Spatial Pricing Efficiency: The Case of U.S. Long Grain Rice

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Abstract

This study applies a time-series framework to analyze long-run price relationships for Arkansas, Mississippi, Louisiana, Texas and California long grain rice. The results showed that when transportation cost is taken into account in the analysis, long grain rice price in Arkansas, Louisiana, Texas and Mississippi do have a long run equilibrium relationship, but not with California’s.

Keywords: rice, spatial, price, integration, cointegration, efficiency
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Background

The U.S. rice industry has operated in a very dynamic market environment over the past two decades. Instability in global rice trade, shifts in rice policy, changes in production and milling technology have resulted in an uneven development of the U.S. rice industry (Wailes and Gauthier). Chavez studied pricing efficiency for the period from 1978 to 1990. He found evidence, using a price margin model, that pricing efficiency had improved significantly in the latter half of the 1980s compared to previous years.

Since 1990, pricing efficiency in the U.S. rice market has not been studied. The industry has continued to experience global trade and price instability. U.S. price and income policy reforms, greater market access through the WTO Uruguay Round agreement for high quality medium grain markets in Japan, long grain trade shocks associated with El Niño and the emergence of Vietnam, China and India as major exporters have restructured the U.S. rice market. Export shares have declined as domestic consumption has increased and U.S. rice has become less competitive in many foreign markets. The share of long grain rice production has increased relative to medium grain in the southern states. Excess milling capacity has persisted and increased in Texas. In light of these changes it is important to reassess pricing efficiency in the U.S. rice industry.

In this study, we focus on spatial price integration in the U.S. long grain rice markets. Long grain rice is produced in each of the five major rice producing states (Arkansas, Louisiana, Mississippi, Texas and California). It is the dominant rice type produced in all states except California where it accounts for less than five percent of rice production. The shifts in production toward more
long grain rice reflect changes in demand for both in the domestic as well as in export markets.

State-level long grain price data over twelve years (1986-1998) showed that the average monthly cash price per cwt received by farmers for long grain rice in Texas is higher than in the other five rice producing states, which is not surprising given the importance of Houston, Texas as a major export center. The average price differential ranges from the lowest in Arkansas at $0.12 to $1.64 in California (Table 1). In terms of price variability, Mississippi’s price is the most volatile followed by that of Texas, Louisiana and Arkansas. Given the production share increases of long grain rice in the U.S. it is important to understand how these price series are related to each other.

Table 1- Basic Statistics of U.S. State-level Long Grain Rice Cash Prices ($ per cwt)$^2$

<table>
<thead>
<tr>
<th>States</th>
<th>Mean</th>
<th>Variance</th>
<th>Minimum</th>
<th>Maximum</th>
</tr>
</thead>
<tbody>
<tr>
<td>Arkansas</td>
<td>7.70</td>
<td>4.31</td>
<td>3.17</td>
<td>11.40</td>
</tr>
<tr>
<td>California</td>
<td>6.18</td>
<td>1.67</td>
<td>2.73</td>
<td>8.25</td>
</tr>
<tr>
<td>Louisiana</td>
<td>7.67</td>
<td>4.36</td>
<td>3.77</td>
<td>11.50</td>
</tr>
<tr>
<td>Mississippi</td>
<td>7.68</td>
<td>4.40</td>
<td>3.05</td>
<td>11.40</td>
</tr>
<tr>
<td>Texas</td>
<td>7.82</td>
<td>4.17</td>
<td>3.84</td>
<td>11.50</td>
</tr>
</tbody>
</table>

The objective of this study is to analyze cash prices in spatially separate markets in the five major rice producing states for long grain rough rice. Two non-stationary price series are said to have a long run relationship, if none of the series move independently. This is true because they are linked with their stochastic trends which cause the dynamic paths of such variables to have some relation to the current deviation from the equilibrium relationship (Enders). For example, market price will tend to increase in the region when excess demand occurs and it tends to decrease in the

$^2$Unadjusted by transportation cost.
region where excess supply exists. However, the price gaps will tend to get closer as time passes, since traders will ship more of the products from the excess supply region to the deficit regions. Consequently, if the price series from two different regions do not follow the law of one price then their relationships can be described as: (a) there is no arbitrage which links the two markets together, (b) the competition in one or both of the markets may be imperfect, and/or (c) there exist trade barriers or asymmetric information which prevent the market to perform efficiently.

Table 1 showed that over the period of analysis, the average long grain rough rice cash price in California is the lowest compared to that of in other markets. If this is always be the case, theoretically California’s rice producers can ship their production to other regions which pay a higher price for their product. What trade restrictions, if any, prevent the rice growers from doing so? In light to the above discussion, this study lets trade and arbitrage to occur among different markets. In this study, rice interstate shipment is allowed to occur so long as the price differential among different regions is greater than the transportation cost$^3$. By doing so, this study lets the price vary depending on the magnitude of the price differentials and the transportation cost. For instance, if the price in California is lower than that of in Arkansas, the model will let California rice growers to ship their products to Arkansas, so long as the transportation cost is lower than the price differential.

Data and Methods:

Monthly rough rice cash price data from the National Agricultural Statistics Service (NASS) of the USDA are utilized in this study. These price series cover the period from 1986-1998 with a

$^3$The long grain rice price in Arkansas is compared against prices of the other markets. This comparison yields positive or negative margins. These margins are compared with the transportation cost between markets. If the margin in one market is higher then the transportation cost, inter-shipments are allowed to prevail between the two regions. By doing so, the study allows trade to occur between the central market and other markets.
total of 151 observations. NASS quoted and compiled the price series based on industry surveys. Assuming that Arkansas is the central market, the transportation cost is applied to adjust long grain rice cash prices for Texas, Louisiana, Mississippi and California. Arkansas is chosen as the central market for two important reasons. First, Arkansas is by far the largest producer of long grain rice in the U.S. Secondly, the largest rice milling facilities in the U.S. are located in Jonesboro and Stuttgart, Arkansas.

In absence of transportation cost, estimating and testing parameter $\beta = 0$ and $\gamma = 1$ from the OLS regression enables one to test the law of one price. Rejection of both null hypotheses suggest that the price series are not cointegrated. Therefore, there must be market restrictions which prevent the two prices move together in the long run.

The stationarity of the different price series is tested using a Dickey-Fuller (DF) unit root test. The DF tests on the unit root are carried out by estimating the following equation:

$$\Delta Y_t = \alpha + \beta_1 Y_{t-1} + \beta_2 t + \sum \gamma_j \Delta Y_{t-j} + U_t$$

After the unit root test is conducted to determine the level of integration, price series can be tested for cointegration if they are determined to have the same level of integration. In this study, the long-run relationships among the price series are tested using the Engle-Granger procedure, the augmented Dickey-Fuller (ADF) test, Johansen’s test, and vector Error Corrections Model (ECM).

Following Engle and Granger, the cointegration is tested based on the following equation:

$$\hat{U}_t = \hat{\beta}_0 + \hat{\beta}_1 \hat{U}_{t-1} + \hat{\beta}_{1+\ell} \hat{U}_{t-\ell} + \hat{\gamma} \Delta Y_{t-j} + \hat{U}_t$$

Cointegration is evaluated by testing the significance of $\hat{\beta}_1$. Rejection of the null hypothesis implies
that the residual term from the cointegrating regression is not white noise which leads to the conclusion that the price series are cointegrated.

The second test on cointegration is carried out by applying the ADF test on the following equation:

\[
(3) \quad \hat{\Delta}_t \equiv \left( \begin{array}{c} 0 \\ \vdots \\ 0 \end{array} \right) \hat{\Delta} \left( \begin{array}{c} \hat{\upsilon}_t \end{array} \right) \left( \begin{array}{c} \hat{\upsilon}_{t-1} \\ \vdots \\ \hat{\upsilon}_{t-p} \end{array} \right) \left( \begin{array}{c} \hat{\upsilon}_{t-p} \end{array} \right) \hat{\upsilon}_t 
\]

Based on Akaike’s AIC criterion, the optimal lag was determined to be one. As in the Engle and Granger procedure, the cointegration test can be pursued by testing parameter $\mu_1$.

The third test on cointegration of price vectors uses Johansen’s procedure based on maximum likelihood estimators. This tests is conducted based on the following system vector autoregressive equations as shown below.

\[
(4) \quad \hat{\Delta}_t \equiv \left( \begin{array}{c} 0 \\ \vdots \\ 0 \end{array} \right) \hat{\Delta} \left( \begin{array}{c} \hat{\upsilon}_{1,t} \\ \hat{\upsilon}_{2,t} \end{array} \right) \left( \begin{array}{c} \hat{\upsilon}_{1,1} \\ \hat{\upsilon}_{1,2} \end{array} \right) \left( \begin{array}{c} \hat{\upsilon}_{1,1} \\ \hat{\upsilon}_{1,2} \end{array} \right) \hat{\upsilon}_t 
\]

\[
(5) \quad \hat{\Delta}_t \equiv \left( \begin{array}{c} 0 \\ \vdots \\ 0 \end{array} \right) \hat{\Delta} \left( \begin{array}{c} \hat{\upsilon}_{1,t} \\ \hat{\upsilon}_{2,t} \end{array} \right) \left( \begin{array}{c} \hat{\upsilon}_{1,1} \\ \hat{\upsilon}_{1,2} \end{array} \right) \left( \begin{array}{c} \hat{\upsilon}_{1,1} \\ \hat{\upsilon}_{1,2} \end{array} \right) \hat{\upsilon}_t 
\]

Where $\hat{\upsilon}_t$ is a vector of $p$ time-ordered price series, $\hat{\upsilon}$ is a $p$-dimensional vector of random residuals. $\hat{\Delta}$ is the first difference operator and $\hat{\upsilon}$s are $p$-dimensional vectors and matrices of coefficients to be estimated, respectively. $\hat{\upsilon}_1$ and $\hat{\upsilon}_2$ are utilized to construct two likelihood ratio tests which are used to determine the number of unique cointegrating vectors in $\hat{\upsilon}_t$. The first test statistics is known as a trace test while the second is maximum likelihood test which is also known as the eigenvalue test. The trace test evaluates the null hypothesis that there are at most $r$ cointegrating vectors which is
expressed by the following equation:

\[ \mathcal{E}_{\text{trace}} (r) = \mathcal{T} n \sum_{i=1}^{r} \ln (1 - \hat{\lambda}_i) \]

where \( T \) is the number of observations and \( \hat{\lambda}_i \) is the \( p - r \) smallest canonical correlations of \( \leq \) with respect to \( \leq_\alpha \). The eigenvalue test evaluates the null hypothesis that there are exactly \( r \) cointegrating vectors in \( P_t \) which is shown below:

\[ \mathcal{E}_{\text{max}} (r, r\alpha) = \mathcal{T} n \sum_{i=1}^{r} \ln (1 - \hat{\lambda}_i) \]

Since the expression of \( \ln(1) \) is equal to zero, then the expression of \( \ln(1 - \hat{\lambda}_i) \) indeed will be zero, if the price series are not cointegrated. Therefore, the further away the estimated characteristics roots are from zero, the larger the \( \mathcal{E}_{\text{trace}} \) and \( \mathcal{E}_{\text{max}} \) statistics.

The fourth cointegration tests is carried out by estimating the ECM model which is in the form of first order VAR vector augmented with a single error correction terms (\( \hat{\lambda}_{t-1} \)). The signs of the speed of adjustment coefficients are in conformity with convergence towards the long run equilibrium. Assuming that there are three price series, then the ECM model can be expressed as follows:

\[ P_{AR_i} = \% \%
\]

\[ P_{AR_i} = \%
\]

\[ P_{AR_i} = \%
\]

\[ P_{AR_i} = \%
\]

\[ P_{AR_i} = \%
\]

\[ P_{AR_i} = \%
\]

\[ P_{MS_i} = \%
\]
\[ P_{MS,t} = \left(1 - \hat{U}_{AR,t-1} \right) P_{MS,t-1} + \hat{U}_{MS,t} P_{MS,t-1} + \hat{U}_{AR,t-1} P_{AR,t-1} + \hat{U}_{TX,t-1} P_{TX,t-1} \]

The subscripts AR, TX and MS correspond to long grain rice price from Arkansas, Texas and Mississippi, respectively and \( \hat{U}_{AR,t-1} \) refers to the lag of the cointegrating residual terms from Arkansas’ price equation. In this model, the long run equilibrium relationship occurs if the coefficient of the error terms is statistically significance which implies that the change in the dependent variable does response to the deviation of long run equilibrium in period t-1 (\( \hat{U}_{AR} \)) and therefore one can conclude that the vector of price series do have common movements in the long run.

**Results and Discussion**

In this paper, the price relationship in different markets are first tested by estimating the OLS regression where all price series are in log. Based on the estimation, simultaneous and individual tests on parameters " =0 and $=1 are conducted. The results of the least squares estimation and the test of the law of one price are presented in Table 2. Most of the estimates are statistically significant at a one percent confidence level. A simultaneous test of the null hypothesis on parameters " =0 and $=1 is rejected at a one percent confidence level. An individually test on parameter " and $ suggests that Mississippi’s price does have a long run price relationships with that of Arkansas’. Other price
series are tested in pair and the results suggest that Arkansas price does not have a common movement with that of Texas’, Louisiana’s and California’s. Note however, the above tests are conducted without taking into account that the price series do not fulfil the assumption of least square estimator property of a constant variance. Therefore, the OLS results are bias toward rejection of the null hypothesis of no-cointegration.

Examining the graph of monthly long grain rice price movements in the last twelve years, California’s price behaves differently compare to the other markets (Figure 1). While other prices move together, California’s price does not follow the same pattern. This conclusion is confirmed after examining the order of integration in the price series. Unit root tests show that California price series is I(0) or stationary without differences, while the other four prices are stationary after first differences as shown by Dickey Fuller stationarity tests presented in Table 3. Tests on the level showed that long grain rice price in Arkansas, Louisiana, Texas and Mississippi have a unit root.

Having tested and found that the price series are I(1), except for California’s, tests on the long run equilibrium relationships among the price series are conducted using four approaches. These approaches are Johansen’s, Engle and Granger, Augmented Dickey Fuller and ECM tests. The result of these tests are presented in Table 4, 5, 6 and 7, respectively. As shown in Table 4, a bivariate Johansen’s tests on long run equilibrium relationships between Arkansas’ and other price series suggest that the pair of price series do move together in the long run, except for that of Arkansas’ and California’s. This finding is expected to be the case, since California’s and Arkansas’ cash price do not have the same order of integration. On the other hand, both the $\delta_{\text{trace}}$ and the $\delta_{\text{max}}$ tests reject the null hypotheses of no-cointegration for the other price pairs. Arkansas’ and Louisiana’s price series indicate a strong long run price relationship as shown by the significance of both $\delta_{\text{trace}}$ and the
This findings lead us to believe that long grain rough rice cash price in Arkansas do move together with long grain price in Louisiana, Mississippi and Texas. A multivariate Johansen’s tests are also carried out on all price series and the results show that at least there exist three cointegrating price vectors (the results of this test are not reported here).

The same findings are also found by applying Engle and Granger and Augmented Dickey Fuller cointegration tests. As shown in Table 5, the null hypothesis of non stationarity of the cointegrating residuals are rejected at a one percent confidence level in all cases. The stationarity of the residuals from Engle and Granger equation (equation 2) is also tested by the Durbin-Watson statistics. Both of these tests suggest that the error terms are white noise. As shown in Table 5, the DW statistics for each price pair is around two which suggests the absence of serial correlations on the error terms. These findings are consistent with the results of bivariate Johansen’s tests discussed earlier. The same conclusions are also reached after examining the results of the ADF tests as presented in Table 6. The ADF tests suggest a strong rejection of the null hypothesis that the residuals from the ADF equation (equation 3) are not white noise. The results are even more convincing as shown by the Durbin Watson statistics, which are all around two, which reject the null hypothesis of the presence of serial correlation.

Based on the above finding one can conclude that the price series in Arkansas, Mississippi, Texas and Louisiana do move together in the long run. As shown by the result of this study, California’s price does not have a long run relationship with the rest of the other prices. One of the reason which can explain the lack of long run equilibrium relationship between California and Arkansas is that it is not only due to its different order of integration, but it also due to different quality that California’s rice has, even though they are classified in the same category. Consumer’s
preference more on medium or short grain than long grain rice can also be the important source of price disintegration. With a large Asian and Hispanic population in California, most table rice which is consumed in California are medium or short grain. Long grain rice in California is only grown in an insignificant amount (about 0.9 percent, in 1999) compare to other grains. As shown in Table 1, over the period of analysis, not only that the mean price in California is the lowest ($6.18/cwt), but it is also the most stable as measured by its variance (1.67). The stable price in California could have been caused by market niches found by producers which sell their product by contract with the end users which reduce price variability. Sumner and Lee argued that in more recent years, California’s rice has been available at a ten percent or more price discount from the Southern long grain price. However, they did not elaborate the reason behind the price discount.

The ECM estimation showed that the cointegrating error terms in all models are significant at a one percent confidence level (Table 7). The model is estimated without California price in it, since it has different order of integration. In this study, variable which will be modeled in the cointegrating regression is selected based on how significance its contribution in explaining the variability in dependent variable (Arkansas’ price). The process of selecting variable is based on backward selection procedures available in SAS. During the selection process, California price is also included in the full model. From the Multivariate Johansen’s cointegration tests, this study found that the null hypotheses of no-cointegration are rejected for up to three price vectors. This prior information is useful to decide how many cointegrating variables need to be included in the final cointegrating regression. Our interest is to select the best three possible combinations of explanatory variables which meet the significance criteria chosen and imposed in the selection process. The results show that California’s rice price can not meet the criteria as does the other three explanatory
variables. Therefore, California’s price is dropped from the final ECM model. The model is estimated in first difference augmented by cointegrating error terms.

Three price series survive the selection procedure and these price series are from Louisiana, Texas and Mississippi. Therefore, they are included in the cointegrating regression\(^4\). The results also showed that the cointegrating residual terms \(\hat{U}_{AB - 1}\) in all price equations in the ECM model are significant as shown by the t-statistics. This result supports the earlier long run equilibrium tests using the other three tests. Therefore, one can conclude that there are three cointegrating price vectors and these price vectors are Arkansas’ price against Texas, Mississippi’s and Louisiana’s. As can be seen in Table 7, these price series response mostly to a negative discrepancies in \(\hat{U}_{AB - 1}\). In all cases, the coefficient of the lagged of the error terms are negative. Long grain price in Mississippi, Texas and Louisiana tend to response negatively to the negative discrepancies in \(\hat{U}_{AB - 1}\).

Conclusion

Given the importance of long grain rice in the U.S. in more recent years, the performance of rough rice market is of interest to many market players, including traders, producers and consumers. This analysis utilized long grain rough rice cash prices to assess price relationships among different markets in the U.S. rice industry such as in Arkansas, Texas, Louisiana, Mississippi and California. The results of stationary tests suggested that four rice price are I(1), not including California’s which is I(0). This might suggest that California differs from the rest of the other markets. Within four

\[^4\text{The estimated cointegrating regression and standard deviation are as follows:}\]

\[
P^{AR} = 0.12 + 0.32 P^{LA} + 0.18 P^{TX} + 0.45 P^{MS} + \hat{U}_{AB - 1} \\
\text{Prob } > \text{F} = 0.0001 \quad \text{R}^2 = 0.93
\]
cointegration tests, this study found that rough rice cash price in Arkansas and the price in Texas, Mississippi and Louisiana are I(1,1) which suggests that they do have a long run price relationships. The fact that four of these price series are cointegrated leads one to conclude that the market in these regions operate efficiently such that players are well informed about market conditions so that none of them has the power to affect price and make abnormal profit. However, California long grain rough rice market seems to operate differently with the rest of the other markets in the U.S. Perhaps, the quality factors and the characteristics of long grain rice demand in California is different with those of in other markets. These distinct characteristics perhaps are important factors which can explain its unique behavior.
Table 2 - OLS Cointegration Tests ($P^A = " + $ P^B$)

<table>
<thead>
<tr>
<th>Independent Variables</th>
<th>Texas</th>
<th>Mississippi</th>
<th>Louisiana</th>
<th>California</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$P^{TX}$</td>
<td>$P^{MS}$</td>
<td>$P^{LA}$</td>
<td>$P^{CA}$</td>
</tr>
<tr>
<td>$&quot;$</td>
<td>-0.34</td>
<td>0.05</td>
<td>-0.21</td>
<td>-0.50</td>
</tr>
<tr>
<td></td>
<td>(-6.61)**</td>
<td>(1.06)</td>
<td>(-4.80)**</td>
<td>(-2.22)*</td>
</tr>
<tr>
<td>$$</td>
<td>1.11</td>
<td>0.97</td>
<td>1.06</td>
<td>1.18</td>
</tr>
<tr>
<td></td>
<td>(45.82)**</td>
<td>(46.10)**</td>
<td>(51.96)**</td>
<td>(11.24)**</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.93</td>
<td>0.93</td>
<td>0.95</td>
<td>0.46</td>
</tr>
<tr>
<td>$F$-statistics$^b$</td>
<td>155***</td>
<td>5.37***</td>
<td>85.92***</td>
<td>21.69***</td>
</tr>
<tr>
<td>$F$-statistics$^c$</td>
<td>20.57***</td>
<td>2.28</td>
<td>9.86***</td>
<td>2.92*</td>
</tr>
<tr>
<td>$F$-statistics$^d$</td>
<td>43.67***</td>
<td>1.13</td>
<td>23.01***</td>
<td>4.94**</td>
</tr>
</tbody>
</table>

$^a$Numbers in parenthesis are t-statistics. $***$, $**$ and $*$ respectively are significant at 1, 5 and 10 % confidence level. For a sample of size 120, the critical t-value at 1, 5 and 10 % respectively are 2.358, 1.657 and 1.289.

$^b$Based on the following hypothesis: $H_0$: $E$=1 and "$=0.

$^c$Based on the following hypothesis: $H_0$: $E$=1.

$^d$Based on the following hypothesis: $H_0$: "$=0.
Figure 1 - Long Grain Rice Cash Price (1986-1998)
Table 3 - Dickey Fuller Tests on Stationary Condition of Various Long Grain Rice Price Series

<table>
<thead>
<tr>
<th>Null hypothesis for unit roots (^{a})</th>
<th>ADF Statistics</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Arkansas</td>
</tr>
<tr>
<td>(\tau_{\mu}, H_0 : (\beta_1 = 0</td>
<td>\beta_2 = 0))</td>
</tr>
<tr>
<td></td>
<td>-2.28</td>
</tr>
<tr>
<td>(\tau_{\tau}, H_0 : (\beta_1 = 0))</td>
<td>-3.21*</td>
</tr>
<tr>
<td>(\phi_1, H_0 : [\alpha, \beta_1 = (0,1)</td>
<td>\beta_2 = 0])</td>
</tr>
<tr>
<td>(\phi_2, H_0 : (\alpha, \beta_1, \beta_2 = (0,0,0))</td>
<td>3.68</td>
</tr>
</tbody>
</table>

\(^{a}\)A number of different tests have been suggested by Dickey and Fuller (p. 1062-63, 1981) and Fuller (p. 373, 1976), based on the following model:

\[
\Delta Y_t = \alpha + \beta_1 Y_{t-1} + \beta_2 t + \sum \gamma_j \Delta Y_{t-j} + U_t
\]

At the 10%, 5% and 1% confidence levels, the critical values of the \(t(J_{\mu} \text{ and } J_{j})\) and F-statistics \(M_i \text{ and } M_j\) for a sample size of 100 are, respectively, \(J_{\mu} : -2.58, -2.89, \text{ and } -3.51; J_{j} : -3.15, -3.45 \text{ and } -4.04; M_i : 3.86, 4.71 \text{ and } 6.70; M_j : 4.16, 4.88 \text{ and } 6.50. \) In either case, the null hypothesis is rejected if the calculated pseudo \(t\) or F-statistics are greater than the critical \(t\) or F-values as reported by Fuller (1976) and Dickey and Fuller (1981). To clarify, \(M_i\) is based on the above equation but without the time trend in the equation. One, two and three stars indicate rejection of the null hypothesis of a unit root (non-stationary) at ten, five and one percent confidence level.
Table 4 - Trace and Maximum Eigenvalues Statistics
Using Johansen's Cointegration Tests

<table>
<thead>
<tr>
<th>Null Hypothesis</th>
<th>Alternative Hypothesis</th>
<th>$\lambda_{\text{max}}$ tests</th>
<th>$\lambda_{\text{trace}}$ value</th>
<th>$\lambda_{\text{max}}$ value</th>
<th>95% Critical value</th>
<th>99% Critical value</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$P^{\text{AR}}$ vs $P^{\text{TX}}$</td>
<td>$P^{\text{AR}}$ vs $P^{\text{MS}}$</td>
<td>$P^{\text{AR}}$ vs $P^{\text{LA}}$</td>
<td>$P^{\text{AR}}$ vs $P^{\text{CA}}$</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$r=0$</td>
<td>r &gt; 0</td>
<td>34.10**</td>
<td>35.24**</td>
<td>47.35***</td>
<td>21.37</td>
<td>29.68</td>
</tr>
<tr>
<td>$r &gt; 1$</td>
<td></td>
<td>6.89</td>
<td>4.656</td>
<td>15.41</td>
<td>20.04</td>
<td></td>
</tr>
<tr>
<td>$r#1$</td>
<td>r &gt; 1</td>
<td>5.31</td>
<td>5.57</td>
<td>14.07</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$r=1$</td>
<td>r = 1</td>
<td>28.79***</td>
<td>29.67***</td>
<td>40.46***</td>
<td>16.72</td>
<td>20.97</td>
</tr>
<tr>
<td>$r &gt; 2$</td>
<td></td>
<td>6.89</td>
<td>4.66</td>
<td>14.07</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Sources: Osterwald-Lenum (1992). *** and ** denote to rejection of the null hypothesis of no cointegration at a one and five percent confidence level. The optimum number of two lags are chosen based on Akaike’s AIC criteria.
The test is based on the following equation:

\[ \hat{U}_t = \mu_0 + \beta_{\text{lag}} V_t + \epsilon_t \]

Rule of thumb suggests the rejection of null hypothesis of no autocorrelation, if DW statistics is around 2.

Criteria for selecting the optimum length of lags.

Three and two stars refer to rejection of the null hypothesis that the cointegrating error terms \((h_{t-1})\) have a unit root. Stationarity implies that the price series are integrated of order \((1,1)\). Engle and Yoo (1987) report the critical t statistics as -4.00 and -3.37.

Test on the significant of \("_1 \) based on the following equation:

\[ \hat{U}_t = \mu_0 + \beta_{\text{lag}} \sum_{j=1}^{p} \hat{U}_{t-j} + \epsilon_t \]

---

**Table 5 - Engle-Granger Cointegration Tests**

<table>
<thead>
<tr>
<th>Price Pairs</th>
<th>t - Statistics</th>
<th>DW-Statistics</th>
<th>Akaike’s AIC</th>
<th>( R^2 )</th>
</tr>
</thead>
<tbody>
<tr>
<td>( P_{AR} ) vs ( P_{TX} )</td>
<td>-7.45***</td>
<td>2.32</td>
<td>0.28</td>
<td>0.27</td>
</tr>
<tr>
<td>( P_{AR} ) vs ( P_{MS} )</td>
<td>-7.30***</td>
<td>2.19</td>
<td>0.29</td>
<td>0.26</td>
</tr>
<tr>
<td>( P_{AR} ) vs ( P_{LA} )</td>
<td>-8.16***</td>
<td>2.05</td>
<td>0.23</td>
<td>0.31</td>
</tr>
</tbody>
</table>

**Table 6 - Augmented Dickey Fuller Cointegration Tests**

<table>
<thead>
<tr>
<th>Price Pairs</th>
<th>t - Statistics</th>
<th>DW-Statistics</th>
<th>Akaike’s AIC</th>
<th>( R^2 )</th>
</tr>
</thead>
<tbody>
<tr>
<td>( P_{AR} ) vs ( P_{TX} )</td>
<td>-4.19***</td>
<td>2.01</td>
<td>0.25</td>
<td>0.36</td>
</tr>
<tr>
<td>( P_{AR} ) vs ( P_{MS} )</td>
<td>-4.56***</td>
<td>1.99</td>
<td>0.28</td>
<td>0.29</td>
</tr>
<tr>
<td>( P_{AR} ) vs ( P_{LA} )</td>
<td>-4.82***</td>
<td>1.97</td>
<td>0.23</td>
<td>0.32</td>
</tr>
</tbody>
</table>

5The test is based on the following equation:

6Rule of thumb suggests the rejection of null hypothesis of no autocorrelation, if DW statistics is around 2.

7Criteria for selecting the optimum length of lags.

8Three and two stars refer to rejection of the null hypothesis that the cointegrating error terms \((h_{t-1})\) have a unit root. Stationarity implies that the price series are integrated of order \((1,1)\). Engle and Yoo (1987) report the critical t statistics as -4.00 and -3.37.

9Test on the significant of \("_1 \) based on the following equation:
Multivariate Johansen’s cointegration procedures are estimated and the null hypothesis of no-cointegration for three pairs of price series is rejected (the results are not reported here). The question becomes which of the combination of price vectors will generate the most stationary residual terms? To encounter such a problem, this study applies model selection procedures which available under Pro Reg (SAS). Based on such an approach, at each step of the elimination procedure, the explanatory variable which contributes the least is deleted based on the F-statistics criteria.

**Table 7 - Estimation Results on Long Grain Rice Price Equations**

<table>
<thead>
<tr>
<th>Independent Variable</th>
<th>Arkansas</th>
<th>Mississippi</th>
<th>Texas</th>
<th>Louisiana</th>
</tr>
</thead>
<tbody>
<tr>
<td>Intercept</td>
<td>-0.12</td>
<td>0.48</td>
<td>0.36</td>
<td>0.63</td>
</tr>
<tr>
<td></td>
<td>(-2.08)**</td>
<td>(9.85)***</td>
<td>(7.52)***</td>
<td>(18.47)***</td>
</tr>
<tr>
<td>( \hat{U}_t - 1 )</td>
<td>-2.41</td>
<td>-0.72</td>
<td>-0.50</td>
<td>-0.71</td>
</tr>
<tr>
<td>( -(10.02) ***</td>
<td>(-3.71)***</td>
<td>(-2.61)***</td>
<td>(-5.25)***</td>
<td></td>
</tr>
<tr>
<td>( P_{AR,t-1} )</td>
<td>1.19</td>
<td>0.43</td>
<td>0.74</td>
<td>0.74</td>
</tr>
<tr>
<td></td>
<td>(22.02)***</td>
<td>(9.87)***</td>
<td>(17.31)***</td>
<td>(24.17)***</td>
</tr>
<tr>
<td>( P_{MS,t-1} )</td>
<td>-0.24</td>
<td>0.03</td>
<td>-0.58</td>
<td>0.19</td>
</tr>
<tr>
<td></td>
<td>(-2.82)***</td>
<td>(0.40)</td>
<td>(-8.74)***</td>
<td>(3.94)***</td>
</tr>
<tr>
<td>( P_{TX,t-1} )</td>
<td>0.95</td>
<td>1.53</td>
<td>0.05</td>
<td>0.83</td>
</tr>
<tr>
<td></td>
<td>(8.56)***</td>
<td>(17.11)***</td>
<td>(0.52)</td>
<td>(13.19)***</td>
</tr>
<tr>
<td>( P_{LA,t-1} )</td>
<td>-0.88</td>
<td>-0.73</td>
<td>0.49</td>
<td>-0.46</td>
</tr>
<tr>
<td></td>
<td>(-8.76)***</td>
<td>(-9.02)***</td>
<td>(6.15)***</td>
<td>(-8.15)***</td>
</tr>
<tr>
<td>( R^2 )</td>
<td>0.84</td>
<td>0.90</td>
<td>0.93</td>
<td>0.97</td>
</tr>
</tbody>
</table>

Numbers in parenthesis are t-statistics. ***, ** and * respectively are significant at 1, 5 and 10 % confidence level. For a sample of size 120, the critical t-value at 1, 5 and 10 % respectively are 2.358, 1.657 and 1.289.


