Time-Varying Risk Premium or Informational Inefficiency?
Further Evidence in Agricultural Futures Markets

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Recent research has provided mixed results regarding the presence of a time-varying risk premium in agricultural futures markets. In this paper we test for the presence of a time-varying risk premium and market efficiency focusing on the properties of the underlying data. Specifically, we examine the same markets and period used by McKenzie and Holt (2002) and extend the analysis through 2004. Our results show that accounting for the structural break in the early seventies plays a key role in the findings. In contrast to McKenzie and Holt, we find no evidence of time-varying risk premium in the four commodities analyzed. The corn market appears to be (weak form) efficient. Hogs, live cattle, and soybean meal futures contracts show evidence of inefficiency, which suggests an inability of these markets to incorporate all available information in the futures prices. Our results identify the importance of careful examination of the data as failure to do so can lead to inappropriate conclusions.

Key words: risk premium, market efficiency, structural change, stationarity

Introduction

The presence of a risk premium in agricultural commodity futures markets is controversial, and difficult to determine. For a specific market, the difficulty emerges because of the relationship between risk premium and market efficiency. Under the risk premium hypothesis (Keynes 1930), risk-averse short hedgers pay speculators to bear spot price risk, resulting in futures prices that systematically fall below the futures spot prices and rise over the life of the contract to meet spot prices at maturity. However, this bias in futures prices may also reflect the inability of the market to incorporate available information at a specific time to forecast subsequent spot prices at maturity. Failure to account for these competing hypotheses can make statistical inference regarding the presence of risk premium and market efficiency problematic, and can cloud our understanding of how and why markets operate.

As evidence of the difficulty in disentangling these hypotheses we consider two recent studies of agricultural futures markets. Beck (1994) tests market efficiency allowing for risk premia and concludes that the inability of futures prices to forecast cash prices is determined by informational inefficiencies rather than by the presence of a risk premium in many commodities (i.e., cattle, hogs, orange juice, corn, copper, and cocoa). However, she assumes that the risk premium is constant. Engle, Lilien and Robins (1987) point out that risk premium may be time-varying because the amount that risk-averse hedgers pay to speculators is a function of the uncertainty in the underlying asset which varies over time. Also, the proportion of hedgers and speculators can change over time, contributing to the time-varying behavior of the risk premium. Building on Beck (1994) and Engle, Lilien and Robins (1987), McKenzie and Holt (2002) test for risk premium and market efficiency using an ARCH-in-mean error correction framework. They conclude that hogs and cattle markets contain a time-varying risk premium, and that the presence of the risk premium can affect the assessment of market efficiency.

Testing these hypotheses involves several empirical difficulties as the characteristics of the data (i.e. stationarity, structural breaks, seasonality) may influence the selection of procedures and findings. Stationarity tests performed over short periods of time may capture trends generated by irrational market agents causing temporal price movements away from their mean values. When longer periods of time are considered, these trends may disappear due to the behavior of more rational traders who force prices back to their mean prices. Using long periods of time to test for market efficiency may therefore reflect more accurately the behavior of futures markets in the equilibrium. However, stationarity tests performed over a long period of time, as in McKenzie and Holt, can be influenced by structural changes which introduce a trend that can bias unit root tests towards non-rejection of the null hypothesis.

Market efficiency tests may also be sensitive to the seasonal behavior of the underlying commodity. Beck (1993) shows that in the hogs market a time-varying risk premium may be confused with seasonal effects if the model does not account for the latter. Further, Newbold et al. (1999b) argue that futures markets inefficiency may be due to the presence of seasonal patterns and efficiency tests should account for this effect.

The purpose of this paper is to reassess both the presence of a time-varying risk premium and market efficiency taking into account the time series properties of the underlying data. Specifically, we examine the same agricultural commodities used by McKenzie and Holt (corn, live cattle, hogs, and soybean meal) during their period of analysis (1959-1995) and then extend the analysis incorporating more recent data to assess the robustness of the findings. Our motivation for re-examining these issues with these data is straightforward. Structural breaks observed in many agricultural commodity prices in the early 1970s can affect stationarity tests (Wang and Tomek, 2004), the procedures used in the analysis, and also may influence inferences regarding the presence of time-varying risk premium and market efficiency.

The next section provides a brief review of the efforts made in the literature to test the risk premium hypothesis. Sections three and four describe the data used and provide the results of the tests for structural change. Since the methods used to test for efficiency and risk premium are driven by the properties of the data (i.e. stationarity), we also perform unit root tests. Based on the stationarity results, in section five we present the econometric models and test the market efficient and risk premium hypotheses, and then offer conclusions in section six.

Literature review

The literature on the presence of risk premium has produced mixed and contentious findings.[1] Early studies by Houthakker (1957) and Rockwell (1967) computed speculator's returns as a measure of risk premium using different datasets and found opposite results. Kolb (1992) tested for the existence of a risk premium using three different measures and a large dataset (1959-1988) and found normal backwardation behavior for meats but no evidence of a risk premium for grains. Deaves and Krinsky (1995) extended Kolb's work by adding more years of data and confirm his results for livestock commodities. Hartzmark (1987) used actual trading histories of individual futures traders obtained from the Commodity Futures Trading Commission (CFTC) to study five markets (oats, wheat, pork bellies, live cattle, and feeder cattle). He found that speculators, on average, do not earn positive profits, whereas the opposite holds for hedgers. Based on these results, he rejected the normal backwardation hypothesis for these markets.

Beck (1994) tested the efficiency and risk premium hypotheses separately using the cointegration and error correction technique developed by Engle and Granger. Her findings attribute the bias in orange juice, corn, copper, and cocoa markets to the inability of futures markets to incorporate all available information and not to the presence of risk premium. However, Beck's conclusions hold for those markets characterized by a constant risk premium.

The notion of a time varying risk premium was explored by Engle, Lilien, and Robins (1987) for the term structure of interest rates. Market participants' perceptions of the uncertainty changes over time, and therefore so does the amount that they are willing to pay (receive) to hold a net short (long) position in the futures market. The underlying market uncertainty is represented by the conditional variance of future prices. In order to incorporate this effect on the futures prices, they propose the ARCH-M model, where the risk of holding a contract is compensated by a higher expected spot price at $t+1$, which in an efficient market equals the current futures price. As the risk increases, the risk premium increases and therefore the deviation of futures prices from future spot prices is higher. Using this ARCH-M procedure and a rational expectations model, Beck (1993) tests the efficient market hypothesis (for a two-month forecast horizon) allowing for the presence of a time-varying risk premium in hog, orange juice, soybean and live cattle markets. In her study, only the soybean market shows a significant time-varying risk premium.

McKenzie and Holt (2002) extended Beck's (1994) work by testing market efficiency and unbiasedness separately and allowing the risk premium to vary (linearly and nonlinearly) over time. They analyze four commodities using a long dataset: live cattle (1965-2000), hogs (1966-2000), corn (1959-2000), and soybean meal (1959-2000). Using three different unit root tests on the whole dataset (i.e., Dickey Fuller, weighted symmetric, and Phillips-Perron) they found that all series contained one unit root and therefore they use cointegrating techniques to test market efficiency. They found that all of these markets are efficient and unbiased in the long-run. In the short-run, futures markets show a different behavior. Corn futures prices appear to be biased and inefficient and soybean meal futures prices are efficient and unbiased. The hog market contains a short-run time varying risk premium, but futures prices are still efficient predictors of future spot prices. Live cattle markets also show a short-run time varying risk premium and futures prices appear to have some informational short-run inefficiencies.

Data

The presence of a time varying risk premium and market efficiency is tested in the hogs, live cattle, corn, and soybean meal markets using futures prices from the CME and CBOT for a two-month forecast horizon.[2] The datasets were constructed to ensure equally spaced observations and to eliminate overlapping data corresponding to different forecasting horizons. However, corn and soybean meal contracts do not expire every other month and therefore a dataset with equal forecast horizons and spacing can not be constructed. Based on Newbold et al. (1999a) and following McKenzie and Holt (2002), we kept the equal spacing by using a longer forecast horizon in December. Contract months and forecast horizon lengths are summarized in table1. Spot prices are futures prices at expiration and futures contracts were pooled for the analysis.[3]

Structural break and unit root tests

The models used by McKenzie and Holt are based on their finding of non-stationary prices. However, non-stationarity may be caused by a structural change in the data. Figure 1 shows the commodity prices used in their study. Commodity prices increased dramatically in the early seventies due to changes in US international trade patterns and the impact of the oil crisis.

To support the notion that a structural change in these markets influenced the stationarity tests, we test for the presence of a structural change using non-parametric methods, which do not assume any particular distribution of the underlying process. In order to test differences in the mean before and after the structural break, we perform the Fligner-Policello test, which is sensitive to differences in the median and does not assume equal variance before and after the break. Changes in the variance of prices after the structural break are assessed using the Miller jackknife test, which does not assume an equal median in the two periods.[4] The structural break period for each commodity used in these tests is indicated in figure 1. As expected, the non-parametric structural break tests results shown in table 3 suggest that the data generating processes before and after the early seventies are different.

In light of our structural break findings, we test for stationarity using three different periods. Period 1 is identical to that used by Mckenzie and Holt and extends from the sixties to the mid nineties. Period 2 starts after the structural break and ends on the same date as Period 1. Period 3 also begins after the structural break but extends to mid 2004.[5] The exact dates and length of each period are shown in table 2.

Stationarity of the data in natural logs was assessed using the Augmented Dickey-Fuller (ADF) test and the Innovational Outlier (IO) test. The ADF test was performed taking into account three different models (with constant, constant and trend and with no constant and no trend) including lags ranging from 1 to 8. The appropriate model was selected using the AIC criterion to determine the number of lags (Enders, 2003). The IO test is similar to the ADF test but allows for structural breaks in the time series (Perron, 1990). We use the IO test for the live cattle market with a structural break in 1978.[6]

Results of the ADF and IO test are reported in table 4. The ADF test for prices during Period 1 shows that all commodities are non-stationary. In contrast, results of the ADF and IO test for the period after the break show that all series are stationary. Apparently, testing for unit roots without accounting for the structural break introduces an upward trend that causes bias towards non-rejection of the null hypothesis of unit root. These results coincide with those found by Wang and Tomek (2004) for selected agricultural commodities.

Market efficiency and risk premium tests

Given the strong evidence of the structural break and unit root test results, we focus our analysis on prices after the structural break to assess the presence of a time varying risk premium and market efficiency in a stationary framework.[7] Here, consistent with the literature discussed above, we consider a weak form of efficiency and assume costless information.

The classical approach to test for market efficiency is to estimate the following equation,

$$S_t = \beta_0 + \beta_1 F_{t-1} + \varepsilon_t \qquad \varepsilon_t \sim N(0, \sigma_t^2) \qquad (1)$$

where S_t are natural logs of spot prices at expiration, F_{t-1} are natural logs of futures prices two months prior to expiration, β_0 and β_1 are estimated coefficients, and ε_t is the unconditional error term. If futures prices are unbiased, then the null hypotheses β_0=0 and β_1=1 will not be rejected. However, efficient markets may reject the above joint hypothesis due to the presence of a risk premium (Beck, 1994).[8]

Another source of inefficiency may be due to seasonality. This effect is incorporated in model (1) using dummy variables. We conducted ANOVA analysis and Tukey test for differences in mean between contract months to detect the seasonal pattern of each commodity. Only hog prices show significant differences between months. This price behavior is incorporated in equation (1) as follows,

$$S_t = \beta_0 + \beta_1 F_{t-1} + \beta_2 D_1 + \beta_3 D_2 + \varepsilon_t \qquad \varepsilon_t \sim N(0, \sigma_t^2) \qquad (2)$$

where D_1 and D_2 are seasonal dummies for hogs. D_1=1 for August, D_2=1 for June and August, $D_1=D_2=0$ otherwise, $\beta_2=0$ in period 3, and $\beta_3=0$ in period 2. For all other commodities $\beta_2=\beta_3=0$.

Market informational inefficiency is due to the presence of information at time *t-1* that contributes to predict S_t and is not contained in F_{t-1}. Any omitted variable in equation (1) contained in the residual ε_t will yield correlation between residuals across time and therefore testing for autocorrelation is a natural test for market efficiency. We use the Breusch-Godfrey LM test where rejection of the null hypothesis of no autocorrelation indicates that the market is informationally inefficient.[9] We assume that the relevant information to predict S_t not contained in F_{t-1} is in past spot and futures prices and use the BIC to select the optimal lagged structure. Equation (3) incorporates the source of inefficiency in the mean,

$$S_t = \beta_0 + \beta_1 F_{t-1} + \beta_2 D_1 + \beta_3 D_2 + \sum_{i=1}^{I} \gamma_i S_{t-i} + \sum_{j=2}^{J} \delta_j F_{t-j} + \varepsilon_t \qquad \varepsilon_t \sim N(0, \sigma_t^2) \qquad (3)$$

However, even when the market shows evidence of informational inefficiency a time-varying risk premium may still exist.[10] The presence of a time-varying risk premium can be tested using Engle, Lilien, and Robins (1987) ARCH-M model. This test is performed using equation (3); hence, the LM ARCH test is performed conditional to the autocorrelation structure.[11] If we reject the null hypothesis of no ARCH effect, then we test whether this behavior is a significant factor in spot price forecasts, reflecting the presence of a time-varying risk premium. This is done using the ARCH-M model (4), where the time-varying risk premium is assessed by testing the significance of the variance term introduced in the mean equation. We also extend equation (4) to allow for GARCH (1,1) effects, as shown in (5),

$$S_t = \beta_0 + \beta_1 F_{t-1} + \beta_2 D_1 + \beta_3 D_2 + \beta_4 \sigma_t^2 + \sum_{i=1}^{I} \gamma_i S_{t-i} + \sum_{j=2}^{J} \delta_j F_{t-j} + \varepsilon_t \qquad \varepsilon_t \backslash \varepsilon_{t\text{-q}} \sim \mathrm{N}(0,\ \sigma_t^2)$$

$$\sigma_t^2 = \alpha_0 + \sum_{q=1}^{Q} \alpha_q \varepsilon_{t-q}^2 \qquad \alpha_0,\ \alpha_q > 0,\ \sum_{q=1}^{Q} \alpha_q < 1 \qquad (4)$$

$$S_t = \beta_0 + \beta_1 F_{t-1} + \beta_2 D_1 + \beta_3 D_2 + \beta_4 \sigma_t^2 + \sum_{i=1}^{I} \gamma_i S_{t-i} + \sum_{j=2}^{J} \delta_j F_{t-j} + \varepsilon_t \qquad \varepsilon_t \backslash \varepsilon_{t\text{-q}} \sim \mathrm{N}(0,\ \sigma_t^2)$$

$$\sigma_t^2 = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \lambda_1 \sigma_{t-1}^2 \qquad \alpha_0,\ \alpha_1,\ \lambda_1 > 0,\ \alpha_1 + \lambda_1 < 1 \qquad (5)$$

where ε_t is the error term conditional on the amount of volatility observed in recent periods (ε_{t-q}), σ_t^2 is the variance of ε_t, and a_0 , a_q and λ_1 are estimated coefficients of the variance.[12] A positive sign of β_4 implies normal backwardation, that is, the market is dominated by short hedgers who pay a risk premium to long speculators to bear the risk. In contrast, a negative sign implies contango, where the market is dominated by long hedgers paying a risk premium to short speculators.

Table 5 shows the Breusch-Godfrey LM test for one to six lags performed using model (2). There is no evidence of autocorrelation for corn in either Periods 2 and 3. For the other commodities there is strong evidence of inefficiency, although the inefficiency seems to have decreased in the hog market for the more recent period. The lag structure for the mean equation and the LM test for ARCH effects are reported in table 6. For the corn market, there is no evidence of changes in volatility. These results are similar to McKenzie and Holt's finding of no risk premium in the corn market.[13] The LM ARCH test shows evidence of time-varying price volatility at the 1% significant level in the hog and live cattle markets for both time periods analyzed. However, for soybean meal there is no evidence of ARCH effects.

Table 7 shows the test of time-varying risk premium for those commodities with significant LM ARCH test results (hogs and live cattle). In both markets the conditional variance has no significant effect in the mean equation, contrasting with McKenzie and Holt's time-varying risk premium in hog and live cattle markets.

Next, we tested the efficient market hypothesis for corn. According to the results in table 8 and the previous finding of no autocorrelation, we can conclude that the efficient market hypothesis holds. Moreover, based on the LM ARCH test result, we also conclude that there is no risk premium in this market. The conclusion of market efficiency coincides with the findings of Zulauf et al. (1999) for the December contract.

As identified, the other markets do not pass the efficiency test (i.e. LM test for autocorrelation). The results are strong and hold for both periods. These findings may not be surprising given the evidence of inefficiency previously reported in the literature. Leuthold and Hartmann (1979) and Leuthold et al. (1989) found inefficiencies in the hog market using different methods and periods. Garcia et al. (1988), and Martin and Garcia (1981) found alternative models that can outperform cattle futures prices. Finally, Rausser and Carter (1983) found that multivariate and ARIMA models outperform the futures markets in the soybean meal market.

Conclusions

The presence of risk premium and the efficient market hypothesis have been studied extensively and results are controversial. Disentangling market informational inefficiencies and risk premium is not straightforward and conclusions about the ability of futures prices to predict subsequent spot prices may be influenced by underlying characteristics of the data.

Here, we investigated the presence of a time-varying risk premium and market efficiency focusing on the time series characteristics of the data. Results show that the properties of the data play a significant role in model specification of efficiency tests. Failure to account for the structural break in the early 1970s introduces a trend into unit root tests, which leads McKenzie and Holt to estimate an error correction model to investigate the efficiency and risk premium hypotheses. Our results show that accounting for the structural break in the early seventies also plays a key role in the findings. We find no evidence of a time-varying risk premium for the four commodities analyzed which contrasts with McKenzie and Holt findings regarding the presence of a time-varying risk premium in hogs and live cattle markets, but is consistent with Beck (1993). These results hold for both the data used by McKenzie and Holt and the expanded data to the most current period.

Hogs, live cattle and soybean meal markets show evidence of inefficiencies, but the corn market appears to be (weak form) efficient. These efficiency findings are consistent with previous research, and along with the work by Beck (1994) who identified the absence of a constant risk premium, suggesting that the bias in these futures markets may reflect their inability to incorporate information effectively rather than the presence of risk premium. Further, our findings identify the importance of careful examination of the characteristics of the data as failure to do so can obstruct our understanding of how and why markets behave.

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Endnotes

[1] The portfolio approach is one way to test the risk premium hypothesis. Dusak (1973) uses the Capital Asset Pricing Model (CAPM) and found no systematic risk in futures contracts for wheat, corn, and soybeans data. However, Carter, Rausser, and Schmitz (1983) analyzed the same commodities and found evidence of systematic risk of futures supporting normal backwardation. Later, Marcus (1984) and Baxter, Conine, and Tamarkin (1985) changed the market index used by Carter, Rausser and Schmitz and found no evidence of risk premium for the same commodities. Still based on the financial literature, Ehrhardt, Jordan, and Walkling (1987) applied the Arbitrage Pricing Theory to the same commodities and rejected the risk premium hypothesis. So (1987) applied the random coefficient method to test futures price behavior of the same commodities for different periods and find no significant risk premium in the overall period (1953-1976) and most subperiods. Bjornson and Carter (1997) found a significant time-varying conditional risk premium in agricultural commodities returns using the asset pricing model.

[2] Data obtained from the Commodity Research Bureau (CRB). The analysis was also performed for a four-month horizon and the results, which are available, do not differ from those presented.

[3] We use futures prices because spot prices for the same grade and location as futures prices are usually not available.

[4] See Hollander and Wolfe, 1999, p.135 (Fligner-Policello) and p. 158 (Miller).

[5] We focus on in-sample analysis through mid 2004 to assess the robustness of our findings and to increase the number of observations for the ARCH type processes, which can be difficult to estimate in smaller samples.

[6] The structural break for cattle can be seen in 1978 in figure 1, panel b, reflecting higher cattle prices due to prior liquidation of the breeding inventories which was influenced by higher feed prices. The IO test was performed using the following equation: $Y_t = \alpha + \varphi D_L + \eta D_P + \gamma y_{t-1} + \Sigma_{i=1}^{r} \zeta_i \Delta y_{t-1}$, where D_P = pulse dummy (1 for t=1978:2, 0 otherwise) and D_L = level dummy (1 for t≥1978:2, 0 otherwise).

[7] In the presence of non-stationary prices differencing and allowing for cointegration between cash and futures is more appropriate.

[8] Rejection of null hypothesis does not necessarily mean that the market is inefficient in the strictest sense. In a weak form context, efficiency is most completely assessed by examining the returns and costs from alternative strategies.

[9] In the Breusch-Godfrey LM test of autocorrelation the OLS residual of equation (2) is regressed on its lagged values and the independent variables of equation (2). The test statistic, the sample size times the R^2 of the regression, is distributed as a chi-square with degrees of freedom equal to the number of lags.

[10] The observed autocorrelation may be due to ARCH effects and not to autocorrelation per se. We examined this possibility by performing the time-varying risk premium test first prior including the additional lagged variables. Our conclusions regarding the presence of a time-varying risk premium do not change.

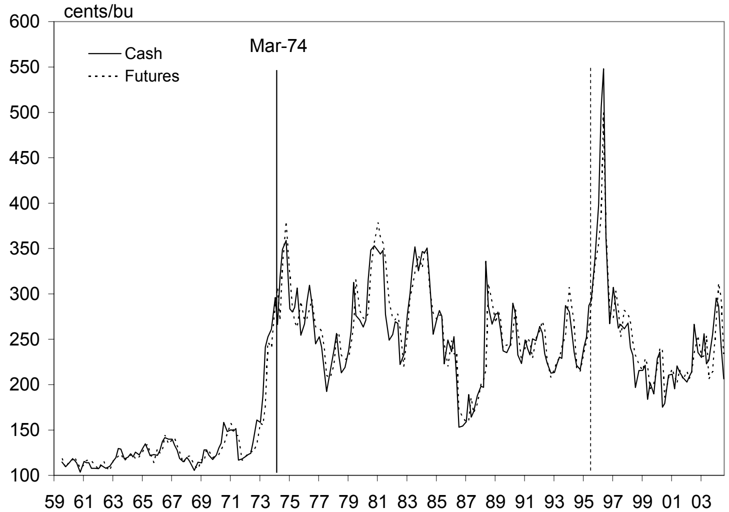
[11] In the LM ARCH test the squared OLS residual of equation (3) is regressed on an intercept and its lagged values. The test statistic, the sample size times the R^2 of the regression, is distributed as a chi-square with degrees of freedom equal to the number of lags.

[12] The stationarity ($\alpha_1 + \lambda_1 < 1$) and nonnegativity ($\alpha_0, \alpha_1, \lambda_1 > 0$) conditions of the GARCH (1,1) model follows from Bollerslev (1986) and Nelson and Cao (1992) respectively. For the ARCH (Q) model these conditions are stated in Hamilton (1994, p.658-659)

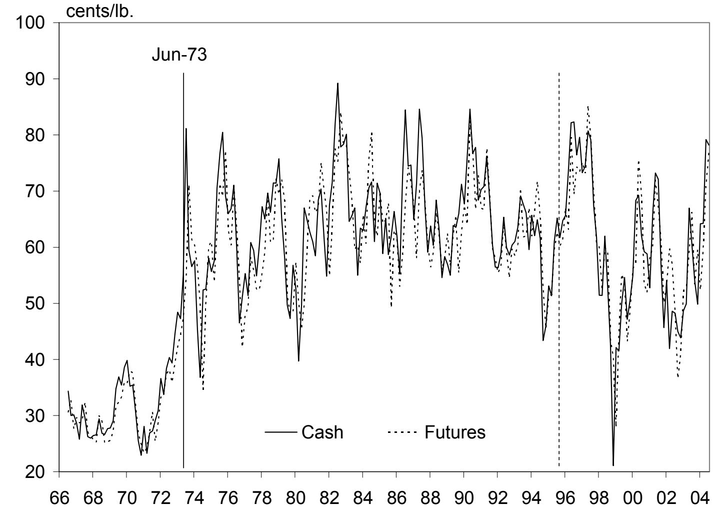
[13] McKenzie and Holts's conclusions focus on the presence of short-run risk premium. The unit root tests performed in this paper show no evidence of prices departing from equilibrium and therefore the long-run and short-run effects coincide.

Figure 1. Commodity Futures Prices at Expiration (Cash) and 2 Months Prior to Expiration (Futures).

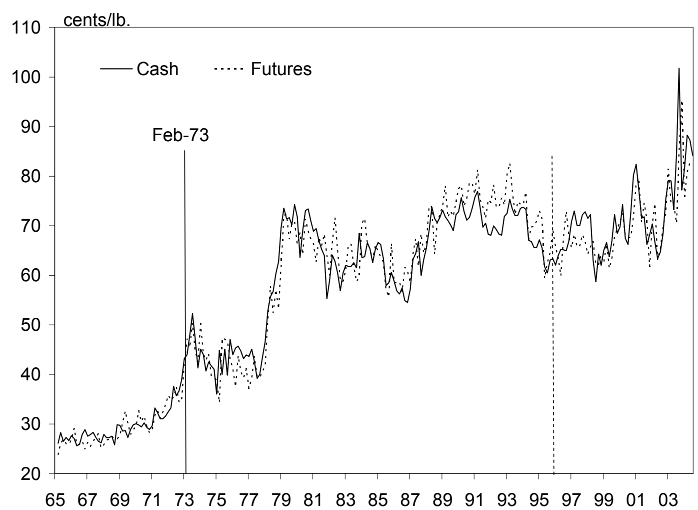
a) Corn (CBOT, Sep 1959 – Sep 2004)



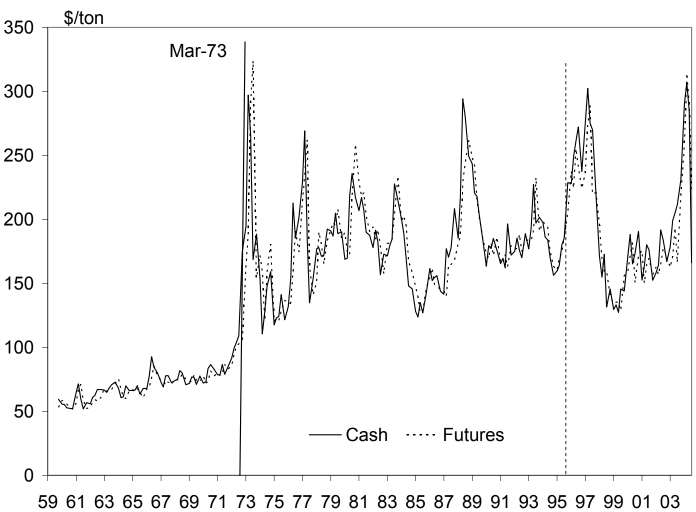
b) Hogs (CME, Aug 1966 – Oct 2004)



c) Live Cattle (CME, Apr 1965 - Oct 2004)



d) Soybean Meal (CBOT, Dec 1959 – Sep 2004)



The solid vertical line represents the structural break and the dotted vertical line represents the end of period used by McKenzieand Holt (2002)

Table 1. Data Characteristics

Commodity	Contract months	Forecast horizon (trading days before expiration)
Corn	Mar, May, Jul, Sep, Dec	40 (Dec: 60)
Hogs	Feb, Apr, Jun, Aug, Oct, Dec	40
Live cattle	Feb, Apr, Jun, Aug, Oct, Dec	34
Soybean meal	Mar, May, Jul, Sep, Dec	40 (Dec: 60)

Table 2. Time Periods

Commodity	Period 1	N_1	Period 2	N_2	Period 3	N_3
Corn	1959:09 – 1995:09	181	1974:03 – 1995:09	109	1974:03 – 2004:09	154
Hogs	1966:08 – 1995:10	176	1973:06 – 1995:10	135	1973:06 – 2004:08	188
Live cattle	1965:04 – 1995:10	184	1973:02 – 1995:10	137	1973:02 – 2004:08	190
Soybean meal	1959:12 – 1995:09	179	1973:03 – 1995:09	114	1973:03 – 2004:09	159

N is the number of observations in each period

Table 3. Non-parametric Structural Break Tests for Period 1

Commodity	Structural break	m	n	FP $\hat{U}$-stat	FP p-value	MI Q-stat	MI p-value
Corn	1974:3	72	109	-58.53	< 0.001	-1.75	0.040
Hogs	1973:6	41	135	-177.25	< 0.001	-2.77	< 0.001
Live Cattle	1973:2	47	137	-837.75	< 0.001	-7.58	< 0.001
Soybean Meal	1973:3	65	114	-77.12	< 0.001	-1.57	0.058

m = # of observations before the break
n = # of observations after the break
FP: Fligner-Policello test for differences in mean between two periods.
MI: Miller jackknife test for differences in variance between two periods.
For large m and n both U and Q converge to a standard normal distribution.

Table 4. Unit Root Tests

	Price	Period 1			Period 2			Period 3		
		Lags	Model	τ-stat	Lags	Model	τ-stat	Lags	Model	τ-stat
Corn	SPOT	1	CNT	-1.97	2	CNT	-3.21*	2	CNT	-3.58*
	FUT	1	NCNT	0.62	3	CNT	-3.05*	3	CNT	-3.38*
Hogs	SPOT	8	CNT	-2.20	6	CNT	-6.81*	7	CNT	-5.79*
	FUT	8	CNT	-2.04	8	CNT	-4.01*	7	CNT	-5.72*
Live	SPOT	1	NCNT	1.13	8	CM	-5.49*	6	CM	-4.93*
cattle	FUT	7	NCNT	1.01	6	CM	-5.15*	6	CM	-5.08*
Soy	SPOT	1	CT	-3.21	6	CNT	-3.68*	6	CNT	-4.71*
meal	FUT	1	CT	-2.97	8	CNT	-3.26*	8	CNT	-3.91*

CM: changing mean, NCNT: no constant and no trend, CNT: constant and no trend, and CT: constant and trend
τ-stat corresponds to IO test for live cattle after the structural break and to ADF test for the rest of the series. Critical values for the IO test are available in Perron (1990), table 4; for the ADF test see, for example, Enders (2003), table A.
* denotes stationary series at the 5% significant level.

Table 5. LM Test of Autocorrelation

$$S_t = \beta_0 + \beta_1 F_{t-1} + \beta_2 D_1 + \beta_3 D_2 + \varepsilon_t$$

Commodity	Period	(1)	(2)	(3)	(4)	(5)	(6)
Corn	2	0.92	0.53	0.59	0.75	0.76	0.85
	3	0.60	0.54	0.73	0.74	0.84	0.77
Hogs	2	0.00	0.01	0.01	0.01	0.02	0.03
	3	0.15	0.11	0.06	0.03	0.03	0.05
Live cattle	2	0.01	0.00	0.00	0.01	0.01	0.00
	3	0.00	0.00	0.00	0.00	0.00	0.00
Soybean meal	2	0.01	0.04	0.00	0.00	0.01	0.01
	3	0.02	0.07	0.00	0.00	0.00	0.01

Numbers in parenthesis are the number of lagged residuals, p, used in the test. P-values of the χ^2 distribution with p degrees of freedom are reported.

Table 6. Conditional LM Test for ARCH

$$S_t = \beta_0 + \beta_1 F_{t-1} + \beta_2 D_1 + \beta_3 D_2 + \sum_{i=1}^{I} \gamma_i S_{t-i} + \sum_{j=2}^{J} \delta_j F_{t-j} + \varepsilon_t$$

Commodity	Period	I	J	(1)	(2)	(3)	(4)	(5)	(6)
Corn	2	-	-	0.88	0.71	0.78	0.81	0.86	0.93
	3	-	-	0.84	0.67	0.81	0.85	0.91	0.96
Hogs	2	3	2	0.00	0.00	0.00	0.09	0.89	0.92
	3	2	-	0.01	0.03	0.07	0.12	0.24	0.41
Live cattle	2	6	-	0.00	0.00	0.00	0.01	0.01	0.02
	3	1	-	0.00	0.00	0.00	0.00	0.00	0.00
Soybean meal	2	4	-	0.14	0.27	0.16	0.21	0.14	0.08
	3	4	-	0.32	0.43	0.33	0.41	0.10	0.06

Numbers in parenthesis are the number of lagged squared residuals, p, used in the test. P-values of the χ^2 distribution with p degrees of freedom are reported.

Table 7. Risk Premium Test in Hog and Live Cattle Markets

$$S_t = \beta_0 + \beta_1 F_{t-1} + \beta_2 D_1 + \beta_3 D_2 + \beta_4 \sigma_t^2 + \sum_{i=1}^{I} \gamma_i S_{t-i} + \sum_{j=2}^{J} \delta_j F_{t-j} + \varepsilon_t$$

$$\sigma_t^2 = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \lambda_1 \sigma_{t-1}^2$$

	Period 2				Period 3		
	Coef.	SE			Coef.	SE	
Hogs							
β_0	0.870	0.357	**	β_0	0.681	0.254	*
β_1	0.641	0.110	*	β_1	0.432	0.090	*
β_2	0.054	0.030	***	β_3	0.053	0.017	*
β_4	-0.126	2.607		β_4	0.775	1.139	
γ_1	0.319	0.137	**	γ_1	0.287	0.110	*
γ_2	0.107	0.174		γ_2	0.114	0.065	***
γ_3	0.108	0.093		α_0	0.005	0.001	*
δ_2	-0.386	0.152	**	α_1	0.656	0.130	*
α_0	0.006	0.001	*	λ_1	0.160	0.096	***
α_1	0.308	0.177	***				
Live cattle							
β_0	0.436	0.100	*	β_0	0.205	0.075	*
β_1	0.421	0.060	*	β_1	0.279	0.053	*
β_4	2.208	2.579		β_4	-0.537	1.845	
γ_1	0.511	0.115	*	γ_1	0.672	0.051	*
γ_2	0.136	0.101		α_0	0.002	0.000	*
γ_3	-0.164	0.094	***	α_1	0.499	0.162	*
γ_4	0.245	0.092	*				
γ_5	-0.030	0.102					
γ_6	-0.226	0.078	*				
α_0	0.001	0.000	*				
α_1	0.498	0.192					
λ_1	0.062	0.166					

Level of significance: 1% (*), 5% (**), and 10% (***)

Table 8. Efficiency Test in the Corn Market

$S_t = \beta_0 + \beta_1 F_{t-1} + \varepsilon_t$

Period	Restriction		
	β_0=0	β_1=1	β_0=0, β_1=1
2	0.68	0.75	1.26
3	3.54 ***	3.67 ***	2.35

F-statistics are reported. Level of significance: 1% (*), 5% (**), and 10% (***)