Journal of Agricultural and Applied Economics, 32,3(December 2000):543–553

Price Transmission, Threshold Behavior, and Asymmetric Adjustment in the U.S. Pork Sector

Barry K. Goodwin and Daniel C. Harper

ABSTRACT

The US pork sector has experienced many significant structural changes in recent years. Such changes may have influenced price dynamics and transmission of shocks through marketing channels. We investigate linkages among farm, wholesale, and retail markets using weekly price data for the period covering 1987 through 1998. Our analysis uses a threshold cointegration model that permits asymmetric adjustment to positive and negative price shocks. Our results reveal important asymmetries. Our results are consistent with existing literature which has determined that price adjustment patterns are unidirectional and that information tends to flow from farm, to wholesale, to retail markets.

Key Words: *asymmetric price transmission, vertical price transmission, error correction, thresholds, pork markets.*

The U.S. pork sector was shocked in late 1998 by historically low prices. Farm-level prices reached a low of $10.50 per hundredweight in December 1998. Only six months earlier, prices were four times this level. These low farm prices brought about considerable financial stress for producers and led many observers to question the extent to which industry concentration, consolidation, and vertical integration may have been related to the events of 1998. In addition, although farm prices experienced significant declines, price movements of a similar magnitude were not experienced in wholesale and retail markets. This fact has prompted many to question whether price transmission and adjustment in pork markets is asymmetric.

The pork sector has experienced numerous structural changes in recent years. Perhaps most significant have been changes that have affected production and marketing, which have been significantly influenced by the substantial vertical integration that the industry has experienced. A large majority of hogs are now grown and marketed under confidential contracts and private agreements between growers and large packers. In this light, terminal auction markets may have diminished in their importance to the price-discovery process. However, terms of these agreements are

---

The authors are professor and graduate research assistant, respectively, in the Department of Agricultural and Resource Economics at North Carolina State University. This research was supported by the Livestock Pricing Research Institute in the Department of Agricultural Economics at Virginia Tech and by the North Carolina Agricultural Research Service. We are grateful to Wayne Purcell for helpful suggestions which initiated and guided our research. We are also grateful to Ted Schroeder, Lawrence Duewer, Michael Sheats and Thomas Morgan (Morgan Consulting) for graciously supplying much of the data used in this analysis. Valuable comments were provided by two anonymous referees. Direct correspondence to Goodwin at P.O. Box 8109, Raleigh, NC 27695, (919) 515-4547, E-mail: barry_goodwin@ncsu.edu.

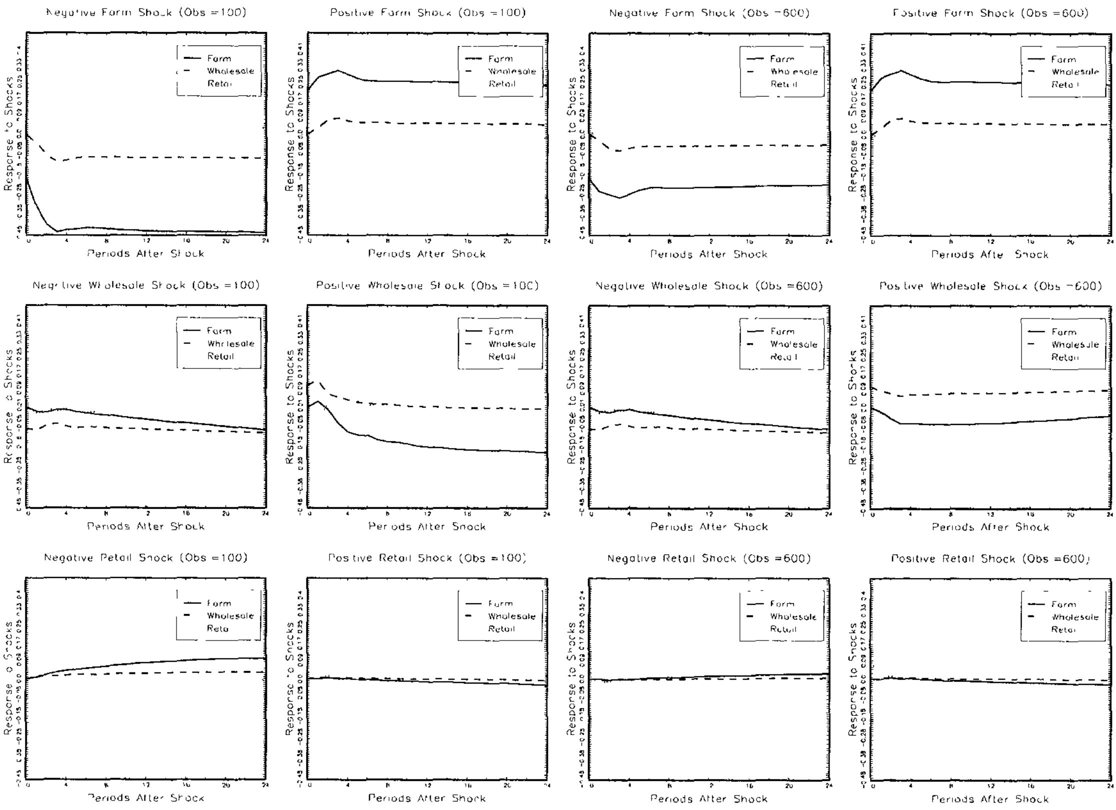

Figure 1. Nonlinear impulse responses for analysis of pork markets: 1987–1998

often based upon spot market conditions, making spot market prices relevant even when prices are set by private agreement. There have also been significant regional shifts in production, with significant growth in the Southeastern U.S. These changes have been accompanied by decreases in the number of producers and, in some cases, by significant increases in the scale of operations.

The vertical transmission of shocks among various levels of the market is an important characteristic describing the overall operation of the market. Market prices have traditionally been the primary mechanisms by which various levels of the market are linked, though the move away from traditional auction markets toward private contracts may have influenced price transmission. The extent of adjustment and the speed with which shocks are transmitted among producer, wholesale, and retail market prices reflects the actions of participants at alternative market levels. The nature, speed, and extent of adjustments to market shocks may also have important implications for price discovery, marketing margins, spreads, and mark-up pricing practices.

An extensive literature has examined market linkages among farm, wholesale, and retail markets. This research has established the existence of significant lags in the adjustment of prices at various levels in the marketing channel [see, for example, Boyd and Brorsen (1988), and Hahn]. These lags are generally attributed to adjustment costs which delay or otherwise inhibit market-price adjustments. Recent research in this area has concentrated on the potential for asymmetric adjustments in prices at various market levels. In particular, the conventional wisdom suggests that responses to price increases may differ from responses to price decreases. Most of these studies utilize some variation of a model originally introduced by Wolffram and later modified by Houck and Ward. These various model specifications typically involve the regression of differenced price data on lagged price differences where allowances are made for differential effects of positive and negative lagged differences. Although a sweeping generalization is difficult to make, most research has revealed the presence of asymmetries in price adjustments at the various market levels though the extent of asymmetry is generally small. In addition, most existing research has found that the direction of causality flows from the farm level to wholesale and retail markets. In particular, farm prices have generally been found to be relatively less responsive to shocks in wholesale and retail markets than is the case for wholesale and retail markets.

A number of institutional and theoretical reasons for asymmetries in price adjustments have been offered.[1] Ward noted that agents in possession of perishable goods may resist the temptation to increase prices for fear of being left with spoiled product. Bailey and Brorsen noted that asymmetries in adjustment costs may underlie asymmetric price adjustments. Imperfectly competitive markets characterized by price leadership roles by major buyers or sellers may also underlie asymmetric price adjustments. Finally, Kinnucan and Forker noted that, where applicable, government intervention through price supports and marketing quotas could lead to asymmetric price adjustments.

Most studies have ignored important time-series properties of the data. In particular, most research has not considered the potential for nonstationarity in individual prices or long-run stationary equilibria (i.e., cointegration) relationships among prices. The typical econometric specification used to evaluate asymmetric price adjustments omits error correction terms and thus is incompatible with long-run cointegration linkages. This limitation of standard models of asymmetry was recently recognized by Cramon-Taubadel in an investigation of asymmetric price adjustment in German producer and wholesale hog markets. Cramon-Taubadel modified the standard Wolffram specification to include an error-correction term and found that wholesale prices reacted more rapidly to positive shocks than to negative shocks originating at the farm level. More recently, Goodwin and Holt evaluated price transmission in beef markets using asymmetric, threshold error correction models. This analysis extends the approach and methods of Goodwin and Holt to consider price transmission in the pork sector.

Although recent research on price transmission has focused on asymmetric adjustments, these models generally require the functional relationships which underlie the price transmission process to be fundamentally *linear.* Recent developments in time series analysis techniques have recognized the potential for nonlinear and threshold-type adjustments in error-correction models. Threshold effects occur when larger shocks (i.e., shocks above some threshold) bring about a different response than do smaller shocks. The resulting dynamic responses may be nonlinear in that they may involve various combinations of adjustments from alternative regimes defined by the thresholds. Threshold models of dynamic economic equilibria are usually motivated by adjustment costs, which may inhibit or otherwise constrain adjustments to small shocks. Put another way, a shock may have to be of a particular size before a significant response is provoked. This analysis evaluates price linkages among producer, wholesale, and retail marketing channels in U.S. pork markets. We use the threshold cointegration methods recently introduced by Balke and Fomby. In particular, a threshold error correction model allowing asymmetric adjustments is estimated and used to evaluate the dynamic time paths of price adjustments to shocks at each level in the U.S. pork sector.

Econometric Methods

Tsay developed an approach to testing for threshold effects and modeling threshold autoregressive processes. Balke and Fomby, noting the correspondence between error-correction models representing cointegration relationships

[1] See Cramon-Taubadel for an extensive discussion of models of asymmetric price transmission.

and autoregressive models of an error-correction term, extended the threshold autoregressive models to a cointegration framework. Balke and Fomby also showed that standard methods for evaluating unit roots and cointegration work reasonably well when threshold cointegration is present.[2]

Consider a standard cointegration relationship representing an economic equilibrium

$$(1)\quad y_{1t} - \beta_1 y_{2t} - \beta_2 y_{3t} - \cdots \beta_k y_{kt} = v_t$$

where

$$v_t = \rho v_{t-1} + e_t.$$

Cointegration of the y_{it} variables depends upon the nature of the autoregressive process for v_t. As ρ approaches 1, deviations from the equilibrium become nonstationary and thus the variables are not cointegrated. Balke and Fomby extend this simple framework to the case where v_t follows a threshold autoregression:

$$(2)\quad \rho = \begin{cases} \rho^{(1)} & \text{if } |v_{t-1}| \leq c \\ \rho^{(2)} & \text{if } |v_{t-1}| > c, \end{cases}$$

where c represents the threshold which delineates alternative regimes.[3] A common case is that of $\rho^{(1)} = 1$, which implies that the relationship for small deviations from equilibrium is characterized by a random walk (i.e., a lack of cointegration). Parity relationships among commodity prices and interest rates have been examined in such a context.[4]

Balke and Fomby note that this simple framework is easily extended to permit multiple thresholds, implying multiple parametric regimes and thus allowing asymmetric adjustment. In the case of k thresholds, $k + 1$ different regimes are implied, each having a unique set of parameters and implying its own dynamics for the system. Multiple thresholds allow one to model asymmetries in relationships among the variables as different regimes may correspond to positive versus negative shocks. Our analysis considers a case of two thresholds (c_1 and c_2), which implies three regimes. In this case, an equivalent vector error correction representation of the threshold model is given by:

$$(3)\quad \Delta y_t = \begin{cases} \sum_{i=1}^{p} \gamma_i^{(1)} \Delta y_{t-i} + \theta^{(1)} v_{t-1} + \epsilon_t^{(1)} & \text{if } v_{t-1} < c_1 \\ \sum_{i=1}^{p} \gamma_i^{(2)} \Delta y_{t-i} + \theta^{(2)} v_{t-1} + \epsilon_t^{(2)} & \text{if } c_1 \leq v_{t-1} \leq c_2 \\ \sum_{i=1}^{p} \gamma_i^{(3)} \Delta y_{t-i} + \theta^{(3)} v_{t-1} + \epsilon_t^{(3)} & \text{if } v_{t-1} > c_2 \end{cases}$$

where ϵ_t is a mean zero residual.

Testing for threshold effects presents a number of challenges. Tsay developed a general nonparametric test for the nonlinearity implied by thresholds in an autoregressive series. Consider a standard autoregressive model of the form:

$$(4)\quad v_t = \alpha + \gamma v_{t-1} + \epsilon_t.$$

In constructing Tsay's test, we denote each combination of v_t and v_{t-1} as a "case" of data. The individual cases of data are ordered according to the variable relevant to the threshold behavior, v_{t-1} in this case. Recursive residuals are obtained by estimating the autoregressive model for an initial sample and then for sequentially updated samples obtained by adding a single observation. A test of nonlinearity is then given by the regression F-statistic obtained by regressing the recursive residuals on the explanatory variables (v_{t-1}). Obstfeld and Taylor note that, as a practical matter, the test should be run with both increasing and decreasing ordering in the ar-

[2] Balke and Fomby also indicate, however, that standard tests may lack power in the presence of asymmetric adjustment.

[3] More generally, thresholds pertain to some delay parameter d in adjustment to v_t, such that $|v_{t-d}| \leq c$ defines the threshold. Although testing for d is discussed below, most applications assume a delay of $d = 1$.

[4] See Obstfeld and Taylor and Goodwin and Grennes for examples of the former and Siklos and Granger for an example of the latter.

ranged autoregression.[5] Tsay's test is also useful in determining the "delay" parameter which defines the threshold autoregression in equation (2). The test is typically run for alternative delays and the delay giving the largest F-statistic is chosen as optimal.

Once the presence of threshold effects is confirmed, some parametric estimation strategy must be considered to estimate the threshold. We use a two-dimensional grid search to estimate the thresholds c_1 and c_2 which define the three regimes. Two alternative grid search techniques have been proposed. Obstfeld and Taylor use a grid search to find the threshold which maximizes a likelihood function. Alternatively, we follow Balke and Fomby and use a grid search which minimizes a sum of squared error criterion.

Our specific estimation strategy can be summarized as follows. First, standard Dickey-Fuller unit root tests and Johansen cointegration tests are used to evaluate the time-series properties of the data. We then follow the general two-step approach of Engle and Granger and use ordinary least squares estimates of a cointegrating relationship among the variables.[6] Lagged residuals from this regression are then used to define the error-correction terms. A two-dimensional grid search is then conducted to define two thresholds. In particular, we search for the first threshold between 5 percent and 95 percent of the largest (in absolute value) negative residual. In like fashion, we search for the second threshold between 5 percent and 95 percent of the largest positive residual. The error-correction model is then estimated conditional on the threshold parameters.

Some method of testing the statistical significance of the differences in parameters across alternative regimes is desirable. In the case of a single threshold, this amounts to a conventional Chow test of parameter differences. As is well known, this testing problem is complicated by the fact that the threshold parameter is not identified under the null hypothesis of no threshold effects and thus conventional test statistics have nonstandard distributions. Hansen has developed an approach to testing the statistical significance of threshold effects. After optimal thresholds have been identified, a conventional Chow-type test of the significance of threshold effects (i.e., the significance of the differences in parameters over alternative regimes) is conducted. Because the test statistic has a nonstandard distribution, simulation methods must be used to approximate the asymptotic distribution and identify appropriate critical values. Hansen recommends running a number of simulations whereby the dependent variables are replaced by standard normal random draws. For each simulated sample, the grid search is used to select optimal thresholds and the standard Chow-type test is used to test the significance of the threshold effects. From this simulated sample of test statistics, the asymptotic p-value is approximated by taking the percentage of test statistics for which the test taken from the estimation sample exceeds the observed test statistics.

Empirical Application

Our empirical analysis uses three series of weekly (logged) pork prices observed from January 1987 through the first week of January 1999, giving a total of 626 observations.[7] Producer prices are quoted for U.S. 1–2 230–250 lb. hogs at the Iowa-Southern Minnesota direct market and were taken from unpublished USDA-AMS data. Wholesale prices were collected from unpublished Economic Research Service data. A retail price index was constructed using a weighted average of

[5] The test is nonparametric in that it depends neither on the number of thresholds nor their values. The alternative ordering allows more power in discerning thresholds for which data are concentrated in a regime at either end of the arranged series. We report only the more significant of the two ordered tests.

[6] In cases of $p > 2$ variables, finding multiple cointegrating relationships suggests that the OLS estimates are not unique. Properties of the OLS estimates in such a case are discussed by Hamilton. As always, the results may also be sensitive to the normalization rule.

[7] Our use of logarithmic prices assumes a proportional relationship among prices at the alternative market levels. Similar results were obtained when a linear relationship among prices was assumed.

retail prices for individual cuts taken from Bridge.[8] The Bridge retail price data are collected from a survey of advertised specials appearing in newspapers. It is worth noting that to the extent these specials are determined before the advertisements appear the quoted prices may not reflect the most up-to-date information that retailers have about their costs.[9] We nevertheless believe them to provide a reasonably accurate depiction of actual price movements.

Standard unit-root tests confirmed a single unit root in each price series. Johansen cointegration tests (Table 1) indicated the existence of a single cointegrating relationship among the three prices. Lag orders for the cointegration tests and threshold error-correction models were chosen using Akaike and Schwartz-Bayesian criteria. The alternative criteria indicated lag orders ranging from 3 to 5. An evaluation of autocorrelation patterns for the residuals led us to adopt a specification with four lags in both the cointegration and error correction models. The equilibrium relationship was normalized on the retail price and ordinary least squares was used to obtain estimates of the cointegrating relationship. These estimates are presented in Table 1.[10]

Tsay's test was conducted using the error-correction terms implied by the OLS estimates. The test (Table 1) strongly rejected linearity and thus implied the presence of one or more thresholds. The largest rejections occurred for delays of a single week, suggesting a delay parameter of 1. The two-dimensional grid search identified thresholds at −0.0636 and 0.0123. A standard likelihood ratio test of the significance of the differences in parameters across regimes was strongly rejected using conventional critical values. As noted above, however, the test statistic is likely to be nonstandard since a search for the thresholds preceded the testing. When Hansen's simulation approach was used, the test statistic exceeded each replicated value, confirming the statistical significance of the test. The thresholds correspond to three regimes of 83, 297, and 246 observations, respectively.[11]

Table 1. Cointegration and Threshold Testing Results: Analysis of Pork Markets (1987–1998)

Test	Test Statistic	Critical Value[a]
Maximum Eigenvalue Test Statistic		
r = 0	35.09	21.28
r = 1	6.13	10.29
r = 2	1.74	7.5
Trace Test Statistic		
r = 0	42.95	31.88
r = 1	7.87	17.79
r = 2	1.74	7.5
Tsay's Nonlinearity Test	21.42	0.001[b]
Hansen's Threshold Test	74.294	0.001[c]

OLS Estimates of Cointegrating Relationship

$$P_t^R = \underset{(0.1195)^d}{2.153} + \underset{(0.0210)}{-0.3203*P_t^F} + \underset{(0.0434)}{0.9966*P_t^W}$$

$R^2 = 0.4884$

Threshold/Regime Estimates

Regime I	$(-\infty < \nu_{t-1} \leq -0.0636)$	$n = 83$
Regime II	$(-0.0636 < \nu_{t-1} \leq 0.0123)$	$n = 297$
Regime III	$(0.0123 < \nu_{t-1} < \infty)$	$n = 246$

[a] Critical values are at the $\alpha = 0.10$ level and are taken from Hansen and Juselius (1995).
[b] Approximate asymptotic p-value for test statistic.
[c] Empirical p-value based upon bootstrap simulation.
[d] Numbers in parentheses are standard errors.

[8] Weights were constructed by the authors using approximate meat cut yields taken from the National Pork Producers Council's 1998–1999 *Pork Facts.* The component prices and weights are available from the authors on request. It should be acknowledged that this approach assumes a fixed-proportions relationship among individual meat cuts.

[9] We thank an anonymous reviewer for pointing this out.

[10] In that deterministic time trends did not appear to be present in the series, we restricted the intercept term to apply to the cointegration relationship only.

[11] Because of the long computing time required for the simulation, a modified grid search was used. The modified search estimated each threshold conditional

Table 2. Threshold Error-Correction Model Parameter Estimates and Summary Statistics

Variable	Regime I Farm Parameter Estimate	Regime I Wholesale Parameter Estimate	Regime I Retail Parameter Estimate
ΔF_{t-1}	−0.0883	0.1118	0.2795
	(0.1159)	(0.1159)	(0.1159)*
ΔW_{t-1}	0.2655	0.0601	−0.4863
	(0.2329)	(0.2329)	(0.2329)*
ΔR_{t-1}	−0.0831	0.0477	−0.5318
	(0.1059)	(0.1059)	(0.1059)*
ΔF_{t-2}	−0.077	0.2221	−0.1125
	(0.1618)	(0.1618)	(0.1618)
ΔW_{t-2}	−0.4981	−0.677	0.0237
	(0.2475)*	(0.2475)*	(0.2475)
ΔR_{t-2}	−0.0804	0.0956	−0.3463
	(0.1073)	(0.1073)	(0.1073)*
ΔF_{t-3}	0.6685	0.1166	−0.0988
	(0.1631)*	(0.1631)	(0.1631)
ΔW_{t-3}	−0.5015	−0.0982	−0.0024
	(0.2938)	(0.2938)	(0.2938)
ΔR_{t-3}	−0.0102	0.0828	−0.1058
	(0.0964)	(0.0964)	(0.0964)
ΔF_{t-4}	−0.2528	0.0076	−0.0829
	(0.1572)	(0.1572)	(0.1572)
ΔW_{t-4}	−0.1245	−0.2402	0.5723
	(0.2540)	(0.254)	(0.2540)*
ΔR_{t-4}	0.0202	0.0554	0.1266
	(0.0936)	(0.0936)	(0.0936)
v_{t-1}	0.0197	−0.0195	−0.1988
	(0.0665)	(0.0665)	(0.0665)*

Variable	Regime II Farm Parameter Estimate	Regime II Wholesale Parameter Estimate	Regime II Retail Parameter Estimate
ΔF_{t-1}	0.6475	0.2297	−0.106
	(0.0696)*	(0.0696)*	(0.0696)
ΔW_{t-1}	−0.1317	−0.1538	−0.1933
	(0.1246)	(0.1246)	(0.1246)
ΔR_{t-1}	0.0632	−0.0355	−0.4803
	(0.0603)	(0.0603)	(0.0603)*
ΔF_{t-2}	0.0316	0.1568	0.0465
	(0.0709)	(0.0709)*	(0.0709)
ΔW_{t-2}	−0.2244	−0.0666	0.0294
	(0.1237)	(0.1237)	(0.1237)
ΔR_{t-2}	−0.0119	−0.0385	−0.3174
	(0.0597)	(0.0597)	(0.0597)*
ΔF_{t-3}	0.262	0.0106	0.3328
	(0.0815)*	(0.0815)	(0.0815)*

Table 2. (Continued)

Variable	Reigme II Farm Parameter Estimate	Reigme II Wholesale Parameter Estimate	Reigme II Retail Parameter Estimate
ΔW_{t-3}	−0.4627	−0.1363	−0.3783
	(0.1133)*	(0.1133)	(0.1133)*
ΔR_{t-3}	−0.062	−0.0435	−0.1736
	(0.0575)	(0.0575)	(0.0575)*
ΔF_{t-4}	0.1468	−0.0423	−0.1299
	(0.0789)	(0.0789)	(0.0789)
ΔW_{t-4}	−0.1807	0.0485	0.113
	(0.1199)	(0.1199)	(0.1199)
ΔR_{t-4}	0.046	0.0068	−0.1056
	(0.0466)	(0.0466)	(0.0466)*
v_{t-1}	0.0554	0.039	−0.2423
	(0.0670)	(0.067)	(0.067)*

Variable	Reigme III Farm Parameter Estimate	Reigme III Wholesale Parameter Estimate	Reigme III Retail Parameter Estimate
ΔF_{t-1}	0.2944	0.1386	−0.0333
	(0.0769)*	(0.0769)	(0.0769)
ΔW_{t-1}	0.1317	0.0339	0.1774
	(0.1355)	(0.1355)	(0.1355)
ΔR_{t-1}	0.0546	0.0431	−0.6576
	(0.0552)	(0.0552)	(0.0552)*
ΔF_{t-2}	−0.0046	0.1328	−0.1156
	(0.0818)	(0.0818)	(0.0818)
ΔW_{t-2}	−0.0489	−0.2318	−0.0773
	(0.1328)	(0.1328)	(0.1328)
ΔR_{t-2}	0.123	0.0653	−0.5741
	(0.0648)	(0.0648)	(0.0648)*
ΔF_{t-3}	0.0502	0.0468	0.079
	(0.0825)	(0.0825)	(0.0825)
ΔW_{t-3}	−0.1846	−0.1435	0.1069
	(0.1327)	(0.1327)	(0.1327)
ΔR_{t-3}	0.0087	0.0044	−0.3415
	(0.0628)	(0.0628)	(0.0628)*
ΔF_{t-4}	−0.0931	−0.0482	−0.0004
	(0.0881)	(0.0881)	(0.0881)
ΔW_{t-4}	0.0054	0.1236	−0.1216
	(0.1369)	(0.1369)	(0.1369)
ΔR_{t-4}	0.0973	0.0841	−0.0522
	(0.0534)	(0.0534)	(0.0534)
v_{t-1}	−0.0579	−0.0041	−0.0991
	(0.0365)	(0.0365)	(0.0365)*

[a] Numbers in parentheses are standard errors (conditional on thresholds and homogeneous variances across regimes). Asterisks indicate parameter estimates more than twice the standard error.

Table 3. Summary of Regime Switching: Percentages of Observations Falling into Each Regime (1987–1998)

Year	Regime I	Regime II	Regime III
1987	5.88	62.75	31.37
1988	16.98	62.26	20.75
1989	15.38	67.31	17.31
1990	0.00	59.62	40.38
1991	1.92	28.85	69.23
1992	7.69	51.92	40.38
1993	7.55	62.26	30.19
1994	32.69	55.77	11.54
1995	50.00	42.31	7.69
1996	15.38	51.92	32.69
1997	0.00	11.54	88.46
1998	5.66	13.21	81.13
1987–1998	13.26	47.44	39.30

Parameter estimates and standard errors for the error-correction model are presented in Table 2. The standard errors estimates are conditional on the thresholds and the assumption of homogeneous error variances across regimes. In light of the aforementioned shortcomings associated with nonstandard test statistics and the inferential limitations inherent in our estimation approach, the standard error estimates should not be used for direct tests but rather only as a rough guide of the likely significance of individual parameter estimates. The estimates indicate significant dynamic relationships among the price series. In general, dynamic interrelationships among the prices reflect relatively more interaction between retail prices and lagged price differences than is the case for farm and wholesale prices and lags—a finding consistent with causality in the direction of farm and wholesale to retail levels. Error-correction terms are especially significant in the retail equation, again a result consistent with retail prices being responsive to shocks at the wholesale and retail levels.

An evaluation of the timing and frequency of shifts between the three alternative regimes is also useful. Table 3 presents a summary of the frequency with which observations fall into each regime for each year in our sample. It is important to recognize that the price adjustment process at any point is unique in that it depends upon the values of the error-correction term and lagged price differences at each observation. This is in contrast to standard vector autoregressive and error-correction models, where responses to shocks are independent of the timing of the shock. At each point in time (i.e., each observation), depending on the value of the error-correction term, one of the three regimes characterizes the relationship among prices. In light of the definition of the error-correction term (i.e., residuals from the cointegrating regression having retail prices as the left-hand-side variable), Regime I corresponds to large (in absolute value) negative errors (retail prices below equilibrium) that exceed the lower threshold while Regime III corresponds to large positive errors (retail prices above equilibrium). Regime II corresponds to errors that are between the thresholds that define Regimes I and III. Regime II dominates, accounting for 47.4 percent of the observations. Observations only rarely fall into Regime I, which accounts for only 13.3 percent of the observations. Regime III, representing 'higher than usual' retail prices, is also frequently observed (39.3 percent of the time). The third regime especially dominates in later periods. In the final year (1998), 81.1 percent of the observations fall into the third regime. The first regime accounts for a significant proportion of the observations in 1994 and 1995, a period of strengthening in farm prices. A consideration of jumps among regimes (not illustrated here) suggested that jumps between Regimes II and III appeared to be much more influential toward the end of the sample, the period dominated by very low farm prices. However, switching between regimes appeared to occur throughout the period of study.

Interpretation of the dynamic interrelationships among prices at alternative market levels is best pursued through a consideration of impulse response functions. Again, in contrast to the linear model case, the response to a shock is dependent upon the history of the series. In addition, the possibly asymmetric nature of re-

upon the other and iterated until identical estimates were obtained. Five hundred replications were used.

sponses implies that the size and sign of the shock will influence the nature of the response. In this light there are many possible impulse response functions. We chose two observations representative of the early (observation 100) and late (observation 600) periods to evaluate responses to shocks. We adopt the nonlinear impulse response function approach of Potter, which defines responses (denoted I_{t+k}) on the basis of observed data ($z_t, z_{t-1}, \ldots$) and a shock (v) as:

$$(5) \quad I_{t+k}(v, Z_t, Z_{t-1}, \ldots) = E[Z_{t+k} | Z_t = z_t + v, Z_{t-1} = z_{t-1}, \ldots] - E[Z_{t+k} | Z_t = z_t, Z_{t-1} = z_{t-1}, \ldots].$$

It should also be noted that in light of the nonstationary nature of the price data and the error-correction properties of the system of equations shocks may elicit either transitory or permanent responses. In particular, *nonstationarity* implies that shocks may permanently alter the time path of variables.

Figure 1 illustrates responses to one standard deviation positive and negative shocks. The first two columns illustrate responses to negative and positive shocks, respectively, at observation 100 (December 2, 1988) while the latter two columns provide the corresponding responses at observation 600 (July 3, 1998). Several implications for price interrelationships emerge from the responses. First, some minor asymmetries are apparent, particularly in the earlier period. Negative farm price shocks appear to elicit a greater movement in farm prices in subsequent periods than do positive shocks. The asymmetric responses are notably less apparent in the later period. An additional, unexpected result is the response of farm prices to wholesale shocks. Especially in the earlier period, positive wholesale price shocks appear to elicit a negative response in farm prices. In the later period this effect is less evident and does not appear to be economically significant.

Shocks elicit permanent adjustments which are complete after four to six weeks in almost all cases. The effect of price shocks does appear to be somewhat damped as one moves up the marketing chain. Farm shocks elicit large adjustments in farm price, relatively little adjustment in wholesale prices, and have even less effect on retail prices. Wholesale shocks, with the single exception noted above, affect primarily wholesale and retail prices, and the response to retail shocks appears to be largely confined to retail markets. With the exception of farm price responses to farm and wholesale shocks in the early period, neither the extent nor the timing of adjustments appears to have changed significantly over the period included in our data.

The implications of the impulse responses are generally in agreement with expectations and with previous research. Our finding that price transmission appears to occur mainly in one direction—from farm to wholesale to retail markets—parallels the findings of Boyd and Brorsen (1985) and Schroeder. Our finding that asymmetries are present is in line with Hahn. In contrast to Hahn, however, we find that observed asymmetries are minor, are not present at all market levels, and are not present in more recent data. This finding could be due to the effect of aforementioned structural changes which have occurred in pork markets in recent years. Responses are generally as one would expect, with positive shocks eliciting positive responses and negative shocks eliciting negative responses.

In all, the results are consistent with conclusions of most of the extensive literature that has examined price transmission among farm, wholesale, and retail meat markets. The results suggest that the flow of price information is unidirectional, going from the farm level, to wholesale markets, and finally to retail markets. Information does not appear to flow in the opposite direction in that retail market shocks do not significantly influence farm and wholesale prices. Classic notions of retail-to-farm price transmission often assume that price-taking farmers receive prices that are largely determined by large processors and retailers that have significant market power which enables them to set prices. Our results, and indeed the results of the large literature on this topic, are not entirely at odds with this view but rather suggest that retail price shocks are not passed back to farm-level markets. The extent to which this indicates the absence or presence of market power and discriminatory pricing practices is unclear and remains a topic for future research.

Conclusion

We have examined price interrelationships and transmission among farm, wholesale, and retail pork markets. We give special attention to the time-series properties of the price data. In particular we estimate a threshold error correction model which recognizes the nonstationary nature of the price data and allows for asymmetric price responses.

We find that threshold effects are significant. While farm, wholesale, and retail pork prices are found to be cointegrated, our model suggests the degree of adjustment in prices at one market level to shocks at another level varies significantly with the size of the shock. This result is consistent with the presence of adjustment lags observed in previous research.

Our results regarding price transmission largely confirm the findings of other research. In particular we confirm previous findings that the transmission of shocks appears to be largely unidirectional with information flowing up the marketing channel from farm to wholesale to retail markets but not in the opposite direction. Farm markets do adjust to wholesale market shocks. The effects of retail market shocks, however, are largely confined to retail markets. Minor asymmetries are present in the response of farm prices to farm and wholesale shocks in the earlier period, but are no longer as apparent in the later period. To the extent that recent structural changes in the pork marketing channel are responsible for this reduction in asymmetric adjustment, this suggests that the changes may have had a beneficial effect on the efficiency of the price transmission mechanism. However, any conclusions regarding improved efficiency of price transmission are tenuous in light of the limited transmission of price information from retail to farm markets implied by our results.

References

Bailey, D. V. and B. W. Brorsen. "Price Asymmetry in Spatial Fed Cattle Markets." *Western Journal of Agricultural Economics* 14(1989): 246–252.

Balke, N. S. and T. B. Fomby. "Threshold Cointegration." *International Economic Review* 38(1997):627–645.

Boyd, M. S. and B. W. Brorsen. "Dynamic Relationship of Weekly Prices in the United States Beef and Pork Marketing Channels." *Canadian Journal of Agricultural Economics* 33(1985): 331–442.

Boyd, M. S. and B. W. Brorsen. "Price Asymmetry in the U.S. Pork Marketing Channel." *North Central Journal of Agricultural Economics* 10(1988):103–110.

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

Goodwin, B. K. and T. J. Grennes. "Tsarist Russia and the World Wheat Market." *Explorations in Economic History* 35(1998):405–430.

Goodwin, B. K. and M. T. Holt. "Asymmetric Adjustment and Price Transmission in the U.S. Beef Sector." *American Journal of Agricultural Economics* 81(1999):630–637.

Hahn, W. F. "Price Transmission Asymmetry in Pork and Beef Markets." *Journal of Agricultural Economics Research* 42(1990):21–30.

Hamilton, J. D. *Time Series Analysis.* Princeton: Princeton University Press, 1994.

Hansen, B. E. "Inference in TAR Models." *Studies in Nonlinear Dynamics and Econometrics* 2(April 1997):1–14.

Hansen, H. and K. Juselius. *CATS in RATS—User's Manual.* Evanston, Illinois: Estima, 1995.

Houck, J. P. "An Approach to Specifying and Estimating Non-Reversible Functions." *American Journal of Agricultural Economics* 59(1977): 570–572.

Kinnucan, H. W. and O. D. Forker. "Asymmetry in Farm-Retail Price Transmission for Major Dairy Products." *American Journal of Agricultural Economics* 69(1987):285–292.

National Pork Producers Council. *1998–99 Pork Facts.* Des Moines, Iowa, 1999.

Obstfeld, M. and A. M. Taylor. "Nonlinear Aspects of Goods-Market Arbitrage and Adjustment; Heckscher's Commodity Points Revisited." *Journal of the Japanese and International Economies* 11(1997):441–479.

Potter, S. M. "A Nonlinear Approach to US GNP."

Journal of Applied Econometrics 10(1995): 109–125.

Schroeder, T. C. "Price Linkages Between Wholesale and Retail Pork Cuts." *Agribusiness* 4(June–July 1988):359–369.

Siklos, P. L. and C. W. J. Granger. "Regime Sensitive Cointegration: With an Application to Interest Rate Parity." *Macroeconomic Dynamics* 1(1997):(online).

Tsay, R. S. "Testing and Modeling Threshold Autoregressive Processes." *Journal of the American Statistical Association* 84(1989):231–240.

von Cramon-Taubadel, S. "Estimating Asymmetric Price Transmission with the Error Correction Representation: An Application to the German Pork Market." *European Review of Agricultural Economics* 25(1998):1–18.

Ward, R. W. "Asymmetry in Retail, Wholesale, and Shipping Point Prices for Fresh Fruits and Vegetables." *American Journal of Agricultural Economics* 62(1982):205–212.

Wolffram, R. "Positivisitic Measures of Aggregate Supply Elasticities—Some New Approaches—Some Critical Notes." *American Journal of Agricultural Economics* 53(1971):356–359.