View metadata, citation and similar papers at core.ac.uk
brought to you by CORE
provided by Research Papers in Economics

Efficiency of Forest Commodity Futures Markets

Dequan He
Ph.D. Candidate
Department of Economics
North Carolina State University

Matt Holt
Professor
Departments of Agricultural and Resource
Economics and Economics
North Carolina State University

Selected paper prepared for presentation at the American Agricultural Economics Association Annual Meeting, Denver, Colorado, August 1-4, 2004

Copyright 2004 by Dequan He and Matt Holt. All rights reserved. Readers may make verbatim copies of this document for non-commercial purposes by any means, provided that this copyright notice appears on all such copies.

INTRODUCTION

One advantage of a well-functioning futures market is to alleviate welfare losses for producers and consumers due to production or consumption fluctuations caused by volatility in speculative commodity markets. The effectiveness of futures markets in reducing risk, however, is dependent in no small part on their "efficiency". If futures markets are efficient, that is, if the futures price is the best unbiased predictor of the corresponding spot price, implying that the current futures price incorporates all relevant information, agents are able to alleviate potential losses by using appropriate hedging instruments. Alternatively, if futures markets are inefficient, they may introduce an extra cost to hedgers, such as losses caused by the price volatility in the spot and futures markets. As a result, testing the efficiency of futures markets is an important research agenda for both market participants and observers.

Numerous studies have examined the efficient market hypothesis for agricultural commodity futures markets with mixed results (Holt and Mckenzie, 2003; Mckenzie and Holt, 2002; Thraen, 1999; Fortenbery and Zapata, 1997; Zapata and Fortenbery, 1996; Beck, 1994). While Mckenzie and Holt (2002) used cointegration and error correction models with GQARCH-in-mean processes to test the efficiency of four agricultural commodity futures markets (live cattle, hogs, corn, and soybean meal), their results indicate each market is efficient and unbiased in the long run. However, cattle, hogs and corn futures markets exhibit short-run inefficiencies and pricing biases. Beck used the Engle-Granger two-step cointegration procedure to test market efficiency for several agricultural commodity futures markets. Her results indicated that the market efficiency hypothesis was rejected most of the time. Zapata and Fortenbery (1996) found evidence of cointegration between spot prices and nearby futures prices for corn and soybean in Chicago and interest rates for most years examined. Fortenbery and Zapata (1997) tested market efficiency in the cheddar cheese futures market. Their results do not support the market efficiency hypothesis. However, Thraen (1999) extended the data span and used the same method to test for market efficiency. His results support the market efficiency hypothesis.

Empirical tests of market efficiency for forest commodity futures markets are, however, limited. In fact, Deckard (2000) is the only known study to examine the forward rate unbiasedness hypothesis (FRUH) in softwood lumber cash and futures markets. In his study Deckard (2000) used the conventional three-stage approach to testing the FRUH. First, both the levels and first differences of each series were examined for stationarity by using standard unit root tests, a necessary precursor for performing cointegration tests and estimating cointegrating relationships. Second, Johansen's Maximum Likelihood (ML) approach was used in the cointegration analysis. Finally, restrictions were imposed on a vector error correction model (VECM) to test the FRUH in the long-run relationship between spot and futures prices. Results provide evidence that spot and futures market prices for U.S. softwood lumber follow a stationary long-run equilibrium and, moreover, that the structure of this equilibrium is consistent with FRUH.

This paper takes an alternative empirical test of market efficiency that allows a non-linear and time-varying risk premium. We believe this alternative better reflects the reality of the market environment, and therefore should lead to a more realistic test of the market efficiency and unbiasedness hypothesis. This approach is an error correction model with generalized-quadratic ARCH-in-mean (GQARCH), which allows the error terms to be conditional heteroskedastic and the dynamic process generating the underlying heteroskedasticity to be asymmetric (Mckenzie and Holt, 2002; Beck, 2001).

In addition, two other important factors also distinguish the present study from Deckard's work on testing the EMH in futures markets for forest products. First, we expand the number of lumber and paper product markets for which the properties of market efficiency and unbiasedness are examined, while Deckard examined only the market for softwood lumber. Specifically, we examine the performance of the following three forest commodity futures markets: softwood lumber; oriented strand board (OSB); and northern bleached softwood kraft pulp (NBSK), of which the later two markets are also representative of economically important products derived from timber. Second, weekly spot and futures prices (June 1997 – October 2001), rather than monthly prices, are used to more appropriately capture short-term volatility in forest commodity futures markets.

The paper is organized as follows. In the next section, the methodology is provided. After this, data and preliminary results are reported. Next, empirical results are presented and a comparison from different approaches is conducted. The final section provides some concluding comments.

METHODOLOGY

At its core, the theory of market efficiency suggests that an efficient forward or futures pricing instrument reflects all available information at any point in time. In general, if the efficient market hypothesis holds the current futures price of a contract expiring at time t, F_{t-1} should equal the expectation of the spot price, P_t, to prevail at time t. Otherwise, market participants will use additional information to profitably buy or sell futures contracts. Consequently, market efficiency in a frictionless economy ensures that either strict equality between F_{t-1} and P_t, or equality in expectations ($F_{t-1} = E_{t-1}(P_t / \Omega_{t-1})$), where Ω_{t-1} is the information set at time t-1, holds, as long as the market is liquid and viable. This assumes, of course, zero transaction costs. Thus, in its purest form market efficiency implies that F_{t-1} incorporates all relevant information including past spot and futures prices so that F_{t-1} is the best predictor of P_t.

Market efficiency implies that the futures price (F_{t-1}) for a contract expiring at time t is the best predictor of the expected spot price (P_t). Mathematically, this can be expressed as:

$$E_{t-1} (P_t) = F_{t-1}, \tag{1}$$

where $E_{t-1} (P_t)$ is the expected future spot price formed at time t-1. Assuming rational price expectations, we obtain:

$$P_t = E_{t-1} (P_t / \Omega_{t-1}) + \varepsilon_t , \tag{2}$$

where Ω_{t-1} is the information set available in period t-1. By combining (1) and (2), we obtain:

$$Pt = Ft\text{-}1 + \varepsilon t. \tag{3}$$

Equation (3) forms the basis for conventional market efficiency and unbiasedness tests between spot (cash) and futures prices. To carry out these tests, various versions of

$$P_t = \alpha + \beta F_{t-1} + \varepsilon_t \tag{4}$$

are typically estimated. If the null hypothesis of market efficiency ($\alpha = 0$ and $\beta = 1$) can not be rejected, the futures price is an unbiased predictor of the spot price. Therefore, the forward rate unbiasedness hypothesis (FRUH) is accepted. If the null hypothesis ($\beta = 1$) can not be rejected, the market efficiency hypothesis is accepted. As a result, the hypothesis that the futures price is an unbiased predictor of spot price is a joint hypothesis that markets are efficient and that there is no risk premium.

What if the null hypothesis ($\beta = 1$) is rejected? In this case, three separate conclusions can be drawn from the results (Mckenzie and Holt, 2002): (1) the market may be inefficient; (2) a constant risk premium may exist which makes market forecasts biased but possibly efficient; or (3) a time-varying risk premium, may be inherent to the market, thus preventing futures prices in isolation from providing unbiased forecasts of future spot prices. These last two conclusions are very important in the forest product futures market for two reasons. One is that similar to agents in other futures markets, participants in forest product futures markets may be risk averse. Another is that lumber, OSB and NBSK are storable. This storable characteristic may require a premium for agents to cover their storage costs. As a result, a risk premium may be required for agents to use the futures contract to hedge their output.

Early studies often used OLS to estimate the parameters α and β in (4). Even though OLS produces asymptotically superconsistent estimates, simulation experiments provide evidence of significant small sample bias due to endogeneity and simultaneity of variables. As Elam and Dixon (1988) demonstrated by using Monte Carlo experiments, a standard F test of $\alpha=0$ and $\beta=1$ is biased toward falsely rejecting the null hypothesis of unbiasedness.

Due to these problems with OLS, latter studies have adopted cointegration techniques, including Granger and Engle two-step procedure, the Johansen maximum likelihood method and error-correction models (ECM). In the remainder of this section, a brief theoretical explanation of Granger and Engle two-step procedure, Johansen cointegration approach and ECM models are presented.

The development of cointegration modeling initially stems from the work by Granger (1986) and Engle and Granger (1987). The main objective was to test for the presence of a long-term equilibrium relationship between a set of related variables. Specifically, a vector Y_t with the property that each component of Y_t, has been differenced d times to achieve stationary may be written as yt~ I(d), where yt is a component of Yt. If all subcomponents happen to be integrated of the same order, cointegration exists. For example, if an (n x 1) I(1) vector yt is cointegrated, there exists a cointegrating vector γ, such that γ'yt ~ I(0). We then say that Z_t = γ'yt is integrated of order zero, implying that Z_t is stationary. Statistically, Z_t can't deviate too far from the trend line, or the long-term equilibrium, but this does not exclude short-term deviations. For example, if futures price for a forest product moves "too far" from the equilibrium level, buyers and sellers may engage in arbitrage so that futures price will return to its long-term equilibrium.

The Granger & Engle two-step estimation procedure can be shown in a two variable cointegrated system. Specifically, define y_{1t} and y_{2t} as:

$$y_{1t} = \varphi y_{2t} + u_{1t} \tag{5}$$
$$y_{2t} = y_{2,t-1} + u_{2t}. \tag{6}$$

The first step is to estimate φ using OLS and obtain the residuals

$$\hat{u}_t = y_{1t} - \hat{\varphi} y_{2t}\,. \tag{7}$$

The augmented Dickey Fuller (ADF) test may then be used to check if y_{1t} and y_{2t} are cointegrated. If ADF test results lead to a rejection of a unit root in $\hat{u}_t$, we conclude there is coitegration between y_{1t} and y_{2t}. Otherwise, one rejects a cointegration relationship. Engle and Granger's (1987) two-step cointegration procedure has several limitations. First, no strong statistical inference can be drawn with respect to the regression parameters α and β, as is required for efficiency testing. Second, the cointegrating vector is assumed to be unique. However, when there are more than two variables, the uniqueness of the cointegration vector cannot be guaranteed using the two-step cointegration procedure (Antonion and Foster, 1994). Finally, though the two–step cointegration procedure yields estimated coefficients which are consistent, the associated error may be misleading for any hypothesis testing (Hall, 1986 and Stock 1987).

Johansen's maximum likelihood method (Johansen, 1988, and Johansen and Juselius, 1990) provides solutions to the problems discussed above. Indeed, the maximum likelihood (ML) estimation has some advantages compared to the Engle and Granger two-step procedure. First, ML parameter estimates of the cointegrating vectors are obtained, so more detailed inference can be drawn from parameters α and β. Second, the procedure may be applied to perform likelihood ratio (LR) tests of various maintained hypothesis. Third, the procedure allows all the distinct cointegrating vectors to be identified against the alternative hypothesis that there is one (or more) cointegrating vectors. Last, but not least, Johansen's method does not impose a specific number of co-integration relationships a priori. Therefore, tests of the number of co-integration relationships are carried out simultaneously.

To illustrate the above, consider a general vector autoregression (VAR) process in levels:

$$Y_t = a + \sum_{j=1}^{p} \Pi_j Y_{t-j} + \mu_t \qquad (8)$$

Where Yt is an NX1 vector of I(1) variables of interest, Πj is an NXN matrix of parameters, a is a vector of constant and ut is an i.i.d. vector of mean-zero errors with covariance matrix Λ; ut ~ N(0, Λ).

Johansen and Juselius (1990) suggest that equation (8) can be written as:

$$\Delta Y_t = a + \sum_{j=1}^{p-1} \Gamma_j \Delta Y_{t-j} + \Pi Y_{t-p} + \mu_t \qquad (9)$$

where $\Gamma_j = -\mathrm{I} + \Pi_1 + \Pi_2 + ... + \Pi_j$ (j=1, 2, … p-1). Since both μ_t and ΔY_{t-j} are stationary, the term Π is of interest in testing for cointegration. In addition, an examination of the matrix Π can reveal the long-run cointegration relationaship between the variables in Y_t. As shown by Johansen and Juselius (1990), there are three possible forms which Π can take. When Π is full rank, or when its rank is zero, no cointegrating relationship between the variables in Y_t can be found. Only when 0 < rank(Π) < N, can a cointegrating relationship between variables in Y_t be found. In the present study, we investigate whether there is a long-run cointegration relationship between spot and futures prices for forest products. Thus, if the rank (Π) = 1, then P_t and F_{t-1} are not stationary, but are cointegrated.

In order to test restrictions on the cointegrating vector, Johansen defines the Π matrix as:

$$\underset{(nxn)}{\Pi} = \underset{nxr}{C}\,\underset{rxn}{D'} \qquad (10)$$

The Johansen's procedure for estimating parameters in Π is a restricted maximum likelihood method. Moreover, restrictions of interest may be imposed and tested vis-à-vis the C and D' matrices.

Based on the definition of the stationary for a linear combination of nonstationary variables D' Y_t, one may define:

$$D' Y_t = 0 \qquad (11)$$

In this study, $Y_t{}^*$ = (P_t, F_{t-1}, 1) and D'= (1, -β, - α), solving equation (11), gives the cointegrating relation described in equation (4).

A further test for long-run efficiency can be undertaken by imposing D' = (1, -1, 0), which normalizes the coefficient of Pt to unity and yields the restrictions α = 0, β=1.

Though Johansen's method has several advantages for testing for cointegration, limitations may also arise. One limitation is that the parameters of C and D may not be identified. The other inherent problem with Johansen procedure is the inability to test for or examine short-term dynamics. As discussed earlier, short-term dynamics perhaps caused by a change in risk premium, or transaction costs, or both may be very important to market participants and observers. A solution to the problem associated with the Johansen procedure is offered through the error-correction model (ECM). An ECM for I(1) variables implies cointegration and similarly cointegration implies error correction. This result is known as the Granger representation theorem (Granger, 1983; Engle and Granger (1987), stating that for any set of I(1) variables, error correction and cointegration are equivalent representations. It is the ECM that affords cointegration theory to reconcile the long-run equilibrium with short-run dynamics.

Engle and Granger (1987) show that error-correction models permit long-run components of variables to obey equilibrium constraints while allowing short-run components to have dynamic specifications as long as all variables cointegrate. In the present study, when the spot price (P_t) and futures price (F_{t-1}) cointegrate, the ECM can be specified as:

$$\Delta P_t = \rho\left[P_{t-1} - \varphi F_{t-2,t}\right] + \beta \Delta F_{t-1} + \sum_{i=1}^{k} \varphi_i \Delta P_{t-i} + \sum_{i=1}^{l} \beta_i \Delta F_{t-i} + \varepsilon_t \qquad (13)$$

where ε_t is a stationary series with mean zero and $\left[P_{t-1} - \varphi F_{t-2,t}\right]$ is the error-correction term. Cointegration in this form implies that ρ is less than zero as the spot price responds to movements from the long-term equilibrium position illustrated in (4). Short-term market efficiency implies that $\rho=1$, $\rho\varphi=\beta\neq 0$ and $\beta_i = \varphi_i = 0$, while short-run unbiasedness indicates that $\alpha=0$, $\rho=\beta=1$ and $\beta_i = \varphi_i = 0$.

While the ECM is capable of reconciling long-run equilibrium with short-run dynamics, it also has some limitations. First, ECM does not allow the premium to be time varying, though it may capture any constant risk effect. Second, ECM can not reflect potential nonlinear dynamics in the conditional variance of spot price changes. Finally, ECM assumes that the distribution of spot price changes is characterized by a constant variance. These limitations reduce the desirability of using the ECM for market efficiency tests per se. As discussed earlier, a risk averse agent in a futures market in which products are storable may demand a risk premium. Furthermore, the premium may be time varying in nature. This may cause more volatility in spot and futures markets. The volatility could be clustering, therefore, some GARCH effects are possible (Beck, 2001).

Recent studies by Mckenzie and Holt (2002); Beck (2001); Goodwin and Schnepf (2000); and Holt and Aradhyula (1998) argue that GARCH effects are common in commodity prices, therefore, the data-generating process (DGP) may be better represented by models allowing for time variation in the conditional second moment. Mckenzie and Holt (2002) and Beck (1993) went further to test time-varying risk premium in agricultural commodity futures market. However, no one has ever applied GARCH models in the forest products futures market. The GQARCH-M terms with an extension of the ECM specification in equation (13) is defined as

$$\Delta P_t = -\rho\mu_{t-1} + \beta\Delta F_{t-1} + \sum_{i=2}^{k}\beta_i\Delta F_{t-i} + \sum_{j=1}^{l}\varphi_i\Delta P_{t-j} + \theta\sqrt{h_t} + \upsilon_t\,, \tag{14}$$

where

$$\upsilon_t = e_t\sqrt{h_t}\,, e_t \sim IN(0,1)\,, \tag{15}$$

and

$$h_t = \omega + \sum_{i=1}^{m}\phi_i h_{t-i} + \sum_{j=1}^{n}c_{jj}\upsilon^2{}_{t-j} + \sum_{j=1}^{n}c_j\upsilon_{t-j} + \sum_{j\neq k}^{n}c_{jk}\upsilon_{t-j}\upsilon_{t-k}\,. \tag{16}$$

Where h_t is the conditional variance of spot price changes in period t. Given the GQARCH-M-ECM specification in equation, short-term unbiasedness implies: ρ=1, β=1 and $\beta_i = \varphi_i = 0$and θ=0. Short-term market efficiency implies ρ=1, $\rho\varphi=\beta\neq 0$ and $\beta_i = \varphi_i = 0$.

The GQARCH-M-ECM has several advantages, compared to ECM. First, the GQARCH-M-ECM is a more general form of the ECM, while allowing for the possible existence of a time-varying risk premium. McKenzie and Holt (2002) show that if the modified short-run market efficiency restriction ρ=1, β=1 and $\beta_i = \varphi_i = 0$ hold, the equation reduces to:

$$P_t = F_{t-1} + \theta\sqrt{h_t} \tag{17}$$

which is similar to EMH model, except that $\theta\sqrt{h_t}$ can be interpreted as a time-varying risk premium. Second, the nonlinear feedback between the conditional mean and conditional variance of spot price changes in GQARCH-M-ECM may be identified and yield superior forecasts of spot prices in comparison to the futures market. Finally, GQARCH-M-ECM may provide more efficient parameter estimates than OLS if the DGP contains GARCH-type effects. As a result, this study will adopt the GQARCH-M-ECM approach to test market efficiency in the forest commodity futures markets.

DATA AND PRELIMINARY RESULTS

The data set used in this paper consists of weekly futures and spot prices for lumber, OSB and NBSK over the period June 1996 to October 2001. Chicago Mercantile Exchange futures settlement price data for lumber and OSB were taken from Bridge database of futures prices. Spot prices for lumber and OSB were taken from weekly market reports at Random Length. Pulpex futures settlement price data for NBSK were downloaded from Pulpex website. The three-moth-five-day average futures price for the nearby contract was taken as a proxy of spot price for NBSK. The summary statistics for the data is shown in Table 1. The plots for spot and futures prices are shown in Appedix A.

Tests for unit roots

A necessary condition to carry out a cointegration test is that the data have to be non-stationary at the levels, but stationary in the differences. Each series described above was first tested for the existence of a unit root by using augmented Dickey and Fuller (ADF, 1981). The results from augmented Dickey-Fuller (ADF) tests are reported in table 2. The optimal number of augmenting lags for the model was determined by using Akaike's information criterion (AIC). They show that both spot and futures prices are nonstationary (have unit roots). However, the first difference of each time series data shows that both the futures price and spot price perform stationary (Table 3), indicating that spot and futures prices appear to be integrated of order one, I(1). Therefore, it can be concluded that the use of standard OLS procedures should be avoided as the standard F test of $\alpha = 0$ and $\beta = 1$ is biased toward falsely rejecting the null hypothesis of unbiasedness (Elam and Dixon, 1988).

Table 1. Summary statistics for the data

Commodity	Price series	Mean	Standard deviation	Minimum	Maximum	Skewness	Kurtosis
Lumber	Spot	291.65	53.21	176.00	435.00	-0.1836	-0.0853
	Futures	301.24	45.44	193.5	417.6	-0.3455	-0.3355
OSB	Spot	203.85	58.26	120.00	355	0.7374	-0.3848
	Futures	206.25	48.89	137.40	331.00	0.6414	-0.5984
NBSK	Spot	548.99	120.07	405.74	891.30	1.0905	0.4785
	Futures	541.91	124.50	398.00	926.00	1.1111	0.5800

Table 2. Results from ADF tests for lumber, OSB and NBSK

Commodity	Price series	Time trend	ADF*	Unit root
Lumber (2, 7)	Spot	w/o trend	-2.30	Yes
	Spot	with trend	-2.47	Yes
Lumber (7)	Futures	w/o trend	-1.97	Yes
	Futures	with trend	-2.02	Yes
OSB (2, 3)	Spot	w/o trend	-2.49	Yes
	Spot	with trend	-2.53	Yes
OSB (3, 4)	Futures	w/o trend	-2.25	Yes
	Futures	with trend	-2.25	Yes
NBSK(2, 10, 17)	Spot	w/o trend	-1.96	Yes
	Spot	with trend	-1.97	Yes
NBSK(2, 4, 7, 14)	Futures	w/o trend	-2.03	Yes
	Futures	with trend	-2.11	Yes

* Five percent critical values of –2.88 and –3.43 are taken from Dickey and Fully (1979) for the w/o time trend and with time trend specifications, respectively.

Table 3. Results from ADF tests for price difference in lumber, OSB and NBSK

Commodity	Price series	Time trend	ADF*	Unit root
Lumber (1)	Spot	w/o trend	-9.47	No
	Futures	w/o trend	-14.11	No
OSB (1)	Spot	w/o trend	-9.17	No
	Futures	w/o trend	-14.09	No
NBSK(1)	Spot	w/o trend	-8.45	No
	Futures	w/o trend	-12.76	No

EMPIRICAL ESTIMATION

To test if the spot price and futures price are cointegrated, Johansen's (1988) procedure was performed. As discussed early, Johansen's procedure is a multivariate approach based on maximum likelihood estimates of the cointegrating regression. The AIC criterion was used to choose optimal lag length. The trace test statistics were presented in Table 4. The results illustrate clearly that the null hypothesis of no cointegrating relationship, i.e., r=0, is rejected at the 1% level for all three commodities. The hypothesis that r=1 is not rejected at 5% level for all the commodities. These empirical results suggest that spot price (P_t) and futures price (F_{t-1}) are likely cointegrated.

Table 4. Johansen test for cointegration

Commodity	α	β	λ_{itrace} $k = 0$	λ_{itrace} $k \leq 1$
Lumber	-0.044	-0.97	25.54* (12.21)	0.77 (4.14)
OSB	-0.055	-0.995	41.63* (12.21)	0.14 (4.14)
NBSK	0.345	-1.011	29.10* (12.21)	0.36 (4.14)

Indicates significant at 5% level. The critical value is in parenthesis

As the long-term relationship between spot price (P_t) and futures price (F_{t-1}) can be represented by $P_t - \beta F_{t-1} - \alpha = 0$, the cointegrating vector D'= (1, -β, - α) normalizes P_t to one. Johansen procedure was used to conduct the tests of long-term unbiasedness hypothesis on the implied (0, 1) restrictions of α and β. The results (Table 5) show that the null hypothesis α = 0 can not be rejected at 5% level for all three commodities, while both the null hypothesis for β = 1 and the joint hypothesis α = 0, and β = 1 are rejected at 1% level. Statistically, one can not reject the null hypothesis, α = 0. However, one can reject the hypothesis that markets are efficient (β = 1) and unbiasedness (α = 0, and β = 1) in the long run.

Table 5. Johansen tests of restrictions on the cointegrating regressions

Commodity	$\alpha = 0$	p-value	$\beta = 1$	p-value	$\alpha = 0, \beta = 1$	p-value
Lumber	1.26	0.26	56.98	0.00	71.04	0.00
OSB	3.31	0.069	46.72	0.00	59.55	0.00
NBSK	0.88	0.348	116.3	0.00	127.86	0.00

The null hypotheses are shown in the tables: $\alpha = 0$; $\beta = 1$; $\alpha = 0$, $\beta = 1$. A χ^2 distribution statistics and p-values for various restrictions are shown with the degrees of freedom equal to the number of restrictions placed on the parameters.

To test short-run market efficiency and unbiasedness, both the standard ECMs and the GQARCH-M-ECM are applied for the three commodity futures markets. In the case of ECM, the AIC criteria were used to select the number of lags, then lags with significant coefficients were retained. The results of estimated coefficients and residual diagnostic tests were reported in Table 6.

As discussed earlier, if the market hypothesis holds, then futures price should not move too far away from the long-run equilibrium-state, which is reflected on the coefficient of the error correction term. In this study, all the coefficients on the error correction terms have the right sign. In addition, estimated parameters are significant for lumber and OSB, but insignificant for NBSK. All the coefficients (β) for the first difference of the futures price also have the right signs and are all significant at the 1% level. However, the Wald test results for both hypothesis $\beta = 1$ and the joint hypothesis $\alpha = 0$ and $\beta = 1$ suggest that all three futures markets fail the test of short-run market efficiency and unbiasedness at the 1% level. The reasons behind the failures, however, vary across the markets. The lumber futures market rejects short-run efficiency because more lags on both spot and futures prices have some predicting power, while for the OSB and NBSK markets, the coefficient on the first difference of the futures price is significantly different from unity.
For the results of residual diagnostic tests, the Ljung-Box test statistic for the 6^{th}-order autocorrelation (Q(6)) indicates that there is no evidence of serial correlation. In addition, the Breusch Pagan test statistics suggests that there is no evidence of heteroscedasticity in the final-form equations for all three commodities at 1% level. However, the Portmanteau Q test statistics show that ARCH effects were detected in the residuals for all the commodities.

Table 6. Error correction models

$$\Delta P_t = -\rho\left[P_{t-1} - \varphi F_{t-2,t}\right] + \beta \Delta F_{t-1} + \sum_{i=1}^{k} \varphi_i \Delta P_{t-i} + \sum_{i=2}^{l} \beta_i \Delta F_{t-i} + \varepsilon_t$$

Parameters	Lumber	OSB	NBSK
ρ	-0.17 (-3.06)	-0.35 (-6.23)	-0.003 (-0.05)
β	0.59 (9.69)	0.39 (4.33)	0.43 (8.56)
β_2	0.22 (4.14)		
β_3			0.14 (4.06)
β_4	0.09 (1.71)		
φ_1		0.51 (7.06)	
φ_6	-0.11 (-1.93)		
R^2	0.38	0.34	0.458
AIC	1623.07	1714.61	1543.60
H_0: ρ=-1	226.54**	137. 52**	239.62**
H_0: β =1	45.08**	45.39**	132.63**
Q (6)	6.11 (0.41)	3.84 (0.70)	7.01 (0.32)
B.P.	4.07 (0.54)	8.26 (0.04)	3.58 (0.31)
ARCH Q(12)	32.51** (0.00)	29.48** (0.00)	30.37** (0.00)
Wald test of Unbiasedness	232.44** (0.00)	185.16** (0.00)	254.18** (0.00)

T statistics are shown in parentheses for parameters. P-values are shown in parentheses for diagnostic test statistics.
** Indicates reject the null at the 1% level. * Indicates reject the null at the 1% level.
AIC represents the Akaike information criteria.
Q (6) represents the Ljung-Box test statistic for 6th-order autocorrelation.
ARCH Q(12) represents the Portmanteau Q test statistic for GARCH effects.
B.P. represents the Breusch pagan test statistic for heteroscedasticity.
Wald test of short-term efficiency and unbiasedness (H0: ρ = -1, β = 1, $\beta_i = \varphi_i$ =0), and associated p-values are shown in parentheses.

To further test market efficiency, unbiasedness and a non-linear and time-varying risk premium, the GQARCH-M-ECM models are applied for these three commodity markets. The model selection was based on a combination of AIC values, Loglikelihood values and residual diagnostics. The results show that GQARCH(1,1)-ECM with one lag best described the data, though all the other forms of GQARCH-M-ECM models with zero to six lags of ΔP_{t-1} and ΔF_{t-1} were tested. The results (Table 7) show that the hypotheses for both market efficiency and unbiasedness in the short run are rejected for all three commodities.

Compared to the standard ECM model, results from the GQARCH-M-ECM model provide some insightful information about three commodity futures markets. First, The coefficient θ, which reflects the non-linear time-varying risk premium, is non-statistical significant for all three commodities. Second, some significant GARCH effects are found to be persistent in the results as the summation of the terms $(\hat{c}_{11} + \hat{\psi}_1)$ ranges from 0.67 to 0.97, which are close to unity. Third, there is no evidence of a significant asymmetric effect between the conditional mean and conditional variance as no significant coefficient on the c_1 parameter is found in the GQARCH-M-ECM models for all the commodities; Finally, there is no evidence of nonlinear quadratic cross-product effects in the conditional variance of spot price changes as no significant coefficient of cross product is found in the GQARCH-M-ECM models.

CONCLUSION

This study has used modern time-series methodologies to investigate market efficiency and unbiasedness in three forest commodity futures markets: softwood lumber, oriented strand board (OSB), and northern bleached softwood kraft pulp (NBSK). Our results show that all these markets do not exhibits a cointegrating relation among the futures and spot markets, implying that there is no long-run equilibrium in these markets. The empirical test results also suggest that the commodity markets for softwood lumber, OSB and NBSK are neither efficient nor unbiased in both long-term and short-term. The conclusion for softwood lumber in this paper is different from the findings in Deckard (2000). Results from the GQARCH-M-ECM also indicate that no short-term time-varying risk premiums are found in these commodity futures markets. However, this study finds that persistent GARCH effects are very important in explaining the spot price changes over time.

Table 7. GQARCH-M-ECMs

$$\Delta P_t = -\rho(P_{t-1} - \varphi F_{t-2,t}) + \beta\Delta F_{t-1} + \sum_{i=2}^{k}\beta_i \Delta F_{t-i} + \sum_{j=1}^{l}\varphi_i \Delta P_{t-j} + \theta\sqrt{h_t} + \upsilon_t ,$$

$$h_t = \omega + \sum_{i=1}^{m}\psi_i h_{t-i} + \sum_{j=1}^{n} c_{jj}\upsilon^2{}_{t-j} + \sum_{j=1}^{n} c_j \upsilon_{t-j} + \sum_{j\neq k}^{n} c_{jk}\upsilon_{t-j}\upsilon_{t-k} .$$

Parameters	Lumber	OSB	NBSK
ρ	-0.55 (-3.97)	-0.49 (-3.39)	-0.12 (-1.28)
β	0.51 (7.63)	0.48 (8.70)	0.27 (4.98)
φ_1	-0.55 (-4.86)	-0.60 (-4.61)	-0.35 (-4.29)
ω	14.73 (0.64)	41.06 (1.83)	3.38 (2.01)
c_{11}	0.06 (0.77)	0.24 (2.37)	0.12 (2.37)
ψ_1	0.78 (2.51)	0.43 (1.87)	0.85 (14.96)
θ	-0.08 (-0.51)	0.05 (0.29)	-0.04 (-0.31)
R^2	0.36	0.31	0.38
AIC	1663.86	1715.78	1635.74
H_0: ρ=-1	510.83**	722.94**	564.14**
H_0: β =1	188.65**	150.87**	271.95**
Q (6)	8.08 (0.23)	3.34 (0.77)	9.71 (0.14)
Wald test of efficiency	156.75** (0.00)	88.64** (0.00)	221.01** (0.00)
Wald test of unbiasedness	249.64** (0.00)	300.00** (0.00)	231.52** (0.00)

T statistics are shown in parentheses for parameters. P-values are shown in parentheses for diagnostic test statistics.
** Indicates reject the null at the 1% level. * Indicates reject the null at the 1% level.
AIC represents the Akaike information criteria.
Q (6) represents the Ljung-Box test statistic for 6th-order autocorrelation.
Wald test of short-term efficiency (H0: ρ = -1, β = 1, $\beta_i = \varphi_i$ =0), and associated p-values are shown in parentheses.
Wald test of short-term unbiasedness (H0: ρ = -1, β = 1, $\beta_i = \varphi_i$ = θ =0), and associated p-values are shown in parentheses.

The non-existence of long-term equilibrium, inefficient futures markets and pricing biases in the three forest commodity markets suggest that futures price for these commodities may not offer a viable tool for price risk management over the time period that we studied. As a result, hedging opportunities using the futures contracts for these commodities may be limited until the cash and futures markets have established an identifiable long-term equilibrium relationship. Because of this, these markets may not be attractive for many participants. This may partly explain why the trading volumes were very thin at both the OSB and NBSK futures markets. As a consequence of this, the Chicago Mercantile Exchange (CME) closed the OSB futures market on November 15 in 2001 and the Pulpex shut down the operations in NBSK futures market in the end of 2003.

Given the results from this study, one may try the other alternative methodologies to look at if these markets are efficient and unbiased in the long run. One possibility is to use Smooth Transition Autoregressive (STAR) Model to test market efficiency and unbiasedness. In addition, more research work could be done by calculating relative market efficiency to find out exactly how much inefficiency the markets are for softwood lumber, OSB and NBSK. Finally, further research work may explore the possibility of using daily data for the empirical tests.

References

Antoniou A., and Foster, A.J. (1994) Short-term and long-term efficiency in commodity spot and futures markets, *Financial Markets, Institutions & Instruments*, V. 3, N. 5, 17-35.

Beck, S. (1993). A rational expectations model of time varying risk premia in commodities futures markets: theory and evidence, *International Economic Review*, 34, 149-68.

Beck, S. (1994). Cointegration and market efficiency in commodities futures market, *Applied Economics*. 26, 249-57.

Beck, S. (2001). Autoregressive conditional heterroscedasticity in commodity spot prices, *Journal of Applied Econometrics*, 16, 115-32.

Dechard, D.L. (2000). The role of CME softwood lumber futures contracts in price risk management, Presented to Southern Forest Economics Workshop (SOFEW).

Dickey, D.A., and Fuller, W. (1981). Likelihood ratio tests for autoregressive time series with a unit root, *Econometrica*, 49: 1057-1072.

Elam, E., and Dixon B. L. (1988). Examining the validity of a test of futures market efficiency, *The Journal of Futures Markets*, 8:365-372.

Engle, R. and Granger, C. (1987). Coitegration and error correction: representation, estimation and testing, *Econometrica*, 55, 251-76.

Escribano, A. (1987). Nonlinear Error-Correction: The Case of Money Demand in the UK (1878-1970). University of California: San Diego, Chapter IV, 1987; Ph.D. Dissertation.

Escribano, A. (1997). Nonlinear Error-Correction: The Case of Money Demand in the UK (1878-1970). Working Paper No. 96-55, Universidad Carlos III de Madrid.

Fortenbery, T.R. and H.O. Zapata. (1997) An evaluation of price linkage between futures and cash markets for cheddar cheese, *Journal of Futures Markets*, 17, 279-301.

Goodwin, B. K. and Schnepf, R. (2000). Determinants of Endogenous Price Risk in Corn and Wheat Futures Markets, *Journal of Futures Markets*. 20, 753-74.

Granger, C. W. J. (1983). Co-integration variables and error-correction models, *UCSD Discussion Paper* 83-113.

Granger, C. W. J. (1986). Developments in the study of cointegrated variables, *Oxford Bulletin of Economics and Statistics*, 48:213-228.

Granger, C. W. J.; Swanson, N.R. (1996). Developments in the study of cointegrated variables, *Oxford Bulletin of Economics and Statistics*, 58:537-553.

Hall, S. G. (1986). An application of the Granger & Engle two-step estimation procedure to United Kingdom aggregate wage data, Oxford Bulletin of Economics and Statistics, 48 (3): 229-240.

Hall, S. G. (1991) The effect of varying length VAR models on the maximum likelihood estimates of cointegrating vectors, *Scottish Journal of Political Economy*, 38, 317-23.

Holt, M.T. and Aradhyula, S.V. (1998) Endogenous risk in rational-expectations commodity models: A multivariate generalized Arch-M approach*, Journal of Empirical Finance*, 5, 99-129.

Johansen, S. (1988). Statistical analysis of cointegration vectors. *Journal of Economic Dynamics and Control*, 12: 231-54.

Johansen, S. and Juselius, K. (1990). Maximum likelihood estimation and inference on cointegration – with applications to the demand for money, *Oxford Bulletin of Economics and Statistics*, 52, 169-210.

Mckenzie, A. M. and Holt, M.T., (2002). Market Efficiency in Agricultural Futures Markets. *Applied Economics*. 34, 1519-1532.

Stock, J. H. (1987). Asymptotic properties of least squares estimators of cointegrating vectors, *Econometrica*, 55: 1035-1056.

Swanson, N.R. (1999). Finite sample properties of a simple LM test for neglected non-linearity in error-correcting regression equations. Statistica Neerlandica, 53: 76-95.

Thraen, C.S. (1999) A note: the CSCE cheddar cheese cash and futures price long-term equilibrium relationship revisited. *Journal of Futures Markets*. 19, 233-244.

Zapata, H.O. and Fortenbery, T.R. (1996) Stochastic interest rates and price discovery in selected commodity markets, *Review of Agricultural Economics*, 18, 634-654.

Appendix A. Charts for spot and futures prices for softwood lumber, OSB and NBSK

Figure 1. Lumber Spot and Futures Prices

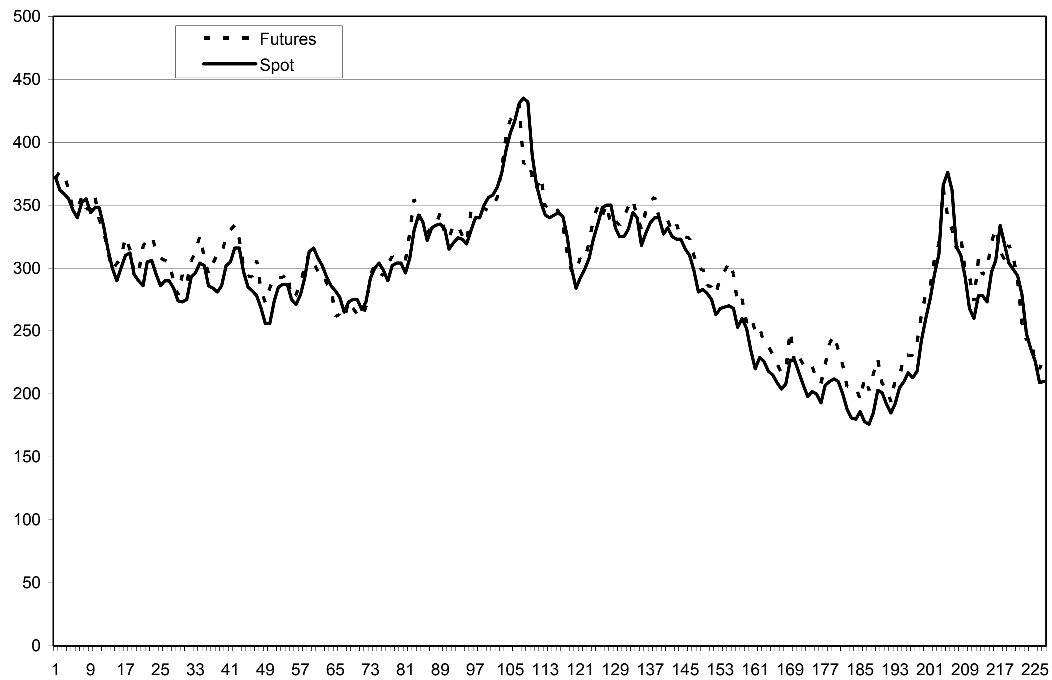

Figure 2. OSB Spot and Futures Prices

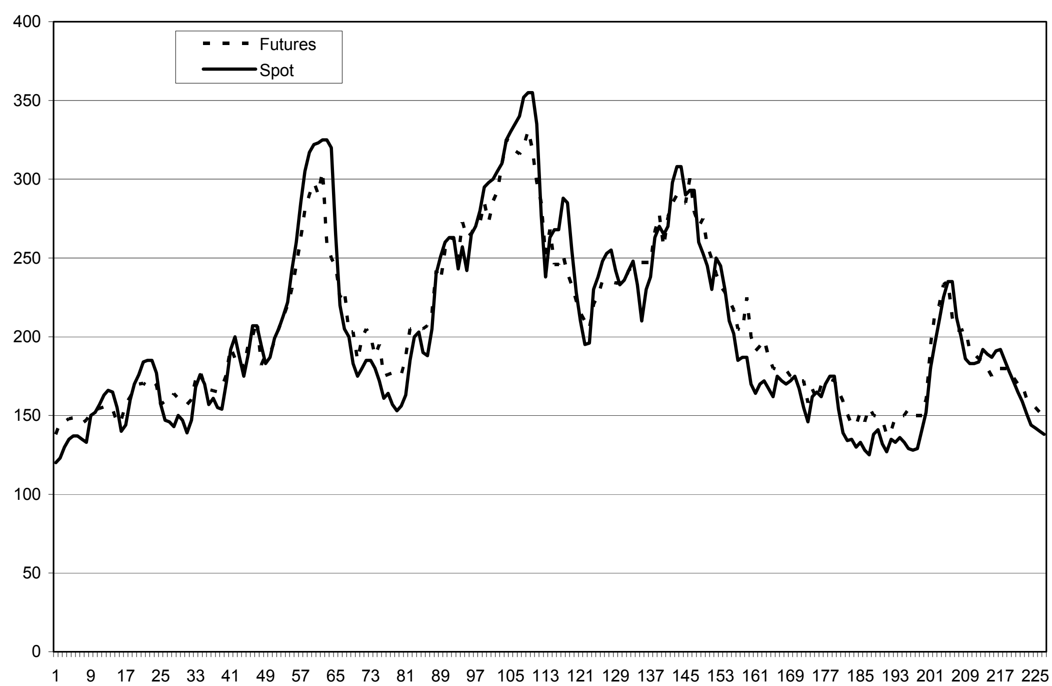

Figure 3. NBSK Spot and Futures Prices