Market Efficiency in Agricultural Futures Markets

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1. Introduction

The concepts of market efficiency and unbiasedness have been difficult to distinguish empirically. Market efficiency implies that futures prices will equal expected future spot prices plus or minus a possibly time-varying risk premia, while futures prices will be unbiased forecasters of future spot prices only if markets are both efficient and have no risk premia. The hypothesis that futures prices provide unbiased forecasts of spot prices is thus a joint hypothesis of market efficiency and risk neutrality. The issue is further complicated by a time dimension, whereby markets may be efficient and unbiased in the long-run, but may experience short-run inefficiencies. The objective here is to empirically test the two separate hypotheses of market efficiency and unbiasedness in both the long and short-term.

This paper uses both the two-stage Engle-Granger and Johansen cointegration procedures to test for long run market efficiency and unbiasedness by allowing for the possible existence of a constant risk premia. The short-run price dynamics are analyzed using an error-correction model within an ARCH framework. Univariate GARCH in mean (GARCH-M), and ARCH in mean (ARCH-M) models are estimated. This modeling approach allows us to test for short-term market efficiency and unbiasedness, taking into account the possible existence of both a constant, and a time varying risk premia. The model is analyzed from the perspective of broiler producers, whereby the efficiency of relevant substitute goods and inputs are tested.

Arbitrage dictates that if futures market speculators are risk neutral, then the current futures price of a commodity must equal the expected future spot price of the same commodity at contract maturity. This condition is illustrated formally as follows:

(1) $\qquad E_{t-1}S_t = F_{t-1}$

Where $E_{t-1}S_t$ is the expectation of the future spot price formed in period $t-1$, and F_{t-1} is the futures price with contract maturity in period t.

This simple specification of market efficiency gives rise to a natural model specification useful for testing market efficiency. If we assume that expectations are rational, so that: $S_t = E_{t-1}\left(S_t \mid \Omega_{t-1}\right) + u_t$, where Ω_{t-1} denotes the information set available in period $t-1$, u_t is a rational expectations error, and u_t is orthogonal to all elements in Ω_{t-1}, including lagged forecast errors, we may write (1) as:

(2) $S_t = \alpha + \delta F_{t-1} + u_t$

Moreover, the joint assumption of market efficiency and unbiasedness is implied by the additional restrictions that $\alpha = 0$, $\delta = 1$. We therefore have a natural test for market efficiency and unbiasedness in the context of the simple linear regression model, equation (2).

It should be noted, however, that three separate conclusions might be inferred from the rejection of the null hypothesis in such tests: (1) the market may indeed be inefficient; (2) a constant risk premium may exist which makes market forecasts biased but possibly efficient; (3) it may be that some possibly time varying risk premium is inherent to the market, thus preventing futures prices in isolation from providing unbiased forecasts of future spot prices. It is the objective of this paper to differentiate empirically between these theoretically plausible conclusions.

Previous studies, which have tested for market efficiency using the model in equation (2), include Tomek and Gray (1970), Kofi (1973), Leuthold (1975), Leuthold and Hartman (1979), and Martin and Garcia (1981). The results have been mixed with evidence that nonstorable livestock futures prices provide biased forecasts of future spot prices. Leuthold found this result for live cattle futures contracts with maturities greater than or equal to three months.

It has been argued in the literature that the appropriate tests for market efficiency and unbiasedness are dependent upon the stationarity of the data. If the price series are non-stationary, hypothesis tests based on equation (2) will give biased results. Regressing a non-stationary variable, which can only be made stationary by differencing on a deterministic trend, generally leads to the problem of a spurious regression, involving invalid inferences based on t and F tests. In such cases the researcher could falsely conclude that a relationship exists between two unrelated non-stationary series. Previous research has attempted to circumvent the stationarity problem by estimating equation (2) in first differences e.g. Hansen and Hodrick (1980). However, such simple stationary-inducing transformations are misspecified if the futures and spot prices are cointegrated. If the two series are in fact cointegrated, an error-correction model (ECM) should be estimated.

2. Methodology.

In this paper, the estimation problems associated with non-stationary data, are addressed by using cointegration techniques and error correction models. Initially, cointegration techniques are used to test for market efficiency in five agricultural commodities futures markets while allowing for a constant risk premia. This follows the approach of Beck (1994), who similarly tested for efficiency in various commodities futures markets

If spot and futures prices are both non-stationary and require differencing to make them stationary, then in general most linear combinations of the two series will also be non-stationary. However, there may exist a cointegrating vector, which makes a specific linear combination of the two series stationary. For example if (3) below is a stationary series, α and δ are the cointegrating terms and the regression $S_t = \alpha + \delta F_{t-1} + u_t$ as in (2) above is the cointegrating or equilibrium regression.

$$(3) \qquad u_t = S_t - \alpha - \delta F_{t-1}$$

It implies that S_t and F_{t-1} cannot move too far apart from each other despite the fact that they are both non-stationary. Cointegration between the two series is a necessary but not a sufficient condition for market efficiency. Spot and futures prices are determined by the same fundamentals and so efficiency implies that they cannot move too far apart. However, cointegration does not rule out short run market inefficiencies, whereby past information can improve future market forecasts of future spot prices.

A time-series model of a cointegrated series can be rewritten in an error correction form as in Granger (1986). Such a transformation renders the series stationary, and allows for normal hypothesis testing. The error correction model is shown in equation (4) below.

$$(4) \qquad \Delta S_t = -\rho u_{t-1} + \beta \Delta F_{t-1} + \sum_{i=2}^{m} \beta_i \Delta F_{t-i} + \sum_{j=1}^{k} \psi_j \Delta S_{t-j} + v_t$$

Where Δ is defined as the change or difference in a variable from one period to the next, u_t is the error correction term, and v_t is a stationary series.

Cointegration implies $\rho > 0$ because spot price changes respond to deviations from the long-run equilibrium equation (2). Market efficiency implies that $\rho = 1$, $\rho \delta = \beta \neq 0$ and $\beta_i = \psi_i = 0$. The coefficient β for the current change in the futures price is non-zero because new information, which also affects the futures price, affects the future change in the spot price. The additional restrictions that $\rho \delta = \beta$, $\rho = 1$, and $\beta_i = \psi_i = 0$ can be seen by rewriting equation (4) as shown below, where $(S_{t-1} - \alpha - \delta F_{t-2})$ is substituted for u_{t-1} from equation (2).

$$(5) \qquad S_t = (1 - \rho)S_{t-1} + \beta F_{t-1} + (\rho\delta - \beta)F_{t-2} + \rho\alpha + \sum_{i=2}^{m} \beta_i \Delta F_{t-i} + \sum_{j=1}^{k} \psi_j \Delta S_{t-j} + v_t$$

If the above restrictions did not hold then past future and spot prices would contain relevant information not completely incorporated into current future prices, which could be used to predict S_t. The efficient markets hypothesis

states that all past information should already be incorporated into the current futures price, and therefore it should have no effect on the future spot price.

Beck (1994) shows that efficiency tests based on equation (4) and the above restrictions allow for the existence of a constant risk premium. This is because unlike equation (2), such tests do not impose the assumption that $\alpha = 0$ and $\delta = 1$. Thus if it is assumed that risk premia are constant and not time varying in nature, the two hypotheses of unbiasedness and efficiency can be tested for separately. Beck (1994) performs such tests on various commodities futures markets. The unbiasedness hypothesis is examined using an error correction model as in equation (4) and testing the restrictions that $\rho = 1$, $\beta = 1$, and $B_i = \Psi_i = 0$. The less restrictive market efficiency hypothesis tests the restrictions that $\rho = 1$, $\beta = \delta$, and $B_i = \Psi_i = 0$. Beck rejects the null hypothesis of unbiasedness, but cannot reject the null hypothesis of market efficiency at the 5% significance level for live cattle futures prices at a two month forecast horizon. Neither null hypothesis can be rejected at the 5% significance level for corn futures prices at a two month forecast horizon.

The assumption that $\alpha = 0$ and $\delta = 1$, can be tested using the Johansen multivariate cointegration procedure. This approach estimates Likelihood Ratio tests for restrictions on the parameters of the cointegrating regression. The Engle Granger two step cointegration procedure cannot be used to test these restrictions, as the test procedure does not have well defined limiting distributions. If the hypothesis that $\alpha = 0$ and $\delta = 1$ cannot be rejected, long run market efficiency and unbiasedness may be inferred. In this case equation (3) reduces to:

(6) $\qquad u_t = S_t - \delta F_{t-1}$

The error correction model can now be estimated with a constant term as in (7) below:

(7) $\qquad \Delta S_t = \lambda - \rho u_{t-1} + \beta \Delta F_{t-1} + \sum_{i=2}^{m} \beta_i \Delta F_{t-i} + \sum_{j=1}^{k} \psi_j \Delta S_{t-j} + v_t$

The market efficiency hypothesis can now be analyzed by testing the restrictions that $\rho = 1$, $\beta = 1$, and $B_i = \Psi_i = 0$.

In this context any short run market inefficiencies cannot be due to long run market bias, and the two concepts of unbiasedness and market efficiency may be regarded as synonymous.

Finally, the above efficiency tests are also estimated using GARCH-M and ARCH-M models to take into account a possibly short-run time varying risk premia. Commodity prices have exhibited extensive volatility over the sample period analyzed. ARCH models provide a useful way to parameterize the time-varying conditional variances

observed in commodity market variables. In this case equation (7) can be rewritten as (8) to include ARCH terms and the time varying risk premia term h_t, which is the conditional standard deviation of the change in spot prices. This was the approach taken by Engle, Lilien and Robins (1987) when modeling the term structure of interest rates, and by Domowitz and Hakkio (1985), who used this model to test for time varying risk premia within foreign exchange markets. Similarly, Beck (1993) used ARCH-M models to test for time varying risk premia in agricultural futures markets. Thus following this approach the risk premium is hypothesized to be a function of the conditional standard deviation of the change in the spot price, or the forecast error. Theoretical justification for using this model in agricultural markets is based on the intertemporal hedging theory first advanced by Keynes (1930). Short hedgers, such as producers, sell futures contracts at a price below the expected future spot price to avoid price risk. The difference between the two prices, the risk premium, compensates purchasers of futures contracts for bearing the spot price risk. An increase in spot price risk, as measured by the conditional variance of spot prices, should increase the risk premium in the relevant futures market.

$$(8) \qquad \Delta S_t = \lambda - \rho u_{t-1} + \beta \Delta F_{t-1} + \sum_{i=2}^{m} \beta_i \Delta F_{t-i} + \sum_{j=1}^{k} \psi_j \Delta S_{t-j} + \theta h_t + v_t$$

Where $v_t = h_t.e_t$, and $h_t^2 = w + \sum_{i=1}^{q} \alpha_i v_{t-i}^2 + \sum_{j=1}^{p} \gamma_j h_{t-j}^2$ where $e_t \sim IN(0,1)$

3. Empirical Tests of Market Efficiency

Data from five different agricultural commodities markets: live cattle, live hogs, corn, soybean meal and iced broilers were analyzed. The data included both futures and spot prices for these markets over the period 1966 to 1995. A two month forecast horizon was analyzed in each case to match the production cycle for broilers. The relative pricing efficiency of these markets is relevant in broiler production and hedging decisions, because efficiency in these markets infers that the futures prices may be taken as proxies for the expected future spot prices of the respective commodities. Rational-expectations commodity models have shown the importance of using the expected prices of relevant substitute and input variables in determining supply. Effective hedging strategies designed to reduce input price risk for broiler producers is also dependent upon the efficiency of the corn and soybean meal futures markets.

Futures contracts for all of these commodities occur at roughly bimonthly intervals throughout the year. To avoid the problem of introducing a moving average process into the residuals, spot price observations were taken on a bimonthly basis to match the future contract time intervals. The average futures prices during the contract expiration month for cattle and hogs were taken to represent spot prices. However, bimonthly average cash prices were used for

corn, soybean meal and wholesale broiler spot prices. It was assumed that average input prices over the whole production period, and substitute good prices at the end of the production period would be of most relevance in production decisions. Corn prices are Chicago number two yellow, while soybean meal prices are Decatur 44% protein for the period August 1966 to December 1992 and Decatur 48% protein for the period January 1993 to October 1995. Broiler prices are nine city wholesale prices, ready to cook for 1968-1977, and twelve city composite wholesale prices, ready to cook for 1978-1982. For months when a relevant futures contract did not exist, the closest following contract month was used as a substitute. All data were converted into natural logs and multiplied by one hundred

An initial consideration must be to test the data for non-stationarity and to determine if the data generating process is difference or trend stationary. It is also important to establish the number of unit roots that a series contains when testing for cointegration. For two non-stationary series to be cointegrated they must be integrated of the same order. Dickey-Pantula tests indicated that all the price series contain at most one unit root. The stationarity of the series was then examined to determine whether they contain a single unit root. Three different unit root tests were used, the augmented Dickey-Fuller test, the weighted symmetric test, and the Phillips-Perron test. The estimated model, included a constant term, and a trend term. The optimal number of augmenting lags for the model was determined by using the Akaike information criterion 2 (AIC2).

All three tests indicated that each of the series contain one unit root and therefore may be regarded as difference stationary. Given that the spot and futures prices are integrated of the same degree, cointegration techniques can be used to determine if a long-run relationship exists between the spot and futures prices. The two-stage Engle-Granger cointegration tests, based on Augmented Dickey-Fuller statistics, indicated that the residuals of OLS regressions on equation (2) were stationary. The Akaike information criterion was used to infer the optimal number of lags. These results show that the futures price series for each commodity appear to be cointegrated with their respective spot prices. The OLS coefficient estimates from equation (2) appear to be close to the (0,1) restriction, which satisfies long run unbiasedness. The results are not sensitive to the reversal of the variables in equation (2).

The issue of cointegration was further explored by using the Johansen cointegration procedure. This is a multivariate approach based on deriving maximum likelihood estimates of the cointegrating regression. The VAR specification was estimated using from one to four lags. The AIC criterion was used to choose the optimal number of lags for each commodity. The maximal eigenvalues and trace statistics presented in Table 1 indicate that the null hypothesis of no cointegration is rejected at the 10% significance level for all commodities. On the other hand, the null

of one cointegrating regression cannot be rejected for all commodities. Thus both the Engle-Granger and the Johansen techniques support the hypothesis that the spot and futures prices for all of the commodities are cointegrated.

The terms α and δ shown in Table 1 are the normalized intercept and futures price coefficients in the cointegrating regressions. Once again these coefficients appear to be close to the (0,1) restriction of unbiasedness. Formal testing of the long run unbiasedness hypothesis was conducted using Johansen likelihood ratio tests on the implied (0,1) restrictions of α and δ. The results, which are shown in Table 2 indicate that the individual null hypotheses that $\alpha = 0$, and $\delta = 1$ cannot be rejected at the 5% level for any of the commodities. However, the joint null hypothesis that $\alpha = 0$ and $\delta = 1$ is rejected at the 5% level for iced broilers. These results suggest that futures prices for live cattle, live hogs, corn and soybean meal provide unbiased forecasts of future spot prices in the long run. However, the evidence suggests that iced broiler futures prices may have been biased predictors of two-month ahead average wholesale broiler prices.

Although the above tests, with the exception of iced broiler futures, provide support for the hypothesis of long run market efficiency and unbiasedness, all of the commodity futures markets may exhibit short-run inefficiencies. To test for short-run inefficiencies the ECM models discussed in the methodology section were estimated, and the results are shown in Table 3. Given the long run efficiency results, the short-run dynamics for live cattle, live hogs, corn and soybean meal were modeled using equation (7) above. F-tests were computed to test the restrictions, $\rho = 1$, $\beta = 1$ and $\beta_i = \psi_i = 0$, imposed by market efficiency. The possible inherent bias discovered in the iced broiler futures market may be due to a constant risk premium, and hence Beck's approach was followed to analyze the short-run dynamics of this market using equation (4) above. In this case the error correction term was computed as $\left(S_{t-1} - 9.380 - 0.991\beta_{t-2}\right)$, the lagged residual from cointegrating regression in equation (2). An F-test was used to test the restrictions $\rho = 1$, $\beta = 0.991$ and $\beta_i = \psi_i = 0$. The models were estimated with zero to seven lags of ΔS_{t-1} and ΔF_{t-1}. Lags with significant coefficients were retained (see Engle and Granger 1987). Residual diagnostics reveal no evidence of serial correlation. ARCH effects were detected in each of the regressions with the exception of iced broilers.

The magnitude of the error correction term coefficient indicates the speed of adjustment of any disequilibrium toward the long run equilibrium state of market efficiency and unbiasedness. The coefficients on the error correction terms are significant in all regressions. This is consistent with the result that the futures and spot markets are

cointegrated. The live cattle futures market appears to have the lowest adjustment rate, with an error correction term coefficient of 0.2213. Plots of the error correction terms suggest that substantial departures from long-run market efficiency took place during the high inflationary period of the mid 1970's. This may be regarded as supporting evidence that futures markets provide inadequate forecasts of future spot prices during periods of unexpectedly high and volatile rises in the general price level.

The results in Table 3 show that all of the commodity markets fail the test of short-run market efficiency and unbiasedness at the 5% level. However, the reasons behind the failures vary among the markets. Short–run efficiency and unbiasedness cannot be rejected for the live hog futures market at the 1% significance level. Rejection at the 5% level can be attributed to significant lags in the hog spot market and to slow adjustment to previous periods forecast errors as shown by the relatively low coefficient value for the error correction term. However, the individual hypothesis that the coefficient on the first differenced futures price should be equal to one cannot be rejected. A similar result occurs for corn futures, with the coefficient on first differenced futures prices being insignificantly different from one. Live cattle and soybean meal on the other hand, fail the short-run efficiency tests because of significant lags on spot prices, slow adjustment to long run equilibrium and a rejection of the hypothesis that the coefficient on the first difference futures price is unity.

ECM for all of the markets with the exception of iced broilers, where no ARCH effects were found, were also modeled within an ARCH-M and GARCH-M framework. These models were estimated in an attempt to see if the short-run market inefficiencies could be the result of time varying risk premia. The estimated values for θ, which represents the coefficient on the time varying risk parameter is shown for the various commodities in Table 3. Estimation indicated that the most suitable model for live cattle, live hogs and corn was an ARCH-M(1) process. A GARCH-M(1) process was modeled for soybean meal. It can be seen that all of the markets once again reject short-run efficiency. However, in each market the coefficients on the first differenced futures price and the error correction term are closer to unity than in the previous model. For live cattle, the coefficient on the first differenced spot price is not significantly different from unity, in contrast to the results reported in Table 3. In addition, spot prices with longer lags replace the previously significant lagged spot prices, suggesting that if ARCH terms are taken into account, recent spot price changes no longer contain relevant information in determining future spot price changes. For the live hogs market the efficiency hypothesis is now rejected at both the 5% and the 1% significance levels. However, both the coefficient on the error correction term and on the first differenced futures price are now not significantly different

from unity. Rejection of market efficiency is due to a significant coefficient on the sixth lag of the first differenced futures price. It is questionable as to whether this lag has economic significance. The coefficient on the time varying risk parameter is only marginally insignificant at the 5% significance level for the live cattle and corn markets. It is insignificant for the live hogs and soybean markets. The risk coefficient has a positive sign for the live cattle market, indicating a positive time varying risk premia existed for the sample period. This would support the theory that livestock futures markets, with a preponderance of short hedgers, would tend to see futures prices biased downwards. The risk coefficient in the corn market has a negative sign. However, if following Domowitz and Hakkio (1985), the time varying risk premia is interpreted as the summation of the risk parameter, θ, and the constant term λ, the risk premia could then change sign over time. This is because the constant term in the corn regression is both significant and has the opposite sign to θ. Plots of the conditional standard deviation, for each market, over the sample period, contain noticeable spikes in the mid 1970's. This suggests there was considerable uncertainty at this time, because of high unexpected inflation, and that this uncertainty may have been reflected in a higher risk premium during the period.

4. Conclusions

Results indicate that the live cattle, live hogs, corn and soybean meal futures markets are both efficient and unbiased in the long-run. These results suggest that in the long-run, risk premia do not exist in these markets. There is evidence that the iced broiler futures market is biased in the long-run. Short-run inefficiencies were found to exist in each market in the short-run. The iced broiler market exhibited short-run inefficiency even allowing for the possible existence of a constant risk premia. There was no evidence of a time varying risk premia in the live hogs and soybean markets to explain these short-run inefficiencies. There was some indication that a short-run time varying risk premia may have existed for live cattle and corn over the sample period.

TABLE 1
Johansen Test for Cointegration

Commodity	α	δ	λ_{trace} $k=0$	λ_{trace} $k\leq 1$	λ_{max} $k=0$	λ_{max} $k=1$
Cattle (4)	4.089	1.001	43.17* (17.79)	7.21 (7.50)	35.97* (10.29)	7.21 (7.50)
Hogs (4)	-1.273	1.009	30.97* (17.79)	6.31 (7.50)	24.67* (10.29)	6.31 (7.50)
Corn (4)	15.010	0.971	32.67* (17.79)	4.56 (7.50)	28.10* (10.29)	4.56 (7.50)
Soy Meal(1)	8.087	0.982	268.29* (17.79)	6.79 (7.50)	261.50* (10.29)	6.79 (7.50)
Broilers (1)	9.380	0.991	41.84* (17.79)	3.79 (7.50)	38.05* (10.29)	3.79 (7.50)

The Trace test was used to test the null hypothesis that the number of cointegrating vectors is less than or equal to k, where k is 0 or 1.

* Indicates that the null hypothesis is rejected at the 10% level.

The critical values at the 10% level are taken from λ_{trace} and λ_{max} tables, Osterwald-Lenum (1992), and are shown in parentheses below the test statistics.

The cointegrating vector is normalized with respect to S_t.

The lag length chosen by the AIC criteria is shown in parentheses after the relevant commodity

TABLE 2
Johansen Tests of Restrictions on the Cointegrating regressions

Commodity	$\alpha=0$	p value	$\delta=1$	p value	$\alpha=0, \delta=1$	p value
Live Cattle	0.40	0.53	0.48	0.49	0.92	0.63
Live Hogs	0.01	0.90	0.09	0.76	4.60	0.10
Corn	1.85	0.17	1.98	0.16	2.62	0.27
Soy Meal	1.52	0.22	1.92	0.17	6.02	0.05
Iced Broil	0.28	0.59	0.04	0.85	16.35	0.00*

The null hypothesis are shown in the tables: $\alpha=0$; $\delta=1$; $\alpha=0, \delta=1$. A Likelihood Ratio test statistic for the various restrictions is shown and has a χ^2 distribution with the degrees of freedom equal to the number of restrictions placed on the parameters.

* Indicates rejection of the null hypothesis at the 5% level.

** Indicates rejection of the null hypothesis at the 1% level.

TABLE 3

Error Correction Models with Johansen Error Correction Term

$$\Delta S_t = \lambda - \rho u_{t-1} + \beta \Delta F_{t-1} + \sum_{i=2}^{m} \beta_i \Delta F_{t-i} + \sum_{j=1}^{k} \psi_j \Delta S_{t-j} + v_t$$

Parameters	Live Cattle	Live Hogs	Corn	Soy Meal	Iced Broil
λ	1.0497	1.8349	-0.0520	-0.0015	0.4180
	(0.5817)	(0.7988)	(0.5647)	(0.8642)	(0.7772)
ρ	-0.2213	-0.6991	-0.5795	-0.4038	-0.5599
	(0.1650)	(0.1156)	(0.1217)	(0.1511)	(0.1244)
β	0.4921	0.7998	0.8262	0.5939	0.8382
	(0.1714)	(0.1057)	(0.1018)	(0.1445)	(0.1022)
β_2	-	-	-	-0.2000	-
				(0.0696)	
ψ_1	-0.3663	-	-	-	-
	(0.1444)				
ψ_2	-0.2873	-0.1334	-	-	-
	(0.0829)	(0.0643)			
ψ_3	-	-	-	0.1462	-
				(0.0710)	
ψ_4	-0.1670	-	-	-	-
	(0.0754)				
ψ_6	-	-	0.1507	-	-
			(0.0646)		
ψ_7	-	-0.1459	-	-	-
		(0.0673)			
R^2	0.205	0.350	0.322	0.171	0.4917
Adjusted R^2	0.181	0.334	0.310	0.152	0.4772
$H_0 : \beta = 1$	-2.96325**	-1.8932	-1.708199	-2.80951**	-1.4958
	(0.003)	(0.058)	(0.088)	(0.005)	(0.135)
$H_0 : \rho = -1$	4.7204**	2.6037**	3.453992**	3.94645**	3.5369**
	(0.000)	(0.009)	(0.001)	(0.000)	(0.000)
F Test	5.4663**	3.0747*	6.3309**	6.7006**	6.2634**
	(0.0001)	(0.0179)	(0.0004)	(0.0001)	(0.0031)

Standard errors are shown in parentheses. *Indicates reject the null at the 5% level. ** Indicates reject the null at the 1% level. T-statistics and p-values shown for individual hypotheses $\rho = -1$ and $\beta = 1$.

P-values for joint hypothesis: $\rho = -1$, $\beta = 1$, all other parameters = 0 are shown in parentheses below the F-statistic.

Note: with respect to iced broilers test statistics are based on the $H_0 : \rho = -1, \beta = 0.991$, all other parameters=0.

TABLE 4

Error Correction Models with Johansen Error Correction Term AndGARCH-M(1,1) - ARCH-M(1) Processes

$$\Delta S_t = \lambda - \rho u_{t-1} + \beta \Delta F_{t-1} + \sum_{i=2}^{m} \beta_i \Delta F_{t-i} + \sum_{j=1}^{k} \psi_j \Delta S_{t-j} + \theta h_t + v_t$$

Where $v_t = h_t . e_t$,

and $h_t^2 = w + \sum_{i=1}^{q} \alpha_i v_{t-i}^2 + \sum_{j=1}^{p} \gamma_j h_{t-j}^2$ where $e_t \sim IN(0,1)$

Parameters	Live Cattle	Live Hogs	Corn	Soy Meal
λ	-1.3410	3.1417	14.4336	1.3569
	(1.2999)	(3.3180)	(7.3603)	(1.5440)
ρ	-0.5367	-0.8134	-0.6223	-0.4530
	(0.1302)	(0.1155)	(0.1133)	(0.1175)
β	0.7582	1.0008	0.8496	0.6753
	(0.1239)	(0.0970)	(0.1017)	(0.0994)
β_6	-	-0.2020	-	-
		(0.0609)		
ψ_1	-0.2681	-	-	-
	(0.0923)			
ψ_5	0.2498	-	-	-
	(0.0609)			
ψ_6	0.2214	-	0.1427	-
	(0.0633)		(0.0552)	
w	21.0018	59.0901	43.3045	16.6481
	(3.6339)	(9.3279)	(4.0381)	(3.9762)
α_1	0.6869	0.3948	0.1464	0.7318
	(0.1813)	(0.1456)	(0.0949)	(0.2314)
γ_1	-	-	-	0.3153
				(0.1106)
θ	0.3906	-0.0966	-2.0640	-0.1727
	(0.2099)	(0.3687)	(1.0946)	(0.2251)
Log Likelihood	-554.509	-612.143	-569.483	-624.349
R^2	0.0649	0.3310	0.3170	0.1336
$H_0 : \beta = 1$	-1.9519	0.0083	-1.4789	-3.2666**
	(0.052)	(0.999)	(0.140)	(0.000)
$H_0 : \rho = -1$	3.5584**	1.6156	3.334**	4.6553**
	(0.000)	(0.111)	(0.001)	(0.000)
$H_0 : \theta = 0$	1.861	-0.262	-1.886	-0.767
	(0.0627)	(0.7934)	(0.0594)	(0.4431)
F Test	9.4606**	5.5551**	6.5278**	10.8372**
	(0.0001)	(0.0012)	(0.0003)	(0.0001)

T statistics and P-values are shown in parentheses for the individual hypotheses: ρ = -1; β =1; $\theta = 0$

All other parameters = 0; P-values for the joint null hypotheses above are shown in parentheses below the F statistic

P-values for the joint null hypotheses above are shown in parentheses below the F statistic.

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