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FUTURES TRADING AND MARKET INFORMATION: SOME NEW EVIDENCE

The question of whether futures trading affects cash prices is a long-standing one. Many people believe that futures trading has adverse effects on the behavior of cash prices, such as increasing variability. Economists typically reach the opposite conclusion (for example, Cox, 1976, Gray, 1963, Powers, 1970), C.C. Cox (1976, p. 1236) states that "Spot markets seem to work more efficiently because of futures trading."

The empirical evidence for such statements is based on analyses of cash prices obtained from time periods with futures trading and time periods without futures trading. Several practical problems arise in such analyses. One is the difficulty of holding "other things" constant over the two periods. M.J. Powers (1970), for example, uses a method of eliminating the systematic variation in prices and compares only the random variation in the two periods; presumably this helps to hold "other things" constant.

A second problem, which perhaps is less well recognized, is to define appropriate sample periods. Futures markets do not always work well at their inception (Powers, 1967), and the volume of trading is often small for several years after the opening of new markets. Thus, one can question whether futures trading would have observable effects initially, and hence whether it is appropriate to include these early periods in the sample. Price behavior can now be observed over longer periods with a larger volume of trading than was available to Cox or Powers.

The major objective of this note is to replicate earlier work by Cox (1976) and Powers (1970) using a more recent sample period which does not include the start-up years of the livestock markets, thereby providing new evidence on the old question of the price effects of futures trading. The data and methods are reviewed, and then the empirical results are provided.

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DATA AND METHODS

It is reasonable to analyze cash prices from a time-series viewpoint. Consider

$$P_t = S_t + e_t, \tag{1}$$

where P = price, S = systematic component, e = stochastic component, and t = time, say weeks. The alternative models and procedures that might be used involve varying assumptions about S and e and their interrelationship. The models used by Cox and by Powers assume that the covariance between S_t and e_t is zero and that the e_t (t = 1, 2, . . . , T) have the "classical" properties of zero mean, constant variance, and zero covariances. The principal difference in their procedures is the specification of the S component. The three models used in this paper—those used by Cox and Powers plus one other—differ only in their specification of the systematic component of variation.

In principle, the initiation of a futures market might influence both S and e. For a storable commodity, for example, the existence of a futures market can improve the allocation of inventories and thereby reduce the variance of the seasonal component of prices (Gray, 1963). But the seasonality in prices would not be eliminated, and with the passage of time, the systematic variation in prices obviously will be influenced by many factors other than the addition of a futures market. It is possible, however, that over short periods of time, the ceteris paribus assumption is tenable, though as suggested above, the initial volume of trading in futures may be very small.

In Cox's empirical analysis, current cash price is made a function of lagged prices. This equation is derived from a theoretical model by J.F. Muth (1961), but the theoretical model has a time dimension implicitly based on the production lag in the supply equation while Cox's empirical analysis uses weekly observations. Also, the theoretical model includes a supply of storage equation, while several of the commodities analyzed do not have storage in the ordinary sense of the term (for example, cattle and hogs). Without the storage equation, the autoregression in prices reduces to a random walk model (Cox, 1976, p. 1223). In contrast, spot prices for many agricultural commodities are known to have systematic components, like a seasonal, because of the biological nature of the production process. Indeed, P.A. Samuelson (1965, 1976) argues that, in general, cash prices should have systematic behavior, while the course of futures prices in a perfect market is essentially a random walk.

Cox's procedure is to fit equations with price lags of various lengths, searching for the number of lag terms that minimizes the mean square error of the residuals in the pre-futures period. This price lag length is then used for the regression in the period with futures trading. Cox (1976, p. 1222) expected the following: ". . .when some traders have not acquired the most recent market information, the current market price will equal a linear combination of past prices plus a random-error term. . .The coefficients of past prices alternate in sign starting with a positive coefficient. . .and the earlier the price, the smaller in absolute value is its coefficient. . . . With an increase in market information, the coefficients of past prices. . .all decrease in absolute value. . . (and) . . .an increase in information decreases the variance of the price-forecast error."[1]

Since some doubt exists about the relationship of Cox's theoretical model to his empirical analysis and since the systematic determinants of price change with time, it is probably best to view the autoregressive equation as just one method of partitioning the variance of prices into the deterministic and stochastic components of Equation (1). In this note, the autoregression

$$P_t = \beta_0 + \sum_{i=1}^{n} \beta_i P_{t-i} + e_t \qquad (2)$$

is fitted to weekly cash prices for choice steers and barrows and gilts.[2] The before futures trading sample period is 1955-64 for cattle and 1955-65 for hogs. Both markets had relatively small volumes of trading at their inception, and the period with futures trading is defined (rather arbitrarily) to be January 1969 and after.

Additional insights are obtained by using shorter time periods, and in these instances the autoregressive model may require estimating numerous parameters relative to the number of observations. Another concern is finding the number of lagged terms that minimize the mean square error of the residuals (see next section). Other procedures without these problems can be used to partition the variance of prices into systematic and random components. Two are explored in this note: the variate difference method used by Powers (1970) (see also Tintner, 1952) and a regression model using seasonal dummy variables and linear trend (Jorgenson, 1964).

The variate difference approach is similar to fitting a polynomial of a given degree and then examining the residual variance. The model implicit in this method is Equation (1) without a rigid assumption about the nature of S, other than that the series is smooth. In practice, the method involves computing successive differences and the variances of the resulting series. When the variances stabilize, the resulting variance is taken as a measure of the variability of the random component of the series.

In the regression model, the weekly observations are arbitrarily grouped into 12 seasonal levels (by using zero-one variables) plus linear trend. The assumed nature of the systematic component is rather restrictive, involving 13 parameters. But the specification may be adequate to remove the systematic behavior in prices for each year (since the equation is fitted to the data by year), and it is a simple, straightforward procedure. The model is

[1] The model, as published, apparently contains an erroneous derivation (private communication from Cox). In the corrected result, the theoretical model implies that an increase in information (say, from a futures market) increases the variance of the error term. This seems counterintuitive, and in Cox's empirical results, the variance of the error term decreased.

[2] The spot price quotations are for quality specifications similar to those in the futures contracts and are for the Omaha market. The weekly price is the mid-point of the range of Monday prices reported by the Agricultural Marketing Service, USDA. A Tuesday price is used when Monday is a legal holiday. Thus, there are 52 or 53 observations per year—the numbers of Mondays per year. Data are from USDA microfilm. Since the data were originally collected for another research project, they start in 1955 rather than in 1949, the starting year in Cox's analysis. The number of observations before and with futures is more nearly equal than in Cox. Powers uses "before" and "with" periods of similar length in his study.

TABLE 1.—COEFFICIENTS OF VARIATION FOR AUTOREGRESSIONS

Commodity and period	Standard error (*$/cwt*)	Mean price (*$/cwt*)	Coefficient of variation (*percent*)	Coefficient of variation from Cox[a] (*percent*)
Choice steers				
Before futures, 1955-64	0.475	24.52	1.936	2.737
With futures, 1969-77	1.056	37.83	2.791	2.019
Barrows and gilts				
Before futures, 1955-65	0.630	17.40	3.620	3.585
With futures, 1969-77	1.385	35.23	3.931	3.355

[a]From C.C. Cox (1976), "Futures Trading and Market Information," *Journal of Political Economy*, Vol. 84, December, Table 4, p. 1233. His sample periods are May 1949-November 1964 (before futures) and December 1964-July 1971 (with futures) for cattle and October 1949-February 1966 and March 1966-May 1970 for hogs. The coefficients of variation for this paper are based on Equation (2) as described in the text.

$$P_{ij} = \alpha_0 + \alpha_i + \beta_t + e_{ij}, \tag{3}$$

where α_i is the seasonal effect ($i = 1, 2, \ldots, 11$) and $t = 1, 2, \ldots 52$ (or 53).

In sum, each model assumes that the covariance between the systematic and random components is zero and that the error component has the classical properties. In effect, the systematic variation in the "before" and "with" futures trading periods is taken into account, and the comparison of periods emphasizes the changes in the variability of the error terms.

EMPIRICAL RESULTS

In fitting Equation (2) to the pre-futures prices, 12 lags are required to minimize the mean square error of the residuals for cattle and 19 lags for hogs. In Cox's results, the lags are 8 and 9 weeks, respectively (his data start in 1949 rather than in 1955). A feature of the data used in the analysis reported here is that several local minima appear. While the mean square errors were checked over 24 weeks (lags), one has the uneasy feeling that longer lags might provide a smaller minimum. If the data do reflect the effects of seasonals or longer cycles, the autoregressions on weekly data may not be a satisfactory procedure.

The equations for the two time periods differ both for cattle and hogs, but not in a way that attributes beneficial effects to futures trading. In general, the absolute values of the regression coefficients did not get smaller with the advent of futures trading, and many of the lagged coefficients continue to have t ratios larger than one in the futures trading period. (If the coefficient of a lagged price

TABLE 2.—SELECTED RESULTS BASED ON TIME-SERIES REGRESSIONS FOR CHOICE STEER PRICES

Year	Standard error ($/cwt)	Mean price ($/cwt)	Coefficient of variation (percent)	R^2	Durbin-Watson
1955	.429	22.79	1.88	.98	1.70
1956	.666	22.03	3.02	.95	1.85
1957	.413	23.21	1.78	.96	1.64
1958	.447	27.05	1.65	.90	2.16
1959	.439	27.54	1.59	.92	1.95
1960	.385	25.78	1.49	.93	2.06
1961	.365	24.28	1.50	.95	2.22
1962	.473	27.09	1.75	.93	2.00
1963	.433	23.36	1.85	.92	2.03
1964	.695	22.46	3.09	.89	1.55
1965	.423	25.31	1.67	.94	1.98
1966	.504	25.68	1.96	.89	1.60
1967	.300	25.21	1.19	.94	1.67
1968	.215	26.90	.80	.95	1.89
1969	.663	29.69	2.23	.93	2.03
1970	.550	29.16	1.89	.89	2.20
1971	.443	32.36	1.37	.92	1.59
1972	.789	35.47	2.22	.78	1.81
1973	1.847	44.24	4.17	.86	1.84
1974	.986	41.79	2.36	.94	1.61
1975	1.358	45.48	2.99	.96	1.82
1976	1.317	38.92	3.38	.77	1.70
1977	.760	40.44	1.88	.88	1.32

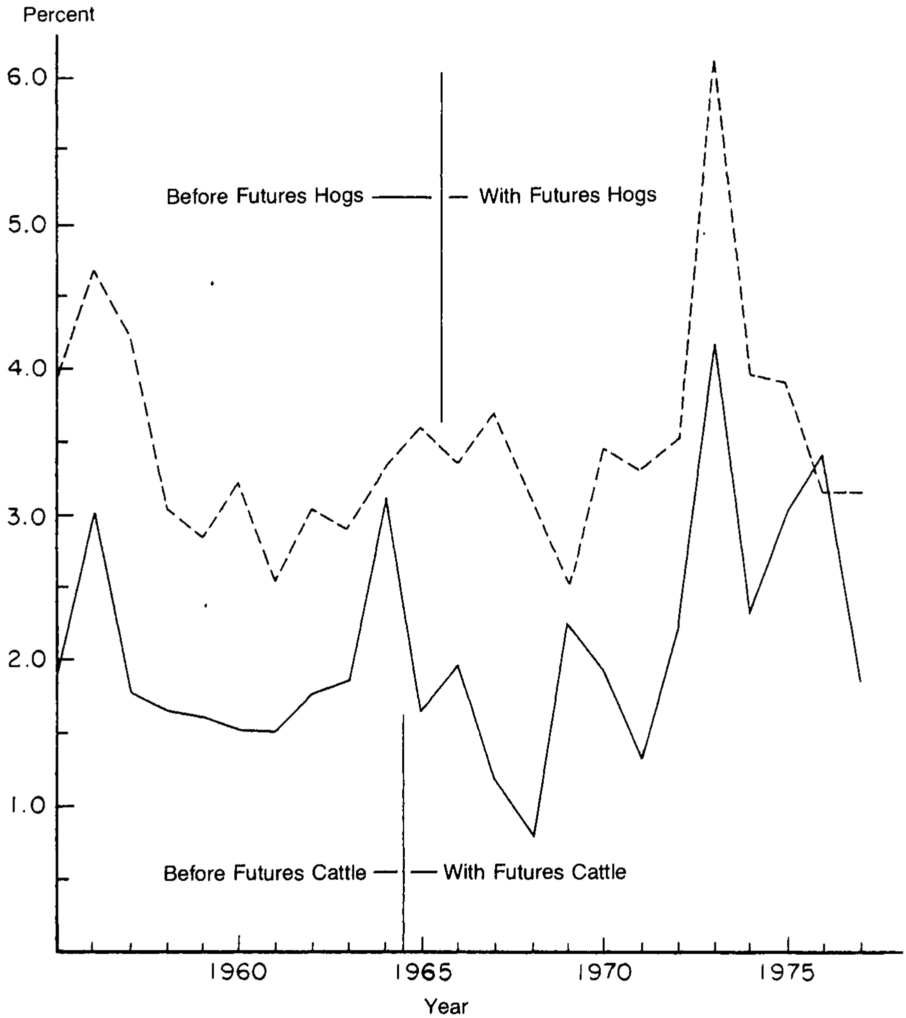
has a t ratio larger than one, then adding the variable will reduce the variance of the residual.)

Coefficients of variation computed from the standard errors of the residuals have increased from the pre-futures to the with-futures period (Table 1). Thus, one might conclude that an active futures market causes the residual variance to increase, which is in contrast to Cox's empirical results.

The results, however, are highly dependent on the sample period used in the analysis. This is illustrated by computing Equation (3) for each year of the sample period. (This model has the advantage of simplicity, and Equation (3) has more degrees of freedom than Equation (2) when (2) contains more than 12 lags.) Selected results for cattle and hogs are reported in Tables 2 and 3, and the coefficients of variation for each year are plotted in Chart 1.

Most equations have large R^2's and Durbin-Watson statistics near two. This implies that Equation (3) was reasonably successful in removing systematic behavior and leaving residuals with little or no first-order autocorrelation. The

CHART 1.—COEFFICIENTS OF VARIATION BASED ON REGRESSION RESIDUALS



standard errors and coefficients of variation computed from the resulting error terms clearly differ from year to year.[3] The coefficients for cattle are exceptionally small in 1968 and 1971, two years in Cox's sample of with-futures trading observations. The variation of the error terms both for cattle and hogs was

[3] The standard errors for the estimated residuals for each year are divided by the average price for the year to obtain the coefficients of variation. A similar procedure is used in computing the coefficients of variation for the variate difference method.

TABLE 3.—SELECTED RESULTS BASED ON TIME-SERIES REGRESSIONS FOR BARROW AND GILT PRICES

Year	Standard error (\$/cwt)	Mean price (\$/cwt)	Coefficient of variation	R^2	Durbin-Watson
1955	0.467	16.38	3.95	.95	1.83
1956	0.727	15.50	4.69	.88	1.59
1957	0.803	18.89	4.25	.82	1.38
1958	0.640	20.85	3.07	.90	1.75
1959	0.433	15.09	2.87	.95	2.44
1960	0.523	16.39	3.19	.91	1.70
1961	0.443	17.52	2.53	.81	1.78
1962	0.527	17.26	3.05	.81	1.97
1963	0.464	15.92	2.91	.93	2.29
1964	0.528	15.94	3.31	.84	1.82
1965	0.787	21.87	3.60	.97	2.03
1966	0.826	24.52	3.37	.93	2.23
1967	0.751	20.26	3.71	.89	1.99
1968	0.609	19.73	3.09	.76	1.61
1969	0.617	24.47	2.52	.96	2.12
1970	0.809	23.46	3.45	.97	1.60
1971	0.643	19.36	3.32	.88	1.63
1972	0.959	27.55	3.48	.89	1.57
1973	2.531	41.56	6.09	.88	1.53
1974	1.459	36.92	3.95	.93	1.64
1975	1.962	50.33	3.90	.96	1.15
1976	1.386	44.93	3.08	.97	1.87
1977	1.303	42.15	3.09	.87	1.59

unusually large in 1973, a year influenced by the existence and release of price controls.[4]

The variate difference method also gives variances and coefficients of variation (not shown) that differ substantially from year to year. The standard errors

[4] The first eight months of 1973 were characterized by high livestock prices and by price controls. Farmers apparently expected prices to rise still further after price controls were dropped and held animals for later marketing. Consequently, the market was glutted with overweight animals and prices dropped rapidly as marketings increased after the termination of price controls. Choice steer prices in Omaha were \$56.75 per cwt on the 33rd Monday of 1973 and by the 39th Monday had dropped to \$38.50. For barrows and gilts, the drop was from \$62.37 to \$41.62 on the same dates. The effects of price controls and erroneous expectations seem sufficiently unusual to consider dropping 1973 from the comparisons of periods.

TABLE 4.—COEFFICIENTS OF VARIATION BY MODEL, SELECTED TIME PERIODS
(Percent)

	Choice steers			Barrows and gilts		
Time period	Auto regression[a]	Trend regression	Variate difference	Auto regression	Trend regression	Variate difference
Before futures						
1955-64	1.94	1.96	1.19	3.62	3.40	2.25
With futures						
1965-71	—	1.59	0.98	—	3.24[b]	1.89[b]
1969-77	2.79	2.50	1.36	3.93	3.65	1.97
1973 excluded	—	2.00	1.27	—	3.36	1.87

[a]Auto regression = Equation (2) in text.
Trend regression = Equation (3) in text.
For the trend and variate methods, coefficients are simple averages of annual coefficients.
[b]Time period for hog prices is 1966-71.

computed by the variate difference approach, however, are consistently smaller than those from Equations (2) or (3). The variances computed for the 1955-64 period for cattle are similar to the variance (.086) reported by Powers (1970, p. 462) for 1961-64.

The coefficients of variation are summarized by taking simple averages for selected time periods (Table 4). If the pre-futures years are compared with the years immediately after futures trading started, the coefficients of variation decline, the conclusion reached by Cox and by Powers. If the comparison is made with 1969-77, the coefficients of variation increase (except for the one computed by the variate difference method for hogs). But, if 1973 is excluded from the with-futures years, then the coefficients of variation are approximately equal in the two periods, increasing slightly for cattle and decreasing slightly for hogs. On balance, the existence of futures trading appears not to have a measurable effect on the variation of the stochastic component of cattle and hog prices, and the conclusion does not depend on the model used.

While it seems logical (at least to this author) that the variability of the random component of spot prices might be reduced by the introduction of an active futures market, the quality of information in a spot market may be very good before futures trading starts. The USDA, for example, provides information on market prices and factors related to supply for cattle and hogs, and the introduction of a futures market may improve information very little. If the marginal effect of futures trading on information is small and if random events of quite different magnitudes occur with the passage of time, then measures of variability may be sensitive to the sample period used and the price effects of futures trading, if any, will be exceedingly difficult to ascertain.

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