MARKET STRUCTURE AND SPATIAL PRICE DYNAMICS

B. Wade Brorsen, Jean-Paul Chavas, and Warren R. Grant

Abstract

A method was developed with time series models to test hypotheses about the relationship between market structure and spatial price dynamics. Long-run dynamic multipliers measuring the magnitude of lagged adjustment for spatial milled rice prices were calculated from the time series model and used as the dependent variable in a regression model that included a number of factors expected to influence price determination. Results show that price adjustments were slower as regional submarket concentration increased and were faster in the regions with a higher market share. Arkansas, the state with the largest market share, was consistently a price leader.

Key words: rice, market structure, spatial pricing

INTRODUCTION

Economists have long been interested in the effect of market structure on equilibrium prices (e.g., Means; Carlton). Several recent articles have examined the relationship between market structure and price dynamics (e.g., Bedrossian and Moschos; Ginsburg and Michel; Kardasz and Stollery). The empirical tests in these studies have used similar data and procedures, relying on a two-step estimation procedure. The lag in dynamic adjustment is typically estimated with time series data for each of a number of industries. A measure of the speed of adjustment for each industry is then used as a dependent variable in a regression against measures of market structure. This study used a two-step procedure but relied on a different measure of the dynamic adjustment. Both Bedrossian and Moschos and Kardasz and Stollery followed past literature and used a partial adjustment model to capture the dynamics, which resulted in an unnecessarily restrictive model. The significant autocorrelation found in the second step of the analysis in both studies could be interpreted as a sign of dynamic misspecification. Time series models are an alternative providing a more flexible model of dynamics. Also, the procedures used in past studies were designed for unrelated markets. The approach used here was appropriate for related markets. Also, past research deals with firm or industry data and has not looked at the effect of market structure on prices in subregions.

Several recent papers have used time series models to study the speed and direction of dynamic adjustments of prices across locations (e.g., Spriggs et al.; Brorsen et al.). Researchers have only been able to investigate whether the prices in one location lead or lag those in another. They then could only speculate about why prices adjusted sooner in one location than in another. For example, Spriggs et al. argued that market structure could explain why wheat prices in the larger U.S. market tended to lead prices in the smaller Canadian market. Past researchers using time series models were unable to test hypotheses about the reasons prices in one market led those in another. The speed of price adjustments should be of concern to antitrust regulators because firms must receive price signals quickly and accurately if they are to make economically efficient decisions (Sporleder and Chavas).

OBJECTIVE

A new procedure to test structural hypotheses about spatial price adjustments using time series models is explained in this paper. The procedure was used to investigate the behavior of milled-rice prices in submarkets of selected U.S. regions. The relative size of the submarkets and submarket seller concentration were found to affect the lead-lag relationships of spatial milled-rice prices.

PROCEDURE

The analysis proceeds in two steps. In the first step, a vector autoregressive (VAR) model of the rice prices is estimated. Dynamic multipliers measuring the magnitude of the delayed price adjustments are calculated from the estimated VAR. In a second step, these multipliers are used as the dependent variable in a regression model investigating the effects of factors influencing the price adjustments.

This procedure allows testing both of the price administration and of the price leadership hypotheses. As used here, the price leadership hypothesis suggests that prices in regions with a larger market share will lead those in markets with a smaller

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market share. The price administration hypothesis suggests that prices in more concentrated markets are sluggish. As applied here, this means that prices in regions with low concentration should lead prices in regions with high concentration.

REASONS FOR LAGGED ADJUSTMENTS

The efficient markets hypothesis (Fama; Samuelson) suggests that prices in perfectly competitive markets should be statistically indistinguishable from a random walk or martingale process.[1] In addition to the standard assumptions of perfect competition, this hypothesis assumes no transaction costs, risk neutral traders, and zero information costs (Danthine). These assumptions are not valid for cash commodity markets where transaction costs (e.g., transportation and storage) can be high. The random walk model for cash commodity prices has been consistently rejected in empirical studies (e.g., Brorsen et al. 1985b; Spriggs et al.). Thus, lags in price adjustment were expected in the study. Possible reasons for prices in one market to lead those in another are discussed next.

Market structure may influence price dynamics. The "administered price" hypothesis (Means) states that output prices in more concentrated industries are less flexible (i.e., change less frequently and by smaller relative magnitude) than in less concentrated markets. In other words, the hypothesis states that seller concentration tends to have a negative influence on the speed of price adjustment. Most previous research investigating the administered price hypothesis using U.S. data has relied on the Bureau of Labor Statistics wholesale price index series (e.g., Depodwin and Selden; Weiss; Ripley and Segal; Qualls). Although this research obtained conflicting results, the most recent work by Carlton and Ross and Krausz found evidence supporting the administered price hypothesis.

Attempts have been made to measure directly the speed of price adjustments. Domberger and Kardasz and Stollery have presented empirical evidence that speed of adjustment is positively related to concentration. However, Dixon and Bedrossian and Moschos found empirical evidence showing the opposite result. Bedrossian and Moschos argued that either a positive or negative relationship is possible. Ginsburgh and Michel presented a theoretical model and argued that net effects in either direction are possible.

In less than perfectly competitive markets, firms (or submarkets as in this paper) may benefit by some form of short-run pricing coordination. One way of coordinating behavior is (barometric) price leadership, where the largest seller is typically the price leader acting as a barometer of market conditions (e.g., Markham).

Because price leadership is characteristic of concentrated markets, the presence of high concentration suggests both the possibility of price leadership and administered prices. Both hypotheses were tested in this study, but they were tested at different levels of aggregation. When considering regional submarkets, market structure can be evaluated within each submarket, or across submarkets in the context of a national or international market. Rice is largely milled in four states, each state having a small number of firms. Thus, rice is a concentrated industry both within each submarket and across regions. The administered price hypothesis is concerned with the negative influence of seller concentration on output price flexibility (Means; Carlton; Ross and Krausz). When applied to regional submarkets, this would suggest a negative relationship between submarket seller concentration and submarket price flexibility. Thus regional markets are treated in the same way as industries in a more traditional analysis.[2]

Since rice milling is concentrated in milling centers, price leadership may exist across regions. Holthausen (p. 341) argued that "the largest firms with the greatest profit cushion against possible loss can best afford the role of price leader." Substituting regional output for firm output, the price leadership hypothesis would thus indicate that regions with higher market shares would be price leaders. In other words, treating submarkets as if they were single firms in a more traditional analysis, the relative size of each submarket is the relevant measure in the analysis of the price leadership hypothesis. Thus, the effects of price leadership (relative market share) and administered pricing (concentration) involve different levels of aggregation in the context of regional submarkets and lead to different testable hypotheses.

Within the United States, the rice production areas of Louisiana and Texas are contiguous, but their milling centers are not, so the state boundaries are a reasonable definition of regional submarkets. Thus, the analysis of milled rice prices was conducted for

[1] The martingale model implies the expected price at time t conditional on the information available at t-1 is the price at time t-1.

[2] In the context of a national market, (rather than submarket) seller concentration measures could also be used to investigate the administered price hypothesis. However, this would require using prices different from the regional prices used in this study.

four milling centers (in Arkansas, California, Louisiana, and Texas), each center constituting a submarket.

Each milling center sells its product on national and international markets where it competes with each of the other milling centers. Centers that are in close competition would be expected to exhibit similar pricing patterns. Accordingly, the cost of transportation of rice from one center to another was used as a proxy for the closeness of competition among centers. The direction of the effect of transportation costs on lagged adjustments cannot be determined *a priori*. As transportation costs between submarkets increase, the degree of competition could decrease and the (total) effect of one submarket's price on another can be expected to decrease. However, as transportation costs increase, this total cross submarket effect may become more lagged and less instantaneous. Thus, the net lagged price effects across locations may either increase or decrease with transportation costs. These hypotheses are empirically investigated next.

EMPIRICAL METHODOLOGY

Because of agro-climatic factors, U.S. rice production is concentrated in just a few regions. Weekly milled-rice prices were analyzed in four rice milling centers in the United States: Stuttgart (Arkansas), Sacramento (California), Lake Charles (Louisiana), and Houston (Texas). Rice processed in these four centers along with Mississippi[3] account for 99 percent of the rice grown in the United States. Each center processes most of the rice produced in the corresponding state. The largest firms tend to operate in only one state. Several firms are cooperatives, but they should not behave any differently from private firms in selling rice. The weekly prices are for the dominant grain type in each location. The prices selected are for U.S. No. 2 long grain in Arkansas and Texas, U.S. No. 2 medium grain in Louisiana, and U.S. No. 1 medium grain in California. The data are from *Rice Market News* (U.S. Department of Agriculture). The prices used are from the midpoint of the range of quoted prices from mills surveyed. The prices quoted by mills are transaction prices if a recent transaction has occurred, otherwise they are offer prices.

The VAR modeling of the rice prices was conducted on seasonally adjusted weekly price changes during two five-year time periods: August 1975-July 1980, and August 1980-July 1985. The analysis of each period separately was motivated by the evidence of structural change found in the empirical analysis (see below). Price changes (first differences) were used and thus each time series was assumed difference stationary. Seasonality was not removed. Seasonality was a small part of weekly price changes and adjusting for seasonality made little difference, so the simpler model was reported.

The estimation of the time series model and of the long-run multipliers is explained next. Long-run multipliers measure the lagged portion of the total change in one price due to a change in another price. If more of the total effect is lagged (i.e., long-run multipliers are large), then adjustments are slower.

The (dx1) vector of stationary time series denoted by $y_t = (y_{1t}, ..., y_{dt})'$. Here, d=4 and y_t is the vector of deseasonalized milled-rice price changes at the four milling centers (Arkansas, California, Louisiana, and Texas) during week t. The VAR model can be written as

$$(1)\quad y_t = \sum_{j=1}^{p} B_j Y_{t-j} + e_t,$$

where p is the order of the autoregressive process generating the vector y_t, B_j is a (dxd) matrix of parameters and e_t is a white noise error vector. The order p of the VAR model was identified using Akaike's Information Criterion (AIC) and the parameters in (1) were estimated by least squares.[4]

Following Chow, equation (1) was rewritten as a first-order difference equation:

$$(2)\quad y_t^* = Ay_{t-1}^* + e_t^*$$

where

$$y_t^* = \begin{bmatrix} y_t \\ : \\ y_{t-p+1} \end{bmatrix},\ A = \begin{bmatrix} B_1 & \dots & B_p \\ I_{d(p-1)} & : & 0 \end{bmatrix},\ e_t^* = \begin{bmatrix} e_t \\ z \end{bmatrix},$$

$I_{d(p-1)}$ is a d(p-1) x d(p-1) identity matrix, 0 is a d(p-1) x d matrix of zeroes, and z is a d(p-1) x 1 vector of zeroes.

Only the first d rows of A contain parameters to be estimated. Accordingly, let A_d be the first d rows of A and let $\tilde{A}_d$ be consistent estimates of the first d

[3] Greenville, Mississippi recently became a major rice milling center. No milled price series are available, however, for the period covered in this report.

[4] Since the explanatory variables are identical in each equation, ordinary least squares estimates of the paramaters are identical to seemingly unrelated regression estimates. Then, conditional on choosing the correct number of lags, the least squares estimates of the a_j's are consistent and asymptotically efficient (Judge et al., p. 680).

rows of A. Then

(3) $Q = E[\,vec(\tilde{A}_d - A_d)\, vec(\tilde{A}_d - A_d)'],$

was defined so that vec is the column stacking operator. By the central limit theorem, it followed that

(4) $\sqrt{T}\,[\,vec(\tilde{A}_d) - vec(A_d)] \xrightarrow{d} N(O,Q),$

where T denoted the number of observations.

The estimated parameters can be used to investigate the dynamics of price adjustments. For that purpose, intermediate-run multipliers can be obtained from the VAR model (1). Intermediate run multipliers were used to measure the combined effects on all prices of a shock through the error term e_t that persists over m periods. The patterns displayed by the estimated m-period intermediate run multipliers can provide useful information on price dynamics. After considering alternative summary measures from these intermediate-run multipliers, long-run multipliers were selected as a simple measure of lagged price adjustments. Long-run price adjustments are the limit (if it exists) of the corresponding m-period intermediate run multiplers as m approaches infinity.

Thus, the goal was to calculate long-run multipliers from the estimated VAR model (see Chow), along with the asymptotic standard errors (Dhrymes). The estimated long-run multipliers were analyzed to investigate the nature of price dynamics. This approach was chosen among several alternatives. For example, Nachane et al., when faced with a similar problem of measuring the strength of the lagged adjustment, opted for Geweke's (1982, 1984) measures of causality strength. Compared to Geweke's approach, multipliers have the advantage of indicating both the direction (sign) and measuring the strength of lagged effects. Also, the asymptotic statistical distribution is known. Nachane et al. were forced to use nonparametric techniques in cross-section regressions. Schroeder and Goodwin used F-statistics that are similar to Geweke's measure of causality strength. Another alternative would be to use the impulse response function of Sims. However, Sims' approach is sensitive to the ordering of equations and does not readily provide a summary measure of the lagged response. The approach used in this study does not consider contemporaneous correlation, because the hypotheses of interest were only about lagged effects.

The long-run multipliers are calculated as

(5) $\pi = (I_{dp} - A)^{-1}\, C,$

where $C = \begin{bmatrix} I_d \\ Z \end{bmatrix}$, Z is a d(p-1) x d matrix of zeroes, and I_{dp} and I_d are identity matrices with dimensions dp x dp and d x d. This follows Brorsen et al. (1985a). Such multipliers measure the impact of a change in a particular price at time t (associated with a shock occurring in the error term) on the future adjustments of all prices. Assuming that the VAR model is stationary, long-run multipliers (denoted here by π_{ij}) measure the lagged effects of a shock in the i^{th} price on the expected value of the j^{th} price after a new equilibrium is reached (Brorsen et al. 1985a, p. 366). Because the prices considered involve substitute commodities, one can expect $\pi_{ij} \geq 0$, reflecting the hypothesis that regional prices tend to move together. A significant multiplier implies Granger causality.

Next, long-run multipliers π_{ij}'s, were used as measures of the magnitude of lagged price adjustments across markets ($i \neq j$) reflecting the lead-lag relationships of one price over another. As argued above, the long-run multipliers provide simple summary measures of the importance of delayed price adjustments across locations.[5]

Brorsen derived the distribution of the long-run multipliers for this particular case based on the more general case developed by Dhrymes. Define $\tilde{\pi} = (I - \tilde{A})^{-1}\, C$ and let $D = (I - A)^{-1}$. Brorsen (p. 71) showed that

(6) $\sqrt{T}\, vec(\tilde{\pi}_d - \pi_d) \xrightarrow{d} N\,[O,(\pi_d' \otimes D)\, Q(\pi_d \otimes D')],$

where $\otimes$ denotes the Kronecker product. The interest here was in π_d, the first d rows of π, because of interest in the changes in Y_t and not its lagged values (the long-run multipliers for Y_t and Y_{t-i}, i=1, ..., p are the same). By substituting consistent estimates of π_d, D, and Q into equation (6), the variance covariance matrix of the estimated multipliers Ω = var $(vec(\tilde{\pi}))$ can be consistently estimated as

(7) $\tilde{\Omega} = (\tilde{\pi}_d' \otimes \tilde{D})\, \tilde{Q}\, (\tilde{\pi}_d \otimes \tilde{D}')/T.$

In order to investigate the factors influencing spatial price dynamics (see previous section), the fol-

[5] This approach could be used similarly with a univariate time series model (or transfer function model, if data on exogenous variables are available) for each price series.

lowing model was specified:

(8) $\pi_{ijt} = \alpha_0 + \alpha_1 QR_{ijt} + \alpha_2 TC_{ijt} + \alpha_3 CR_{ijt} + \alpha_4 t + \mu_{ijt}$, $i,j = 1,...,4, i \neq j$,

where t = 1 denotes the 1975-1980 sample period and t = 2 denotes the 1980-1985 sample period. In equation (8), QR_{ijt} measures relative volume as the ratio of the quantity of milled rice shipped by state i to the quantity of milled rice shipped by state j in period t, TC_{ijt} is the cost of shipping rice from state i to state j, CR_{ijt} is the ratio of the market share of the top three firms in state i to the market share of the top three firms in state j, μ_{ijt} is a random error term with mean zero, and the α's are parameters. Here, the variable CR is a measure of relative submarket seller concentration as suggested by the administered price hypothesis. Also, the variable QR is a measure of relative size of the relevant submarkets as suggested by the price leadership hypothesis. Finally, as discussed above, transportation cost TC is a proxy of the closeness of competition among submarkets.

The off-diagonal elements of the π matrix are denoted by $\bar{\pi}_d$. Note that equation (8) involves the true long-run multipliers in (5). However, these true multipliers were not known; they were only estimated by $\tilde{\bar{\pi}}_d$, the consistent estimates of $\bar{\pi}_d$ derived from the VAR model (1) as discussed above. Having noted that $vec(\tilde{\bar{\pi}}_d) = vec(\bar{\pi}_d) + \varepsilon$, where ε has zero mean asymptotically, it follows that least squares estimation of (8) using estimated rather than actual multipliers provides consistent estimates of the parameters. But least squares produce inconsistent estimates of the standard errors in general. Consistent estimates of the standard errors and asymptotically more efficient estimates of the parameters are available by using the estimated variance of $\tilde{\pi}$. This study defined

$$\Phi = \begin{bmatrix} var(vec(\tilde{\bar{\pi}}_{d1})) & 0 \\ 0 & var(vec(\tilde{\bar{\pi}}_{d2})) \end{bmatrix},$$

where $vec(\tilde{\bar{\pi}}_{d1})$ and $vec(\tilde{\bar{\pi}}_{d2})$ are the estimated long-run multipliers for periods one and two.

Equation (8) was rewritten in matrix form as

$$vec(\bar{\pi}_d) = X\beta + \mu,$$

where X is a matrix of independent variables and $\beta = [\alpha_0,\alpha_1,\alpha_2,\alpha_3,\alpha_4]'$. Since $vec(\tilde{\bar{\pi}}_d) = vec(\bar{\pi}_d) + \varepsilon$, it followed that

(9) $vec(\tilde{\bar{\pi}}_d) = X\beta + V$,

where $V = \varepsilon + \mu$ is the residual vector. Assuming that ε and μ are uncorrelated and that μ is distributed $N(0, \sigma_\mu^2)$, then

(10) $E(V V') = \Sigma = \sigma_\mu^2 I + \Phi$.

Since consistent estimates of Φ can be obtained (say, $\tilde{\Phi}$), σ_μ^2 can be consistently estimated as

(11) $\tilde{\sigma}_\mu^2 = TRACE(\tilde{V}\tilde{V}' - \tilde{\Phi})/T$,

where $\tilde{\sigma}_\mu^2$ is the estimated variance, $\tilde{V}$ are consistent estimates of the residuals obtained from least squares estimation of (9), and T is the total number of observations. In small samples, the estimated variance from equation (11) could be negative. For this reason, Saxonhouse suggested as an alternative the approach used by Bedrossian and Moschos. Saxonhouse pointed out that Φ will converge to zero asymptotically and thus his procedure corresponds to estimating σ_μ^2 as $TRACE(\tilde{V}\tilde{V}')/T$. Here, estimates of σ_μ^2 were calculated using both equation (11), which can produce unbiased estimates of σ_μ^2 with a degrees of freedom adjustment, and Saxonhouse's consistent approach. In either case, the estimated variance-covariance matrix ($\tilde{\Sigma}$) is

Table 1. Quantities Shipped, Transportation Costs, and Concentration Ratios for U.S. Milled Rice

State	Annual Average Quantity Shipped		Transportation Costs Destination				Concentration Ratio (1980)
	1975-79	1980-84	LA	AK	CA	TX	
	--- million cwt. ----		-----------dollars per cwt.-----------				percent
LA	11.2	10.9	—	1.06	3.18	1.37	39
AK	24.1	28.7	1.06	—	3.18	1.06	82
CA	17.7	20.4	3.66	3.66	—	3.66	81
TX	22.7	20.4	0.97	1.06	3.18	—	87

Sources: Holder; Rice Miller's Association.

(12) $\tilde{\Sigma} = (\tilde{\sigma}^2_\mu I + \tilde{\Phi})$,

which can be used in a generalized least squares estimation of (9):

$$(13)\ \tilde{\beta}_{GLS} = (X' \tilde{\Sigma}^{-1} X)^{-1} X' \tilde{\Sigma}^{-1} \begin{bmatrix} vec(\tilde{\bar{\pi}}_{d1}) \\ \\ vec(\tilde{\bar{\pi}}_{d2}) \end{bmatrix}.$$

The resulting parameter estimates are consistent and asymptotically more efficient than ordinary least squares estimates.[6]

The data for the explanatory variables in (8) are exhibited in Table 1. The changes in quantity milled in the two periods was not large. Transportation costs are for 1982 and concentration ratios are for 1980. They could only be obtained for one point in time and were assumed to be the same across the two periods. Thus, these variables only capture cross-sectional differences. Distance was considered as an alternative to transportation costs and the conclusions were unchanged. As shown in Table 1, transportation costs were not always symmetric and costs were much higher for California. Submarket seller concentration was high for all states except Louisiana. (This high concentration can be partially explained by significant economies of scale in rice milling.)

The AIC criterion identified an order (p) of 4 for the 1975-1980 period, and of 9 for the 1980-1985 period.[7] The null hypothesis that the parameters were the same across the two periods was rejected (the estimated x^2 value was 509 with 144 degrees of freedom). This test result was used to justify the analysis of each period separately. The residuals of the VAR model were tested for white noise with Fisher's Kappa and Bartlett's Kolmogorov-Smirnov test statistics (Fuller). The null hypothesis of white noise could not be rejected for any of the equations using either test.

RESULTS

The long-run multipliers which are calculated from the estimated VAR models are presented in Table 2. Long-run multipliers measure the lagged effects of a shock in price i on the expected value of price j. For example, the multiplier showing the effect of Arkansas on Louisiana is 0.77. This means that if the Louisiana price increases $.01 then in equilibrium the estimated increase in the Arkansas price following the initial shock will be $.0077.

When significant, the multipliers are positive (see Table 2). The only negative multiplier is small and not significantly different from zero. Thus, these

Table 2. Estimated Long-Run Multipliers Showing the Lagged Impact of Changes in Milled Rice Prices

		Affected State			
Time Period	Affecting State	LA	AK	CA	TX
Aug 1975 - July 1980	Louisiana		1.08	0.63*	1.04*
			(0.15)[a]	(0.12)	(0.14)
	Arkansas	0.77*		0.57*	1.09*
		(0.12)		(0.09)	(0.14)
	California	0.38*	0.40*		0.56*
		(0.14)	(0.16)		(0.15)
	Texas	0.31*	0.16	-0.02	
		(0.13)	(0.14)	(0.09)	
Aug 1980 - July 1985	Louisiana		0.06	0.68	0.15
			(0.68)	(0.76)	(0.61)
	Arkansas	2.41*		2.46*	1.75*
		(0.74)		(0.77)	(0.59)
	California	-0.16	-0.06		-0.05
		(0.77)	(0.68)		(0.62)
	Texas	1.07	0.94*	1.37*	
		(0.69)	(0.55)	(0.75)	

*Denotes significance at the 5 percent level using a one-tailed t-test.

[6] Since the dependent variables in (8) are estimated parameters it may be possible to get asymptotically more efficient estimates by jointly estimating equations (1) and (8). Joint estimation is not feasible in this case due to differing numbers of observations and the fact that the π_{ijt}, are a highly nonlinear transformation of the Bij's.

[7] The models were also estimated using 9 lags for both time periods and the results were similar.

results basically support the hypothesis that rice in these four locations are substitutes as prices tend to move in the same direction.

The multipliers in the second period have larger standard errors than in the first. Some of the multipliers in the second period are much larger. None of them are significantly greater than one (using a two-tailed test).[8] The large multipliers are estimated imprecisely as shown by their large standard errors. This illustrates the importance of estimating equation (8) with a method that accounts for the differences in variances.

Arkansas prices have a significant lagged effect on each of the other prices in both periods. Thus, Arkansas tends to be the price leader which is as expected since Arkansas mills the most rice. Louisiana has statistically significant multipliers for 1975-1980. Louisiana mills the least rice, but is considerably less concentrated than the other states. The tendency for Louisiana prices to lead the prices of other states offers some support for the hypothesis that less concentrated regions lead more concentrated ones. Thus, based solely on Table 2, market share would appear to be the most important factor affecting price leadership.

The hypothesized effects of relative submarket size, relative seller concentration, and transportation costs were formally tested. Estimates of the parameters in equation (8) were obtained using the two generalized least squares procedures discussed above and are reported in Table 3. They provide a basis for tests of hypotheses about the effects of selected factors on dynamic price adjustments. Using two-tailed t-tests, the coefficients for quantity and concentration are significant at the 5 percent level for both procedures. The coefficients for transportation cost and the dummy variable representing differences in time are not significant.[9] The negative coefficient for the concentration variable indicates that highly concentrated submarkets lag behind less concentrated submarkets. This provides statistical evidence that the administered price hypothesis may hold as well for submarket pricing patterns: prices in highly concentrated submarkets tend to be sluggish and thus lag behind prices in less concentrated submarkets. Also, the negative coefficient for the quantity variable shows that prices in a larger submarket tend to lead those in a smaller one (as measured by long-run multipliers). This is evidence in favor of the price leadership hypothesis among regions. The two estimation procedures yield parameters with different magnitudes, but still suggest the same conclusions. The absolute values of elasticities for the relative quantities (QR) and relative concentration (CR) are similar. Thus, the magnitude of dynamic price adjustments appears to be similarly responsive to changes in these two variables.

CONCLUSIONS

This paper reported a new two-step estimation procedure for testing hypotheses about the effect of market structure on dynamic price adjustments. The procedure provides a more flexible way of modelling dynamics than did past studies that assumed a partial adjustment process. Applied to weekly milled rice prices, the results provide evidence that submar-

Table 3. Estimated Coefficients for Regressions of Long-run Multipliers

Independent Variable	Estimate	Elasticity	Saxonhouse Approach Estimate	Saxonhouse Approach Elasticity
Intercept	2.05 (1.79)[a]		1.35 (1.80)	
Quantity Ratio	1.18* (2.55)	1.82	1.32* (2.78)	2.04
Transportation Costs	-0.09 (-0.90)	-0.27	-0.11 (-1.03)	-0.33
Ratio of Concentration Ratios	-1.03* (-2.15)	-1.60	-1.11* (-2.41)	1.72
Time Dummy	-1.28 (-1.09)		-0.69 (-0.97)	

[a]The variables in parentheses are t-values.
*Denotes significance at the 5 percent level using a two-tailed t-test.

[8] Ravallion argued that if two submarkets are perfectly integrated then the lagged change and the instantaneous change should sum to one.

[9] Results for the other variables change very little whether the dummy variable for time was included or not. The results also change little whether a dummy variable was included for locations producing the same type of rice. The coefficient on the dummy variable for type of rice was not statistically significant.

ket concentration plays a significant role in the determination of regional milled rice prices: the higher the seller concentration in a given region, the slower prices adjust in that region. This supports the administered price hypothesis. Firms must receive price signals quickly and accurately if they are to make optimal decisions. This research suggests that price adjustments are slower in imperfectly competitive markets. It also shows the impact of submarket concentration rather than just national concentration. This may need to be considered in developing antitrust policy. High concentration within a region is shown to result in slower price adjustments. This suggests that regional concentration measures as well as national concentration measures should be considered when making decisions about allowing mergers and acquisitions.

Results show that prices in larger submarkets tend to lead those in smaller submarkets. In particular, prices in Arkansas, which has the largest market share, consistently led prices in other states. This is interpreted here as evidence in favor of price leadership among regions. These results illustrate that a time series approach combined with alternative measures of market structure can shed new light on dynamic price behavior. In particular, the procedure may be quite promising in any study of price determination when prices are observed at short time intervals while other variables influencing price formation (e.g., costs or quantities) are observed less frequently.

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