วารสารเศรษฐศาสตร์ มหาวิทยาลัยเกษตรศาสตร์ 15 (2): 55-78

ISSN 0858-9291

Market Efficiency Test on RSS3 in Agricultural Futures Exchange
การทดสอบประสิทธิภาพยางแผ่นรมควันชั้น 3 ในตลาดสินค้าเกษตรล่วงหน้า

*Suppanunta Romprasert** Ph.D. Program in Economics, NIDA

Abstract The study investigates the efficiency of futures pricing for ribbed smoked rubber sheet no.3 (RSS3) during the period 2004-2007. It addresses the question "Is price discovery process in RSS3 futures market efficient?" Time series data from RSS3 of TOCOM was used as a leading indicator for the rubber price on the Agricultural Futures Exchange of Thailand. The results indicate that the monthly futures prices served as unbiased estimators of futures spot prices and there was no dependence on daily price changes. The tests consistently supported the unbiased hypothesis which implies that Thailand's RSS3 futures market is efficient and aids the process of price discovery. This study would fill the information gap in the prediction of futures spot prices with a guide to understanding how the futures market behaves.
Keywords: agricultural futures market, market efficiency, unbiased hypothesis, RSS3

บทคัดย่อ การศึกษาครั้งนี้มุ่งทดสอบประสิทธิภาพของราคายางแผ่นรมควันชั้น 3 ในตลาดซื้อขายล่วงหน้าระหว่างปี พ.ศ.2547-2550 โดยมีประเด็นคำถามสำคัญ คือ กระบวนการกำเนิดราคาของยางแผ่นรมควันชั้น 3 ในตลาดสินค้าเกษตรล่วงหน้ามีประสิทธิภาพแท้จริงหรือไม่ การวิเคราะห์ใช้ราคาตลาดซื้อขายล่วงหน้าโตเกียวเป็นตัวชี้นำราคาของตลาดสินค้าเกษตรล่วงหน้าในประเทศไทยผ่านทางราคาจากตลาดจริง ผลการศึกษาพบถึงความเป็นอิสระต่อกันระหว่างราคาตลาดสินค้าเกษตรล่วงหน้าแบบรายวันและแบบรายเดือนในการทดสอบความเอนเอียงกับราคาในตลาดจริง สะท้อนให้เห็นว่ากระบวนการกำหนดราคาในตลาดสินค้าเกษตรล่วงหน้าของประเทศไทยมีประสิทธิภาพ ดังนั้นจึงสามารถนำราคาดังกล่าวมาใช้เป็นตัวทำนายราคาในตลาดจริงในอนาคตได้ นอกจากนี้ผลการศึกษายังช่วยให้เข้าใจถึงพฤติกรรมของตลาดสินค้าเกษตรล่วงหน้า
คำสำคัญ: ตลาดสินค้าเกษตรล่วงหน้า ประสิทธิภาพตลาด สมมุติฐานการไม่เอนเอียง ยางแผ่นรมควันชั้น 3

* Corresponding address: *Suppanunta Romprasert* Lecturer, School of Bussiness Economics, Assumption University (Suvarnabhumi Campus) Bang Na Trad Road., Km.26 Bang Sao Tong, Samut Prakan Province, Thailand, 10540. Tel: (+66)02 7232124, Fax: (+66)02 7359599, e-mail: thailandsuppanunta@yahoo.com

Introduction

Thailand had bought and sold rubber in future contracts with traders from China, Japan, and the United States of America, but had to do so through brokers in these countries. Thailand had been less competitive than these other countries but the establishment of the futures market in Thailand provided an opportunity for Thai traders to reduce brokers' fees, plan their buying and selling, and plan on stocking rubber in the country. The development of futures markets in Thailand, and their unique institutional characteristics, prompted researchers to study the basic properties of how price behaves. At the moment, there are few published literatures on futures market in Thailand and fewer yet that are based on statistical characteristics of prices. The study would provide better information and fill some gap in the literature by making a detailed examination of futures prices in Thailand.

The Department of Internal Trade, Ministry of Commerce had initiated and consistently pursued the project to establish a commodity futures market. The Agricultural Futures Trading Act B.E. 2542 was enacted in 1999 with a provision that established the Agricultural Futures Exchange of Thailand (AFET). AFET was an independent institution established to run the agricultural futures exchange in Thailand, which was regulated by the Agricultural Futures Trading Commission. The Exchange would be the marketplace to trade agricultural futures with established rules and regulations that assured fairness for buyers and sellers. On September 20, 2001 the first Board of Directors of the AFET was appointed. The AFET launched its first futures trading in natural rubber ribbed smoked sheet no.3 (RSS3) on May 28, 2004. In addition, the second futures contract, white rice 5% broken, was listed on August 26, 2004 and tapioca starch premium grade on March 25, 2005. Since Thailand had gradually lifted price controls, the prices especially for RSS3 were decided by supply and demand as market-adjusted price. The scope of market-adjusted prices expanded continuously.

This paper seeks to answer questions on efficiency price discovery of RSS3 in Thailand. The study considered daily and monthly prices over the period January 2004-December 2007. A thorough analysis conducted on the development of Thailand's commodity futures market provided the background. Also it examined the random walk and unbiased hypotheses for RSS3. This study might have limited relevance for Thailand's markets because this commodity futures market is new and the institutional details, trade practices, and the types of investors who participate in the market are different. Based on the empirical evidence, the paper argues that Thailand's RSS3 futures market is efficient, and aids the process of price discovery because futures prices could be unbiased predictors of future spot prices.

A comprehensive test of the efficiency of rubber futures was conducted by examining a period of time over which rubber futures had existed. The process was organized as follows. First, a background is provided on the development of Thailand's commodity futures markets followed by the review of literature, the concept of efficiency test, and then the methods and data used in the study. The last two sections describe the results of the analysis, a discussion on the findings and their implications on RSS3 markets.

Futures Market Efficiency

Whether futures prices were unbiased estimates of subsequent cash prices remained to be determined empirically. Fama (1970) classified three types of test concerning efficiency of market as (i) strong-form tests in which the current information set included everything relevant; (ii) semi-strong-form tests in which the obviously publicly available information were considered; and (iii) weak-form tests in which the current information set contained the historical price series only. Most of the studies used the weak-form tests since both the strong and semi-strong tests were difficult to conduct. Many studies had examined the pricing accuracy of futures market.

Tomek and Gray (1970), Leuthold (1974), and Kofi (1973) investigated the forecasting ability of futures markets within the context of allocative efficiency. Tomek and Gray compared price relationships of two storable commodities, corn and soybeans, with a non-storable commodity, Maine potatoes. All three were produced seasonally but while stocks of corn and soybean were held continuously from harvest to harvest, stocks of potatoes were not. They found that corn and soybean market prices were relatively a better forecast than potato market prices. The difference in pricing performance between these markets indicated the significance of stock on the price spread and the influence of expectations on the price level. Kofi's study provided further support to Tomek and Gray's finding. He examined Chicago's futures market for wheat and Maine potatoes during 1953 to 1969 and found that storable commodity such as wheat provided relatively reliable forecasts of cash prices at any point in time. Kofi also showed that the longer the horizon, the less effective the futures market performed as a predictor of spot prices.

Leuthold (1972) also found that futures price were efficient forecasts of spot prices only for near-maturity dates. He also compared futures prices of live cattle representing a non-storable commodity with corn, a commodity with continuous production. Despite the clear differences between the two with respect to storage and production, he found no significant difference in the pricing accuracy. He showed that futures prices were efficient forecasters of spot prices only for near-maturity dates. Stein (1981) carried the analysis a step further and placed emphasis not only on the bias of futures market forecasts, but also on the variance of the forecast error. Stein concluded that futures prices earlier than four months prior to delivery were unreliable forecasts of closing prices.

Leuthold and Hartman (1979) examined monthly averages of daily futures prices of live hogs during 1971 to 1978. Using an economic forecasting model as a performance norm, they found that the futures market did not at all times fully reflect available information. Peck (1975) found that futures prices for eggs were as accurate as several econometric models examined. Giles and Goss (1980) studied the forward pricing functions of wool using general instrumental variables estimator (GIVE). The results supported the view that lagged futures prices were unbiased estimates of delivery date spot prices for wool with lags one to twelve months, and for live cattle with lags from one to three months. This hypothesis was generally accepted for wool except for lags of three or twelve months and was rejected for beef. Recent works by Bigman, Goldfarb, and Schechtman (1983) further supported the previous finding on futures price efficiency, i.e. the market was inefficient for the more distant futures contract. They used a simple linear regression model to test the efficiency of wheat, corn and soybeans trading at the CBOT. Based on F test they concluded that futures prices generally provided inefficient estimates of the spot price at maturity. Later, Maberly (1985), Elam and Dixon (1988), and Shen and Wang (1990) pointed out that the result based on such conventional F-test was invalid when the price series were nonstationary.

The cointegration theory developed by Engle and Granger (1987) provided a new technique for testing market efficiency. Aulton, Ennew, and Rayner (1997) reinvestigated the efficiency of UK Agricultural Commodity Futures Markets using the cointegration methodology. They found that the market was efficient for wheat but not for pork and potatoes. The cointegration method could effectively account for the nonstationary in price series. But one limitation of this approach was that no strong inferences could be drawn for the parameters, which were the central point of the efficiency tests (Lai and Lai, 1991). Johansen and Juselius (1990) derived statistical procedures for testing cointegration using the maximum likelihood method. These procedures were based on a vector autoregressive (VAR) model that allowed for possible interactions in the determination of spot prices and futures prices. Lai and Lai (1991) suggested using Johansen (1992)'s approach to test for market efficiency and illustrated the procedure with an example of the forward currency market in the US. Based on Johansen's approach, Fortenbery and Zapata (1993) evaluated the relationship of two North Carolina corn and soybean markets with respect to the CBOT. Cointegration existed between any pair of these markets and no strong evidence was found to reject the efficiency hypothesis. Mckenzie and Holt (2002) tested the efficiencies of the USA futures markets for cattle, hogs, corn, soybean meal and broilers. Their results indicated that futures markets for all the commodities except broiler were efficient and unbiased in the long run. Kellard et al. (1999) examined the efficiency of several widely traded commodities in different markets, including soybeans on the CBOT and live hogs and live cattle on the Chicago Mercantile Exchange. The results showed that the long-run equilibrium condition holds, but there was evidence of short-run inefficiency for most of the markets studied. The degree of the inefficiency was measured based on the forecast error variances.

Among the more recent studies were those by Ke and Wang (2002) on the efficiency of the Chinese wheat and soybean futures markets using Johansen (1992)'s cointegration approach. Their results suggested the existence of a long-term equilibrium relationship between the futures price and cash price for soybeans, and a weak short-term efficiency of the soybean futures market. Hourvouliades (2006) offered evidence for the efficient market hypothesis, with the series following a random walk, being cointegrated and having a long-term equilibrium.

In view of the mixed findings, it was highly probable that inaccuracies in forward pricing have caused social losses. Particularly in the non-inventory markets, social losses were minimized where futures prices were reliable forecasts. Kamara (1982) saw the need for more research to determine the magnitude of the avoidable social losses and, importantly, to examine the causes of these inefficiencies. Factors that contributed to the inefficiencies included the insufficient number of producers and firms actively trading in the market, as suggested by Leuthold and Hartman (1979) and Cappoza and Cornell (1979). Other causes of inefficiencies included irrational trades, unreliable speculation, tax effects, misinterpretation of information, transaction costs and others.

In fact, theoretical and empirical research has yet to be carried out to explore the causes of these inefficiencies.

Unit Root Test

This test showed the process of I(1) having unit root. This meant that if the hypothesis could not be rejected, for example, assuming that one variable (x) was unit root, it shows that this variable was nonstationary. There are several other methods to test besides Dicky-Fuller (DF) and augmented Dicky-Fuller (ADF), including the decision tree approach proposed by Holden and Perman (1994) and used by Mukherjee, White, and Wuyts (1998: 352-356). However, the most acceptable worldwide was DF and ADF as follows:

Null hypothesis on DF test was $H_0 : \rho = 1$ and

$$X_t = \rho x_{t-1} + et \quad (1)$$

This is called unit root test. If $|\rho| < 1, X_t$ was stationary and if $|\rho| = 1, X_t$ was nonstationary. However, this test could be done in another way that would give the same result as Equation 1.

$$\Delta X_t = \theta X_{t-1} + et \quad (2)$$

It was $X_t = 1 + \theta X_{t-1} + e_t$ that was Equation 1 where $\rho = (1+\theta)$. If θ in Equation 2 was of negative value, it would get ρ in Equation 1 having a value less than 1. So, it could be concluded that if $H_0 : \theta = 0$ was rejected, then $H_a : \theta < 0$ is accepted. It means that $\rho < 1$ and X_t had integration of order zero (Charemza and Deadman, 1992: 141) which clarified that X_t was stationary, but if $H_0 : \theta = 0$ could not be rejected, it meant X_t was nonstationary. If X_t was random walk with drift, the equation could be written as:

$$\Delta X_t = \alpha + \theta X_{t-1} + et \quad (3)$$

and if X_t was random walk with drift and had linear time trend, the equation could be written as:

$$\Delta X_t = \alpha + \beta_t + \theta X_{t-1} + et \quad (4)$$

where t was the time during which the test was carried out on $H_0 : \theta = 0$ where $H_a : \theta < 0$. Then, Dickey and Fuller (1979) considered testing the unit root test in three different forms of Equation 1 – 3 as follows: $\Delta X_t = \theta X_{t-1} + et$; $\Delta X_t = \alpha + \theta X_{t-1} + et$; and $\Delta X_t = \alpha + \beta_t + \theta X_{t-1} + et$

The interesting parameter in all the equations is θ. If $\theta = 0$, X_t would have unit root by comparing t-statistic with t-value in Dickey-Fuller tables (Enders, 1995: 221) or with MacKinnon critical values (Gujarati, 1995: 769). Moreover, Enders (1995: 221) and Gujarati (1995: 720) said that the critical values did not change when Equation 2, 3, and 4 were substituted by autoregressive processes as in the following:

$$\Delta X_t = \theta X_{t-1} + \sum_{i=1}^{p} \varphi_i + \Delta X_{t-1} + e_t \tag{5}$$

$$\Delta X_t = \alpha + \theta X_{t-1} + \sum_{i=1}^{p} \varphi_i + \Delta X_{t-1} + e_t \tag{6}$$

$$\Delta X_t = \alpha + \beta_{\tau} + \theta X_{t-1} + \sum_{i=1}^{p} \varphi_i + \Delta X_{t-1} + e_t \tag{7}$$

Also, when DF is joined with the Equation 5 – 7, it would be called ADF. The statistical value testing of ADF has asymptotic distribution as DF statistic, so that it could use the same style of critical values (Gujarati, 1995: 720).

Cointegration Test

Nonstationary data or trend being stochastic or deterministic might result in spurious regression. T-statistic would not be standard distribution or others could explain the higher goodness of fit, so that it was difficult to estimate the result from regression. However, if two variables were nonstationary both would have the higher value integration of the same order and if the difference between the two variables did not integrate, the linear combination of both variables might be stationary (Charemza and Deadman, 1992: 143). Following the definition of Engle and Granger (1987) on cointegration between two variables, if X_t and y_t were time series, X_t and y_t would be cointegrated of order d, b which could be written as $X_t, y_t \sim CI(d,b)$. If X_t and y_t were integrated of order d that could be written as I(d) and had to be linear combination on both variables assuming that $\alpha, X_t + \beta y_t$ would be integrated of order (d-b) where d>b>0. Vector $[\alpha,\beta]$ would be called cointegrating vector. Charemza and Deadman (1992: 144) gave an example "if X_t and y_t were both I(1) and error term of linear regression of both variable is a stationary process I(0), then X_t and y_t would be cointegrated of order (1,1) or $X_t, y_t \sim CI(1,1)$. Therefore, cointegration regression was an estimation technique for long-term equilibrium relationship between nonstationary series by deviations from long-term equilibrium path being stationary (Ling, Leung, and Shang, 1998). For cointegration test, Gujarati (1995: 727) used residuals from regression for cointegration test that was

$$\Delta \hat{e}_t = \gamma \hat{e}_t + v_t \tag{8}$$

Testing stationarity of $\hat{e}_t$ by t-statistic from γ standard error of γ comparing with MacKinnon critical values where the null hypothesis of no cointegration $H_0 : \theta = 0$. If γ was negative value and the negative value of t-statistic was significant, H_0 is rejected which means the variable was stationary and cointegrated (Johnston and DiNardo, 1997: 264-265).

$$H_0 : \gamma = 0 \quad (9)$$

$$H_a : \gamma < 0 \quad (10)$$

If, however, H_0 was not rejected the variable was nonstationary. But if there was multivariate analysis (i.e. it would have k-1 cointegrating vectors), it should use Johansen (1992) (Mukherjee *et al.,* 1998: 399). However, the residuals in Equation 8 were not white noise, and augmented Dickey-Fuller test could be used instead of using Equation 8. Assuming that v_t of Equation 8 had serial correlation, this equation would be used:

$$\Delta \hat{e}_t = \gamma \hat{e}_{t-1} + \sum_{i=1}^{p} a_i \Delta \hat{e}_{t-1} + v_t \quad (11)$$

and if $-2 < \gamma < 0$, one could conclude that the residual was stationary and y_t and X_t would be CI(1,1). Note that Equation 10 and 11 did not have intercept term because $\hat{e}_t$ was the residual from the regression of structural model (Enders, 1995: 345).

Error Correction Mechanism

If y_t and X_t were cointegrated, it meant both variables had long term equilibrium relationship, but in the short term might be disequilibrium. Then, one could use the error term in cointegrated equation being equilibrium error and could bring the error term joining with short term and long term behavior (Gujarati, 1995: 728). The main characteristic of cointegrated variables was time path that was influenced from deviations on long-run equilibrium. If the system went back to long-run equilibrium, the movement of some variables would respond to the size of disequilibrium in error correction model. The short-term dynamics of variables would be influenced from deviation on equilibrium in error correction mechanism (ECM), as proposed by Ling *et al.* (1998) and written as:

$$\Delta y_t = a_1 + a_2 \hat{e}_{t-1} + a_3 \Delta X_t + \sum_{h=1}^{p} a_{4h} \Delta X_{t-h} + \sum_{l}^{q} a_{5l} \Delta y_{t-1} + \mu_t \quad (12)$$

where $\hat{e}_t$ was residual of cointegrating regression equation. The a_z value meant a_z of discrepancy between actual value of y_t and long run value or equilibrium in the previous period that would be eliminated or corrected in the next period. ECM was proposed by Gujarati (1995: 729) and written as:

$$\Delta y_t = a_1 + a_2 \hat{e}_{t-1} + a_3 \Delta x_t + \mu_t \quad (13)$$

ECM was mentioned by Charemza and Deadman (1992: 146) as not having lagged Δx showing as:

$$\Delta y_t = a_1 \hat{e}_{t-1} + a_2 \Delta x_t + \mu_t \quad (14)$$

where a_1 was negative value; $1 \leq a_1 < 0$ (Patterson, 2000: 341). The reason a_1 was negative value is that if $\hat{e}_{t-1} > 0$, then $y_{t-1} > \alpha + \beta x_{t-1}$. That is, y_{t-1} had the higher value than the target value. For letting *y* on the target value, y_t is needed to reduce the value. The lower limit of a_1 was -1 meaning the elimination was perfect disequilibrium on previous period and the absolute size of a_1 was negative showing the eliminated speed was disequilibrium or speed of adjustment. The equilibrium would be restored when the absolute value of a_1 had an increasing value (Patterson, 2000: 341; Enders, 1995: 367). Enders (1995: 375) set error correction model as:

$$\Delta y_t = a_1 + a_2 \hat{e}_{t-1} + \sum_{h=1}^{p} a_{4h} \Delta x_{t-h} + \sum_{l}^{q} a_{5l} \Delta y_{t-l} + \mu_{y-t} \quad (15)$$

$$\Delta x_t = b_1 + b_2 \hat{e}_{t-1} + \sum_{m=1}^{r} b_{4m} \Delta x_{t-m} + \sum_{l=1}^{q} b_{5l} \Delta y_{t-l} + \mu_{x-t} \quad (16)$$

There was no Δx_t in Equation 15 and Δy_t in Equation 16 that was different from model used by Ling *et al.* (1998). Tambi (1999) also built error correction model the same as Equation 15.

The study used a variety of approaches to test the efficiency of RSS3 futures market. The relative performances of futures markets in forecasting different cash markets or spot markets were also evaluated along with the Johansen cointegration technique (see Johansen, 1991), VAR, VEC, and ECM to examine the unbiased hypothesis.

Methods and Data

The study focused on RSS3 futures because it was difficult to obtain sufficient monthly data observations for other futures products. AFET started trading in the RSS3 contract on May 28, 2004. The trading unit was 5 tons and quoted in Thai Baht per kilogram. There were six consecutive contract months from the nearest contract month. The daily turnover of the RSS3 contract at AFET since its initiation were 27,689 contracts traded and trading 138,445 tons at an average of 113 contracts and roughly dealing with 565 tons a day. The trading started slow in the first few months; trading average was less than 100 contracts or 500 tons a day. It gradually improved and reached the highest daily average of 442 contracts and 2,210 tons on March 30, 2005. The data on RSS3 futures prices on daily basis were collected from August 1, 2007 to January 7, 2008. The monthly data were collected from futures database of the Rubber Research Institute (RRI), Department of Agriculture.

Following the previous studies of (Lee and Mathur, 1999 and Lee *et al.,* 2000), a daily nearby closing price was selected to construct a rollover time series. First, the nearby futures contract, which was an active trading contract with the nearest delivery month to the day of trading, was specified. Prices for the nearby futures contract were selected until the contract reached the first day of the delivery month. Then, a switch from the nearby contracts to the contracts next nearest to delivery date was made during the delivery month of the nearby contracts. By constructing data in this way, all price data within the delivery month were excluded to avoid the possibility of noise during the delivery month. The nearby futures contract was selected because it was the most active and had high liquidity. The study used the FOB (BKK) as a proxy for the daily and monthly spot prices.

Regarding the above information, the data would be referred as the agricultural commodity futures market, TOCOM for RSS3. A cash market, the FOB, was chosen to test the efficiency of futures markets for RSS3. The FOB is located in Bangkok, which is the major rubber port area and the main agricultural wholesale trading market in the country. The daily and monthly futures price data of RSS3 during January 2004-December 2007 were provided by the TOCOM. Cash prices were obtained from the Rubber Research Institute using one month forward contract for RSS3 futures. Cash prices were taken at the same period of the one month forward futures contracts.

The study hypothesized that it had one taking, at one-step ahead forecast in all maturity months, by adding the forecast value to the end of time. This paper also had futures price series, which consisted of prices taken at a particular period prior to the cash price observation, assumed to be at the maturity. The number of observations of each series was the total number to be used in running the data set. The statistical properties of RSS3 futures price from 2004 to 2007 were also investigated. In particular, we examined whether (i) there was any dependence in daily and monthly futures price changes; and (ii) whether futures prices were unbiased predictors of future spot prices of the delivery dates.

Results and Discussion

Various tests were conducted to investigate the dependence in prices. Early research used serial correlation coefficients and run tests to investigate whether price series followed a random walk (Fama, 1965). More explicit tests of random walks examined whether unit roots existed in price series. Dickey and Fuller (1979, 1981) proposed unit root tests and their procedure (ADF) had the null hypothesis that a series had a unit root. A complementary test developed by Kwiatkowski *et al.* (1992)[1] used the null hypothesis that the time series of prices was stationary, and would use both the ADF and KPSS measures. The ADF and KPSS tests were based on the assumption of a normal distribution, but this might not be strictly valid for many financial time series. An alternative procedure to test the random walk hypothesis is the variance ratio test developed by Lo and MacKinley (1988, 1989). This test allowed for heteroskedasticity in the data and did not require the assumption of normality, in which case a scatter diagram with a regression line is selected. As Figure 1 depicts, the diagram clearly shows a linear relationship between the two prices.

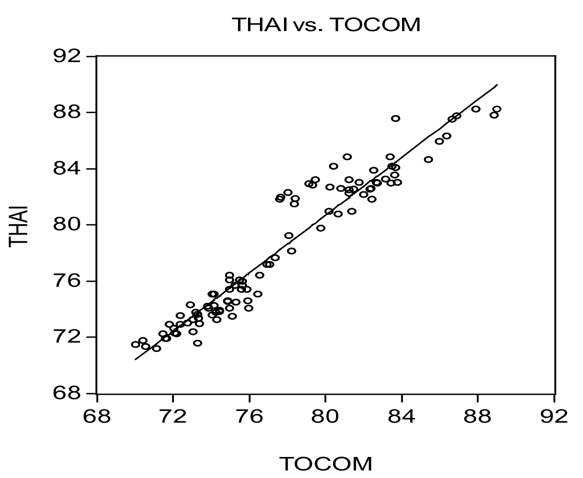

Figure 1 A scatter diagram of FOB (THAI) and TOCOM prices in daily

Figure 2, below, shows both variables having the price scaling on the Y-axis. It clarifies the idea behind the relationship between FOB (THAI) and TOCOM; the two prices had a similar pattern whereas the 'bases' for the price formation of both commodities were not identical. The data uses the daily basis, considered as having a role in the relationship between these prices.

[1] Later so-called KPSS test

FOB (THAI) and TOCOM were prices that may be considered, to a certain extent, in forecasting each other. An important property of the spot market (FOB in THAI) and RSS3 TOCOM could be price signals for RSS3. Another important feature was that both had been influenced by the world market, which had been highly fluctuating.

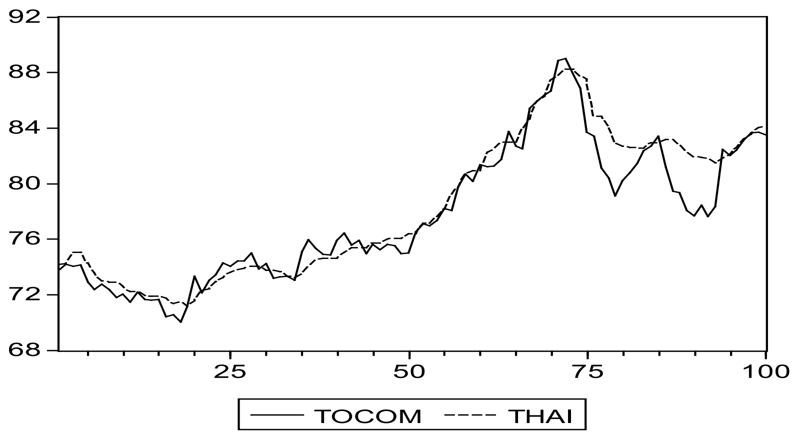

Figure 2 A combined graph of FOB (THAI) and TOCOM prices

The standard errors were the square roots of the elements on the main diagonal of the estimated covariance matrix. It can be observed that the sample had been adjusted and started in the fourth number of data and had lost three observations in the attempt to make the model dynamic with the specification of lags. In restrictions on parameters were very simple test with the Wald F-test, consider again a model in general notation on $Spotprice_{t+n} = \alpha_0 + \beta_1 Futuresmarketprice_t + \mu_t$:

$$THAI_{t+n} = \alpha_0 + \beta_1 TOCOM_t + \mu_t, \quad (17)$$

where $THAI_{t+n}$ = spot price at time t+n; n = 1,2,3,…

$TOCOM_{t,n}$ = futures market price at time t

The hypothesis to be tested would be: $H_0 : \beta_1 = 0$. A null hypothesis was tested for the contemporaneous and lagged TOCOM in the relationship between THAI and TOCOM. If H_0 was not rejected then the first differences of the TOCOM price could be specified instead of $TOCOM_t$ and $TOCOM_{t-1}$ individually. The F-statistic had a p-value of 0.01 percent, so that the null hypothesis was notrejected at the 5 percent significance level, although it was not a very convincing result. The chi-square statistic was explained at significant 0.01 and 0.05, respectively. The differences with the output of the unrestricted model were rather small, so the restriction did not disturb the original unrestricted fit to the data.

The main issue addressed by this study is estimating and solving a dynamic short-run model of an economic variable to obtain a long-run model. Developing a dynamic short-run model for simulation experiments, the study sought know and understand the properties of the model with respect to its long-run behavior. A short-run dynamic model could explode or converge in a straight or a cyclical way, which were the properties of different equations. To test the null hypothesis of a unit-root, the β_i was estimated using OLS: $H_0 : \beta_i = 1$; $H_a : 0 < \beta_i < 1$. The level and 'Intercept' were selected including the automatic selection that yielded an acceptable result: i.e. the lowest value of the Schwartz Information Criterion (SIC) 1.426 was chosen for a regression: The null hypothesis of a unit root in FOB (THAI) was not rejected. An AR (3) process was assumed for the FOB (THAI) price ((p-1) = 2). The next process was to test for a second unit root by testing for a unit root in the first differences of the FOB (THAI) price. After a first regression, it was obvious that the constant term could be dropped from the regression or a zero coefficient and t-value, so the final result was for 1st difference only, then the FOB (THAI) price integrated of first order: FOB ($THAI_t$)~ I(1).

In the dynamic model, the null hypothesis was tested that FOB (THAI) followed a unit-root process. The test on unit root of price variable on cash market in Thailand and TOCOM futures market was carried out to determine if the spot market and TOCOM futures market had stationarity at I(1) and in line with the condition to test the cointegration of Engle-Granger that the variables must have the same level at I(d). The analysis, using Equation 18 - 19 below found that the TOCOM futures market had a significant relationship with spot market in Thailand from method of least squares on dependent variable of FOB ($THAI_t$), from the output of the Wald F-test, and from the result of the serial correlation LM test or known as BG(p)-test.

$$THAI = \alpha_0 0 + \beta_0 0TOCOM + \beta_1 TOCOM(-1) + \beta_2 TOCOM(-2) + \beta_3 TOCOM(-3) \quad (18)$$

$$THAI = -5.6263 + 0.3847TOCOM^{**} + 0.2112TOCOM(-1) + 0.064TOCOM(-2) + 0.4221TOCOM(-3)^{**}$$

$$(-3.5444) \quad (4.2381) \quad (1.5945) \quad (0.4832) \quad (4.6281) \quad (19)$$

R-square = 0.9685 LM-Test: NR square = 0.016 Wald Test: Chi-square = 17.9614

Note: ** significant at the 1% level

From Equation 17, estimate μ_t using unit root. The result is shown in Table 1 below:

Table 1 Random walk tests for RSS3 daily and monthly near by closing prices

RSS3	Lag	ADF value	Brandwidth	KPSS value
Short daily	0	-0.864077**	9	1.177248**
Long monthly	1	-2.088818*	5	0.537095*

Notes: * significant at the 5% level and ** significant at the 1% level

Table 1 shows the results, RSS3 was the daily and monthly nearby closing price series. For the ADF test with constant, the critical values of short and long are -2.88 and -3.48; -2.94 and -3.60 at the 5% and 1% significance levels, respectively. The null hypothesis stated that the series was non-stationary and rejected if the test statistic was greater than the critical value. For the KPSS test, the critical values of short and long were 0.46 and 0.74; 0.46 and 0.74 at 5% and 1% significance levels, respectively. The null hypothesis stated that the series was stationary and rejected if the test statistic was greater than the critical value. using ADF and KPSS, of the daily and monthly nearby futures closing price series. Based on the results of the ADF test, the null hypothesis of a unit root (nonstationary and random walk) could not be rejected for RSS3. And based on the results of the KPSS test, the null hypothesis of stationarity (no unit root) could be strongly rejected for RSS3. The ADF and KPSS tests provided consistent empirical evidence that supported the random walk hypothesis. However, this conclusion was based on the unit root tests and should be interpreted with caution because the assumption of normality was not valid. Based on the above results, estimated μ had stationarity at I(0) when tested with ADF-test and KPSS and were significant at 1% and 5%, respectively.

Therefore, it is concluded that the TOCOM futures market had long run relationship and cointegration with the spot market in Thailand. It should be noted that Equation 19 presented serial test by using LM-test for testing residual, which could explain that the Equation 19 in testing unit roots showed no serial correlation itself. Considering that the value of NR square was less than Chi-square (q = the number of lag), the null hypothesis is accepted that there was no serial correlation in the model which Chi-square (3) test found significant at 1% and 5% equal to 11.34 and 7.81, respectively. To accept the stationary qualification could be explained by the fact that the market was efficient (Hakkio and Rush, 1989) and the FOB (THAI) had a long run relationship to TOCOM. However, while it was necessary, it must also be tested for bias.

The “unbiased” hypothesis was expressed through the Johansen cointegration technique, VAR, VEC ECM to examine the unbiased hypothesis. Did the test on the lack of bias of the futures market identify the futures market? The testing of $H_0 = \alpha_0 = 0$ and $\beta_0 = 1$ from Equation 19 where: α_0 = -5.6263 and β_0 = 0.3847 using Wald-test found that the Chi-square calculated had a value of 17.9614. Its value was greater than Chi-square (q=the number of limitation) significant at 0.05 equal to 12.84. As such the hypothesis was accepted that TOCOM futures market was unbiased for FOB (BKK) of Thailand.

The long run relationship between FOB(BKK) and TOCOM got along with the hypothesis of market efficiency when $FOB(BKK)_{t+n} - TOCOM_{t,n} = 0$. Sabuhoro and Larue (1997) mentioned three ways to test the hypothesis, namely, (i) the Engle and Granger; (ii) the ECM and the restriction on ($-0\alpha_0, -a\beta_0$) = (0,-a) and (iii) the Johansen and Juselius test using ML ratio test, which assumed that ($\beta_1, \beta_0, \alpha_0$) = (1,-1,0). It meant the elasticity between TOCOM and FOB (THAI) is equal to 1 ($\beta_1 = \beta_0$) and price in TOCOM was an indicator that was unbiased where ($\alpha_0 = 0$).

Table 2, 3, 4 and 5 show the results of Johansen cointegration test, VAR, VEC and ECM between the monthly spot and futures prices for RSS3 futures, respectively. Based on the trace statistics in Table 2 (with the max-Eigen value test, I obtained the same conclusion), that one cointegrated equation existed between monthly spot and 1-month-forward futures prices at the 5 percent significance level for RSS3 futures markets. This would indicate which monthly spot and futures prices were cointegrated and which monthly futures prices could be considered as unbiased predictors of future spot prices. Because the unbiased hypothesis was consistently supported, it is concluded that Thai's RSS3 futures markets were efficient, and it played an important role in price discovery.

Table 2 Johansen cointegration test between the monthly spot and futures prices for RSS3 futures

Number of CE(s)	Trace statistics	5% critical value	1% critical value
None*	18.73607	15.41	20.04
At most 1*	4.046373	3.76	6.65

Notes: * significant at the 5% level; ** significant at the 1% level

Table 3 Error correction model (ECM) between the monthly spot and futures prices for RSS3 futures

Variable	Coefficient	t-statistic	Prob.
LTOCOM	1.008473	5.42**	0.0000
RESID01	1.000000	3.19**	0.0000
R-squared	1.000000		
Adj. R-squared	1.000000		
F-statistic	1.47		
Dubin-Watson test	0.184317		

Notes: * significant at the 5% level, ** significant at the 1% level, and LTOCOM = monthly futures prices for RSS3 TOCOM

Table 4 Vector autoregression (VAR) between the monthly spot and futures prices for RSS3 futures

Tests	LTHAI	LTOCOM
F-statistic	59.50524**	53.18303**
R-squared	0.868623	0.855266
Adj R-squared	0.854026	0.839184
Akaike AIC	6.336984	6.418247

Notes: * significant at 5% level, ** significant at 1% level, LTHAI = monthly spot prices, and LTOCOM = monthly futures prices for RSS3 TOCOM

Table 5 Vector error correction (VEC) between the monthly spot and futures prices for RSS3 futures

Test	LTHAI	LTOCOM
F-statistic	1.478865	1.546337
R-squared	0.178631	0.185271
Adj. R-squared	0.057842	0.065458
Akaike AIC	6.492978	6.563192

Notes: * significant at the 5% level; ** significant at the 1% level, LTHAI = monthly spot prices, and LTOCOM = monthly futures prices for RSS3 TOCOM

An ECM was a neat way of combining the long run, cointegrating relationship between the levels variables and the short run relationship between the first differences of the variables. It also had the advantage that all the variables in the estimated equation were stationary, hence there was no problem with spurious correlation. First Order ECMA two variables, first order ECM was specified as follows:

$$\Delta y_t = \beta_1 \Delta x_t - (1-\alpha)\left[y_{t-1} - \gamma_1 - \gamma_2 x_{t-1}\right] + u_t \quad (20)$$

$$\Delta FOB_t = \beta_1 \Delta LTOCOM_t - (1-\alpha)\left[FOB_{t-1} - \gamma_1 - \gamma_2 LTOCOM_{t-1}\right] + u_t \quad (21)$$

where the term in the square brackets was the disequilibrium error in the previous period, i.e. the deviation of FOB_{t-1} from its long run equilibrium value, and μ_t was the standard error term. β_1 measured the short run impact of changes in FOB on LTOCOM, γ_z measured the long run impact. If the variables are in logs, then these were short run and long run elasticities, respectively. $(1 - \alpha)$ was the fraction of the previous period's disequilibrium error that was made up in this period. The coefficient was expected to be negative (note the minus sign in front), and, most likely, less than one. Thus, if FOB_{t-1} was above the long run value predicted by ($\gamma_1 + \gamma_2 TOCOMt\text{-}1$) the disequilibrium error was positive. Hence this period FOB_t falls (ΔFOB_t is negative) in order to move y back to its long run equilibrium value. If FOB and LTOCOM were

I(1) variables, then both ΔFOB and ΔTOCOM are stationary. The term in brackets is stationary if the variables were cointegrated. Thus, all the variables in the equation were stationary and a valid OLS estimation was possible. In other words, an ECM can only be estimated between variables that are all I(1), and cointegrated. Hence, the requirement tested for the order of integration (via unit root tests) of the variables, and then for cointegration prior to running the ECM.

The results presented in Table 3 suggest two main views as follows (i) the change in TOCOM affected the change in spot price in the same direction, but the effect was insignificant because t-statistic calculation was higher than t-statistic critical point; and (ii) no matter what situations have caused long term volatility in spot price, the movement back to static of rubber price would adjust in each period by increasing the size of 1.000000 or by a faster movement on spot price in the long term. However, this result might have autocorrelation problem because the Durbin-Watson statistic was substantially less than 2; there was an evidence of positive serial correlation.

A VAR with p lags could always be equivalently rewritten as a VAR with only one lag by appropriately redefining the dependent variable. The transformation amounts to merely stacking the lags of the VAR(p) variable in the new VAR(1) dependent variable and appending identities to complete the number of equations. In the paper, the VAR(2) model is written as:

$$y_t = \alpha + \beta_1 y_{t-1} + \beta_2 y_{t-2} + e_t \qquad (22)$$

$$LTHAI_t = \alpha + \beta_1 LTHAI_{t-1} + \beta_2 LTHAI_{t-2} + e_t \qquad (23)$$

In Table 4 the following F-statistic for two-lag model exclude the null hypothesis even at the 0.5% level, which meant that a very robust model at the level of regression was obtained.

A VEC model could lead to a better understanding of the nature of any nonstationarity among the different component series as well as improve longer term forecasting over an unconstrained model. The VECM(p) form was written as:

$$\Delta y_t = \delta + \Pi y_{t-1} + \sum_{i=1}^{p-1} \Phi_i^n \Delta y_{t-i} + \varepsilon_t \qquad (24)$$

where Δ= the differencing operator, such that $\Delta y_t = y_t - y_{t-1}$. It had an equivalent VAR(p) representation as described in the preceding.

$$\Delta LTHAI_t = \delta + \Pi LTHAI_{t-1} + \sum_{i=1}^{p-1} \Phi_i^n \Delta LTHAI_{t-i} + \varepsilon_t \qquad (25)$$

A VEC model is a restricted VAR that has cointegration restrictions built into the specification, and is designed for the use of nonstationary series that are known to be cointegrated. The VEC model was relevant to the research study. This VEC specification restricted the long run behavior of the spot price to converge to its cointegrated relationships while allowing a wide range of short run dynamics. The cointegration term is known as the error correction term since the deviation from long run equilibrium was corrected gradually through a series of partial short run adjustments. The VEC model was under-specified and an augmented VEC model called the VECM-lead model was fitted, where the first cointegrated vector is treated as a leading exogenous variable. As Table 5 shows, the LTHAI was fully captured by the lead augmentation, and the lead and current cointegrated vector 1 was common to both the LTOCOM and LTHAI process. The error correction process could be said to mirror the data generation process for LTOCOM and LTHAI series. Hence the error correction process along with the return generation process reflected a specific uncertainty in price. The hypothesis on ECM could be tested by (a) testing on the restriction "no risk premium" in the test of unbiased and, if rejected, (b) testing the statement "Was bias created by risk premium?"

1) Under "no risk premium", the hypothesis was $a\alpha = 0, -a\beta_1 = -a$, and $\beta_k = \gamma_k = 0$ (Hakkio and Rush, 1989 and Sabuhoro and Larue, 1997). Accepting the hypothesis meant that the market had no risk premium and that the market was efficient. Rejecting it however does not mean that the market was inefficient because it does not have the real efficiency or because it has risk premium (Sabuhoro and Larue, 1997).

2) Assuming "risk premium", the hypothesis is $a = 1, a\beta_1 = b$, and $\beta_k = \gamma_k = 0$ (Beck, 1994). Accepting the hypothesis meant there was risk premium, meaning that the market was efficient even if it had risk premium, but rejecting the hypothesis could mean that the market was inefficient.

Testing $a > 0$ showed that change in spot market responded to the deviation from the long term equilibrium with restriction on b not equal to zero. Receiving the news results in a change in the futures market, the test on lag variables showed that the past information was already included in futures market. If the restriction on lag variable did not correspond to the restriction showing that the change in past price on futures market provides an indication of the change in future price on cash market. Futures market would not fully affect to future price of spot market, which meant futures market would be inefficient. For the same reason, the efficiency market must have no serial correlation or the qualification of et must be serially uncorrelated. Testing for serial correlation by using LM-test, and the hypothesis is $-a = b = 1$ being the case of "no risk premium" (Sabuhoro and Larue, 1997) $a = 1, a\beta_1 = b$ being the case of "having risk premium" (Beck, 1994) and would use the Wald test on the hypothesis.

The result on testing the unbiased by depicting on "risk premium" was coming along with the hypothesis: $a = 1, \beta_1 = b \neq 0$ and $\beta_k = \gamma_k = 0$. It was found that the model showed the relationship between FOB (THAI) and TOCOM rejecting the hypothesis unbiased depicting on "risk premium" significant at 0.01. However, rejecting the hypothesis could not support the conclusion that it is caused by real inefficiency of the market or because of risk premium. Testing the lag variables on $\beta_k = 0$ and $\gamma_k = 0$ for lag (1) showed that the change in the past price on FOB (THAI) and TOCOM affected the change in the present price on FOB (THAI). The test a = 0, model of FOB (THAI) and TOCOM accepted the hypothesis, which meant that the change in FOB (THAI) did not respond to the long run equilibrium bias.Besides all the results, the test on b = 0 found that t-ratio on FOB (THAI) and TOCOM rejected the hypothesis b = 0, showing that the one month changing on TOCOM affected the changing FOB (THAI) at that current time. Since cointegration was a necessary condition for futures marketefficiency, the cointegration testing results rejected the null hypothesis r = 0 at 0.05 level ofsignificance while the corresponding hypothesis r = 1 could not be rejected. This suggests that the futures price of TOCOM was cointegrated with the cash price in the FOB (BKK). These results also indicated that the futures price had a closer relationship with the cash price in a shorter forecasting horizon than in a longer one. This finding is reasonable because the futures price contained better information about the supply and demand of the commodity when it got closer to its maturity.

Besides cointegration, efficiency also required the futures price to be an unbiased predictor of the cash price. The study tested hypotheses a = 0 and b = 1, the results of which were that the null hypotheses a = 0 and b = 1 were not rejected at a significance level of 0.01 nor at 0.05 level. This meant that the TOCOM futures price was an unbiased predictor of cash prices. However, the unbiased assumption was too strong to imply market efficiency. As pointed out earlier, the unbiased hypothesis may be rejected with the existence of a risk premium even when the market was efficient. Therefore, more inferences could be drawn from the separate tests of a = 0 and b = 1. The null hypothesis a = 0 was rejected in all cases at 0.05 or higher level of significance and at 0.1 level. This indicated a non-zero risk premium from traders in the futures market. The null hypothesis b = 1 could be rejected at 0.1 level. This indicated the TOCOM futures market was efficient in the very long-term. Although was still rejected at 0.05 or even at 0.01 in all other cases, the p-values were generally larger than those corresponding cases for a = 0. This suggested that the market efficiency was less likely to be rejected than the zero risk premiums, thus the latter was the main contributor to the rejection of joint test hypothesis.

Conclusion

The rapid growth of Thailand's agricultural output has been driven by large increases in the export of basic commodities such as natural rubber and rice. The demand for these commodities had resulted in a dramatic increase in spot prices as well as price volatility in recent years. Thus the development of futures markets was seen as a vital step in reducing uncertainty and aiding price discovery. The new commodities futures market was able to provide efficient prices and became an integral mechanism in Thailand's economic development. The analysis was based on the market efficiency theory. The specific purpose of this study was to examine the efficiency of futures pricing for RSS3 products during the period 2004-2007. In this regard, the result indicated that (i) there was no dependence in daily price changes for RSS3 futures, and (ii) monthly futures prices served as unbiased estimators of futures spot prices. Therefore, Thailand's RSS3 futures market was efficient and it aided the process of price discovery.

Along with the test on bias, it was found that changes in the cash market in Thailand involved changes in TOCOM combined with past information in both markets. It meant that change in past prices of TOCOM would indicate change in the futures price in the spot market. Moreover, tests showed there was indeed a risk premium.

The positive results of this study could contribute to the skills and insights of the participants in the market. They would help the traders of the commodities in the futures market make better decisions. The traders include those from companies, state owned enterprises, and individual investors. One assumption of the study was that the futures market players had sufficient knowledge and experience in trading but needed more and better information as a basis for their decisions. This study aims to fill the gap with a guide to understand how the futures market behaves.

This study proved that the RSS3 futures market is efficient. This indicates that the RSS3 producers in Thailand would receive new and valuable information on RSS3 prices in futures market. Government intervention would also be necessary. Further analysis should combine the price discovery and market integration. The variance matrix of the error terms can be developed and used to study the current and long-run behavior of auction markets.

References

Aulton, A. J., C. T. Ennew, and A. J. Rayner. 1997. "Efficiency Tests of Futures Markets for UK Agricultural Commodities." **Journal of Agricultural Economics** 48 (1-3): 498-524.

Beck, S. E. 1994. "Cointegration and Market Efficiency in Commodities Futures Markets." **Applied Economics** 26 (3): 249-257.

Bigman, D., D. Goldfarb, and E. Schechtman. 1983. "Futures Market Efficiency and the Time Content of Information Sets." **Journal of Futures Market** 3 (3): 321-334.

Cappoza, D. R. and B. Cornell. 1979. "Treasury Bill Pricing in the Spot and Futures Markets." **Review of Economics and Statistics** 725 (4): 513-520.

Charemza, W. W. and D. F. Deadman. 1992. **New Directions in Econometric Practice: General to Specific Modeling, Cointegration.** Aldershot: Edward Elgar.

Dickey, D. A. and W. A. Fuller. 1981. "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root." **Econometrica** 49 (4): 1057-1072.

Dickey, D. and W. Fuller. 1979. "Distribution of the Estimates for Autoregressive Time Series with Unit Root." **Journal of American Statistical Association** 74 (366): 427-431.

Elam, E. and B. L. Dixon. 1988. "Examining the Validity of a Test of Futures Market Efficiency." **The Journal of Futures Markets** 8 (3): 365-372.

Enders, W. 1995. **Applied Econometric Time Series.** New York: John Wiley & Sons.

Engle, R. F. and C. W. J. Granger. 1987. "Co-integration and Error Correction: Representation, Estimation and Testing." **Econometrica** 55 (2): 251-276.

Fama, E. 1965. "The Behavior of Stock Market Prices." **Journal of Business** 38 (1): 34-105.

Fama, E. 1970. "Efficient Capital Markets: A Review of Theory and Empirical Work." **Journal of Finance** 25 (2): 383-417.

Fortenbery, T. R. and H. O. Zapata. 1993. "An Examination of Cointegration Relations between Futures and Local Grain Markets." **Journal of Futures Markets** 13 (8): 921-932.

Giles, D. E. A. and B. A. Goss. 1980. "The Predictive Quality of Futures Prices with an Application to the Sydney Wool Futures Markets." **Australian Economic Papers** 9 (35): 291-300.

Gujarati, D. N. 1995. **Basic Econometrics.** London: McGraw-Hill.

Hakkio, C. S. and M. Rush. 1989. "Market Efficiency and Cointegration: An Application to the Sterling and Deutschmark Exchange Rates." **Journal of International Money and Finance** 8 (1): 75-88.

Holden, D. and R. Perman. 1994. "Unit Roots and Cointegration for the Economist." In B. H. Rao. (ed.). **Cointegration for the Applied Economist.** New York: St. Martin's Press, 47-112.

Hourvouliades, N. L. 2006. "Stock Index Futures, Trading Volume and Open Interest: Efficiency and Forecasting in ADEX." **The 5th Annual Conference of the Hellenic Finance and Accounting Association,** Thessaloniki, Greece, December 2006.

Johansen, S. 1991. "Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models." **Econometrica** 59 (6): 1551-1580.

Johansen, S. 1992. "Determination of Cointegration Rank in the Presence of a Linear Trend." **Oxford Bulletin of Economics and Statistics** 54 (3): 383-397.

Johansen, S. and K. Juselius. 1990. "Maximum Likelihood Estimation and Inference on Cointegration: With Applications to the Demand for Money." **Oxford Bulletin of Economics and Statistics** 52 (2): 169-210.

Johnston and DiNardo. 1997. **Econometric Methods.** Mcgraw-Hill.

Kamara, A. 1982. "Issues in Futures Market: A Survey." **Journal of Futures Market** 2 (3): 261-294.

Ke, C. and Y. Wang. 2002. "Nonparametric Nonlinear Regression Models." **Technical Report No.385.** Department of Statistics and Applied Probability, University of California, Santa Barbara.

Kellard, N., P. Newbold, T. Rayner, and C. Ennew. 1999. "The Relative Efficiency of Commodity Futures Markets." **The Journal of Futures Markets** 19 (4): 413-432.

Kofi, T. 1973. "A Framework for Comparing the Efficiency of Futures Markets." **American Journal of Agricultural Economics** 55 (4): 584-594.

Kwiatkowski, D., P. C. B. Phillips, P. Schmidt, and Y. Shin. 1992. "Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root: How Sure are We that Economic Time-series have a Unit Root?." **Journal of Econometrics** 19 (1): 159-178.

Lai, K. S. and M. Lai. 1991. "A Cointegration Test for Market Efficiency." **The Journal of Futures Markets** 11 (5): 567-575.

Lee, C. I. and I. Mathur. 1999. “Efficiency Tests in the Spanish Futures Markets.” **Journal of Futures Markets** 19 (1): 59-77.

Lee, C. I., C. G. Kimberly, and I. Mathur. 2000. “Efficiency Tests in the French Derivatives Market.” **Journal of Banking and Finance** 1 (5): 787-807.

Leuthold, R. 1974. “The Price Performance on the Futures Markets of a Nonstorable Commodity: Live Beef Cattle.” **American Journal of Agricultural Economics** 56 (2): 271-279.

Leuthold, R. M. 1972. “Random Walk and Price Trends: Live Cattle Futures Market.” **Journal of Finance** 27 (4): 879-889.

Leuthold, R. M. and P. .A. Hartman. 1979. “A Semi-Strong Form Evaluation of the Efficiency of the Hog Futures Market.” **American Journal of Agricultural Economics** 6 (3): 482-489.

Ling, B., P. Leung, and Y. Shang. 1998. “Behaviour of Price Transmissions in Vertically Coordinated Markets: The Case of Frozen Black Tiger Shrimp *(Penaeus monodon).*” **Aquaculture Economics and Management** 2 (3): 119-128.

Lo, A. W. and A. C. MacKinley. 1988. “Stock Market Prices Do Not Follow Random Walks: Evidence from a Simple Specification Test.” **Review of Financial Studies** 1 (1): 41-66.

Lo, A. W. and A. C. MacKinley. 1989. “The Size and Power of the Variance Ratio Test in Finite Samples: A Monte Carlo Investigation.” **Journal of Econometrics** 40 (2): 203-238.

Maberly, E. D. 1985. “Testing Futures Market Efficiency - A Restatement.” **Journal of Futures Markets** 5 (2): 425-432.

Mckenzie, A. M. and M. T. Holt. 2002. “Market Efficiency in Agricultural Futures Markets.” **Journal Applied Economics** 34 (12): 1519-1532.

Mukherjee, C., H. White, and M. Wuyts. 1998. **Econometrics and Data Analysis for Developing Countries.** London: Rutledge.

Patterson, K. 2000. **An Introduction to Applied Econometrics**. London: Macmillan Press.

Peck, A. E. 1975. “Hedging and Income Stability: Concept, Implications and an Example.” **American Journal of Agricultural Economics** 57 (3): 410-419.

Sabuhoro, J. B. and Larue, B. 1997. “The market efficiency hypothesis: The case of coffee and cocoa futures.” **Agricultural Economics** 16 (3): 171-184.

Shen, C. and L. Wang, 1990, “Examining the Validity of a Test of Futures Market Efficiency: A Comment.” **The Journal of Futures Markets** 10 (2): 195-196.

Stein, C. M. 1981. "Estimation of the Mean of a Multivariate Normal Distribution." **The Annals of Statistics** 9 (6): 1135-1151.

Tambi, N. E. 1999. "Co-integration and error-correction modeling of agricultural export supply in Cameroon." **Agricultural Economics** 20 (1): 57-67.

Tomek, W. and R. Gray. 1970. "Temporal Relationships among Prices on Commodity Futures Markets: Their Allocative and Stabilizing Roles." **American Journal of Agricultural Economics** 52 (2): 372-380.