basis is more predictable.[21] Thus, this study shows modest improvement in basis variability after cash settlement for most individual markets, along with a distinctly smaller basis that is usually nearer zero. That is, adoption of cash settlement has brought improved contract performance for some hedgers, but not all, as results differ across reporting stations. Most of the improvement has come in closer convergence of cash and futures prices.

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[21]Rich (1988) found anecdotal evidence to support this assertion. He found that basis explanation models could explain a larger proportion of basis variability in a postcash settlement time period than in a pre-cash settlement period.

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