Spatial Market Integration in the Presence of Threshold Effects

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A large body of research has evaluated price linkages in spatially separate markets. Much recent research has applied models appropriate for nonstationary data. Such analyses have been criticized for their ignorance of transactions costs, which may inhibit price adjustments and thus affect tests of integration. This analysis utilizes threshold autoregression and cointegration models to account for a neutral band representing transactions costs. We evaluate daily price linkages among four corn and four soybean markets in North Carolina. Nonlinear impulse response functions are used to investigate dynamic patterns of adjustments to shocks. Our results confirm the presence of thresholds and indicate strong support for market integration, though adjustments following shocks may take many days to be complete. In every case, the threshold models suggest much faster adjustments in response to deviations from equilibrium than is the case when threshold behavior is ignored.

Key words: market integration, thresholds, transactions costs.

A large amount of empirical research has evaluated the extent to which spatially separate markets are integrated. Though the term is used loosely in the literature, tests of "market integration" usually consider the extent to which shocks are transmitted among spatially separate markets. The integration of markets can have important implications for price discovery and the operation of the market since persistent deviations from integration may imply riskless profit opportunities for spatial traders.

Early studies of integration typically adopted price correlation or regression-based tests. More recent studies have built upon the realization that the price data typically used to evaluate spatial integration are often nonstationary, leading to inferential problems in empirical tests. To address such inferential problems, a variety of econometric procedures appropriate to nonstationary and cointegrated data have been adopted to evaluate spatial integration (see, for example, Ardeni, Goodwin and Schroeder, Baffes).

Regression and cointegration-based tests have been criticized recently for their ignorance of transactions costs (McNew and Fackler, Fackler and Goodwin, Barrett).[1] The primary mechanism ensuring integration is spatial trade and arbitrage.[2] The presence of transactions costs, which typically are unobservable to the empirical researcher, may lead to a "neutral band" within which prices are not linked to one another. Price equalizing arbitrage activities are triggered only when localized shocks result in price differences which exceed the neutral band.

Recognition of the important but often neglected role of transactions costs has led to the application of new empirical approaches which explicitly recognize the influences of transactions costs on spatial market linkages. Spiller and Wood, Sexton, Kling, and Carman, and Baulch applied endogenous switching models which account for the multiple regimes that may result from transactions costs. In another line of research, Obstfeld and Taylor and Goodwin and Grennes used

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[1] The idea is not new. Heckscher noted that transactions costs could create "commodity points"—a neutral band which causes deviations from market integration.

[2] This mechanism may involve explicit arbitrage where traders transport grain between terminal markets in response to price differences. Alternatively, and probably more likely for the markets considered here, integration may result from the actions of a large number of widely dispersed producers who evaluate price conditions among several terminal markets and sell in the market with the highest net price. Such collective actions lead to equalization of net marginal returns and thus prices across space. In such a case, market prices will be expected to differ by no more than the difference in the costs of selling in one market versus another.

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threshold autoregression models to examine market integration. Such models recognize thresholds, caused by transactions costs, that deviations must exceed before provoking equilibrating price adjustments which lead to market integration. 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 involve various combinations of adjustments from alternative regimes defined by the thresholds.

Our analysis is motivated, in part, by unobservable transactions costs and the important influence that their presence may exert on equilibrium spatial price relationships. As has been widely recognized in recent years, regression-based tests of market integration may result in misleading inferences when transactions costs are ignored. Our analysis addresses this shortcoming by estimating "neutral bands" that represent the effect of transactions costs on price relationships. It is important to recognize that our approach, like all other studies that attempt to account for unobservable transactions costs, relies upon assumptions that are potentially restrictive. In particular, we assume that the threshold parameters representing relative transactions costs are constant (in proportional terms), implying a fixed neutral band over the period of our study. As Li and Barrett have recently pointed out, transactions costs and the neutral bands which result may not be constant in the long run and may, in fact, even be nonstationary.

The objective of this analysis is to evaluate price linkages among several local corn and soybean markets in North Carolina. Our approach explicitly accounts for the "neutral band" resulting from transactions costs which may inhibit market integration. We utilize the threshold autoregression and related threshold cointegration methods recently introduced by Balke and Fomby. In particular, we first evaluate deviations from a price parity condition for threshold effects in an autoregressive type of model. Tests for threshold effects confirm their significance and estimates allowing for symmetric thresholds reveal that large deviations from parity are eliminated much more rapidly than are small deviations. Likewise, the threshold models imply much faster adjustment to deviations from equilibrium conditions than is the case when thresholds are ignored. A richer multiple-threshold error correction model allowing asymmetric adjustments is then estimated and used to evaluate the dynamic time paths of price adjustments in response to spatially isolated shocks in each of the markets. We utilize a large sample of daily prices quoted at the four principal North Carolina corn and soybean markets over a seven year period.

The next section discusses econometric methods appropriate to the analysis of threshold autoregression models. These methods are then applied in an evaluation of spatial integration using daily corn and soybean prices. The final section summarizes the analysis and offers some concluding remarks.

Econometric Methods

An extensive literature has applied time series procedures appropriate for nonstationary data to the analysis of market integration. In particular, cointegration analysis techniques have been used to examine whether long-run equilibria exist among two or more prices. As already noted, the ignorance of transactions costs, which may inhibit price transmission across spatially separate markets, has been a major limitation of these tests. To address this limitation, we adopt threshold autoregression and cointegration models.

Threshold models are analogous to more conventional regime switching models. In the case of threshold models, the regime switch is triggered when a forcing variable (or variables) crosses a "threshold." In the case of models that include autoregressive structures, a sequence of observations at any point in time may involve various combinations of alternative regimes, making the underlying structure of the model intrinsically nonlinear.

Tong originally introduced nonlinear threshold time series models. Tsay developed techniques for testing autoregressive models for threshold effects and modeling threshold autoregressive processes. Balke and Fomby, noting the correspondence between error correction models representing cointegration relationships and autoregressive models of an error correction term, extended the threshold autoregressive models to a cointegration framework.[3] Balke and Fomby also showed that standard methods for evaluating

[3] To see this correspondence, consider a simple equilibrium relationship of the form $y_t = \alpha + \beta x_t + e_t$. The extent of

unit roots and cointegration work reasonably well when threshold cointegration is present.[4]

Consider a standard cointegration relationship representing an economic equilibrium

$$(1) \quad y_{1t} - \beta_1 y_{2t} - \beta_2 y_{3t} - \cdots \beta_k y_{kt} = \nu_t \quad \text{where } \nu_t = \rho \nu_{t-1} + e_t.$$

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

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

where c represents the threshold which delineates alternative regimes.[5] 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., an absence of cointegration). Parity relationships among commodity prices and interest rates have been examined in such a context.[6]

An equivalent vector error correction representation of the threshold model can be written as

$$(3) \quad \Delta y_t = \begin{cases} \sum_{i=1}^{p} \gamma_i^{(1)} \Delta y_{t-i} + \theta^{(1)} \nu_{t-1} + \epsilon_t^{(1)} & \text{if } |\nu_{t-1}| \le c \\ \sum_{i=1}^{q} \gamma_i^{(2)} \Delta y_{t-i} + \theta^{(2)} \nu_{t-1} + \epsilon_t^{(2)} & \text{if } |\nu_{t-1}| > c \end{cases}$$

where ϵ_t is a mean zero residual. Balke and Fomby discuss a number of extensions to this framework, including models with multiple thresholds which imply multiple parametric regimes and thus allow asymmetric adjustment.[7] In our analysis, we follow Martens, Kofman, and Vorst and utilize two thresholds (c_1 and c_2) which allows three regimes and thus permits asymmetric adjustment.[8]

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

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

In constructing Tsay's test, we denote each combination of ν_t and ν_{t-1} as a "case" of data. The individual cases of data are ordered according to the variable relevant to the threshold behavior, ν_{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 (ν_{t-1}). Obstfeld and Taylor note that, as a practical matter, the test should be run with both increasing and decreasing ordering in the arranged autoregression.[9] Tsay's test may also be useful in determining the 'delay' parameter d which defines the number of lags appropriate to the error correction term in the threshold autoregression given by 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. Following the standard approach, we utilize a grid search to estimate the thresholds which define the alternative regimes. Two alternative grid search techniques have been proposed. Obstfeld and

cointegration between y_t and x_t depends upon the autoregressive root ρ in $e_t = \rho e_{t-1} + \nu_t$. As ρ approaches 1, deviations from the equilibrium become nonstationary and thus the series are not cointegrated. We work with the analogous error correction version of this relationship, which is expressed as $\Delta e_t = \lambda e_{t-1}$, where $\rho = 1$ implies $\lambda = 0$.

[4] Balke and Fomby and Enders and Granger have also shown, however, that standard tests may lack power in the presence of asymmetric adjustment.

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

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

[7] In the case of k thresholds, $k+1$ different regimes are implied, each of which may imply its own set of dynamics for the system. Extensions to this framework include "band-TAR" models in which adjustment is toward the edge of the threshold and "returning-drift" TAR models which model the processes as random walks with drift toward the thresholds.

[8] The number of thresholds considered is typically constrained by the number of available observations, 1773 in our case.

[9] Additional lags of the error correction term may be added to produce white noise residuals. The test is nonparametric in that it depends neither on the number of thresholds nor their values. The alternative ordering of the data in the arranged regressions allows more power in discerning thresholds for which data are concentrated in a particular regime at either end of the arranged series. We report only the more significant of the two ordered tests.

Taylor use a grid search to find the threshold maximizing a likelihood function. Alternatively, we follow Balke and Fomby and use a grid search minimizing 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 consider ordinary least squares estimates of a cointegrating relationship among the variables.[10] Lagged residuals from this regression are then used to define the error correction terms. Symmetric threshold autoregressions of price differentials and the error correction terms are then considered. The nature of adjustments within and outside the bands identified in the threshold autoregression models are considered. We anticipate that the recognition of threshold behavior should result in faster adjustments toward equilibrium when shocks result in differentials that are outside the neutral band.

We then estimate a richer asymmetric threshold error correction model that includes lagged differentials of prices to represent short-run dynamics. A two-dimensional grid search is conducted to define the two thresholds c_1 and c_2. In particular, we search for the first threshold between 1% and 99% of the largest (in absolute value) negative error correction term. In like fashion, we search for the second threshold between 1% and 99% of the largest positive error correction term.[11] The error correction model is then estimated conditional on the threshold parameters.

A test of the statistical significance of the differences in parameters across alternative regimes is desirable. A standard test of parameter differences across regimes is equivalent to a conventional Chow test. As is well known, this testing problem is complicated by the fact that the alternative regime parameters are not identified under the null hypothesis of no threshold effects and thus conventional test statistics have nonstandard distributions.[12] 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 are used to approximate the asymptotic null 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 application is to daily corn and soybean prices observed at four important North Carolina terminal markets. In the case of corn, prices were quoted at Williamston, Candor, Cofield, and Kinston. For soybeans, prices were quoted at Fayetteville, Raleigh, Greenville, and Kinston. In each case, the largest markets in terms of volume (Williamston for corn and Fayetteville for soybeans) were taken as the central market against which the remaining three markets were compared.[13] Our evaluations are of a pair-wise nature; we compare prices in each market to the central market price. The prices were observed continuously between 2 January 1992 and 4 March 1999. A small number of price quotes were missing in each of the markets. On days where prices

[10] In that the cointegrating relationship represents an equilibrium where $\alpha = 0$ and $\beta = 1$ is expected, we consider price differentials as well as residuals from the cointegrating regression as error correction terms. As always, when residuals are used, the results may be sensitive to the normalization rule. In addition, it is important to note that the residuals are generated from a first stage regression and thus second stage standard errors may be understated in the models of the error correction terms.

[11] The grid search must be restricted to ensure adequate observations for estimating parameters in each regime. We constrained the search to require that each regime possess at least twenty-five observations. Lowering this bound to the number of parameters being estimated in each regime did not have a discernible effect on the results.

[12] Our grid search using the SSE criteria is equivalent to a sup-Wald Chow test approach, where the largest test statistic (i.e., smallest SSE) is used to define a break.

[13] Volume data are proprietary and thus were unavailable. Market analysts at the North Carolina Department of Agriculture ranked markets in terms of usual volumes.

Table 1. Local Market Characteristics

Market	Total Capacity[a]	Road Distance Between Markets[b]
	Corn Markets	
Williamston	1375	—
Candor	200	239.8
Cofield	2000	50.7
Kinston	750	90.2
	Soybean Markets	
Fayetteville	4500	—
Raleigh	1500	62.4
Greenville	750	136.2
Kinston	750	96.1

[a] Capacity is in thousand bushels and pertains to the total amount of grain or oilseeds (all commodities) that can be held at any one time. Specific commodity volumes and capacities are proprietary and were not available.
[b] Distances are measured as miles to the central market (Williamston for corn and Fayetteville for soybeans).

Table 2. OLS Estimates of Cointegrating Relationship: $P_t^1 = \alpha + \beta P_t^2$

Markets	α	β	R^2
Corn			
Candor–Williamston	0.1750	0.9018	0.980
	(0.0031)[a]	(0.0031)	
Corn			
Cofield–Williamston	0.0460	0.9911	0.974
	(0.0040)	(0.0039)	
Corn			
Kinston–Williamston	0.0159	0.9965	0.998
	(0.0011)	(0.0011)	
Soybeans			
Raleigh–Fayetteville	−0.0240	1.0119	0.998
	(0.0019)	(0.0010)	
Soybeans			
Greenville–Fayetteville	−0.0381	1.0004	0.988
	(0.0049)	(0.0026)	
Soybeans			
Kinston–Fayetteville	−0.0445	1.0033	0.987
	(0.0051)	(0.0027)	

[a] Numbers in parentheses are standard errors.

were missing in every market (typically holidays), the observations were omitted and a smooth continuity of prices was assumed. The remaining missing observations were replaced using cubic spline interpolation.[14] Characteristics of the markets, including the road distances between each market, are presented in table 1. Note that, with the exception of Candor, which is located in a separate region of the state, the regional markets are all in close proximity to the central market.

The empirical analysis is based upon logarithmic transformations of the prices. Standard unit-root tests confirmed a single unit root in each series. Ordinary least squares estimates of the cointegrating relationships are presented in table 2. In all but one case, the intercept terms are close to zero and the slope parameters are very near to one. Of course, in light of the nonstationary nature of the data, conventional testing procedures are inappropriate. However, we can note that the standard errors are very small in light of the large sample, suggesting conventional tests may reject $\alpha = 0$ and $\beta = 1$ in spite of the estimates' proximity to these values.

Cointegration testing results are presented in table 3. Balke and Fomby found that standard cointegration tests are also capable of detecting threshold cointegration.[15] Evidence of long-run equilibria among the pairs of prices is strong. In particular, cointegration among the pairs of prices is supported in every case, though evidence is weaker for corn price linkages between Kinston and Williamston. We also evaluated the stationarity of price differentials using standard augmented Dickey–Fuller (ADF) tests. The results are consistent with the cointegration tests, though the tests again indicate a lack of support for a stationary differential between Kinston and Williamston prices. As will be discussed in greater detail below, there is evidence suggesting that the basis relationship between Kinston and Williamston shifted in late 1998, perhaps influencing the cointegration tests. Confirmation of cointegration, in spite of the many weaknesses of such tests (McNew and Fackler), does indicate the existence of stable long-run equilibria among the prices and thus is consistent with integration.[16]

[14] The percentage of observations missing varied from 0.5% to 1.6% of the total sample.
[15] Balke and Fomby used the two-step cointegration testing procedures of Engel and Granger. In our application, results for the Engel and Granger tests were identical to those obtained with the Johansen procedures and thus are not presented here. Lag orders were chosen using AIC and SBC criteria. In that deterministic linear time trends did not appear to be present in the series, we restricted the intercept term to apply to the cointegration relationship only.
[16] Though, as Barrett has pointed out, cointegration need not necessarily imply integration since two series can be coincidentally cointegrated without implying economic integration.

Table 3. Cointegration Testing Results

Markets	Test	Test Statistic[a]
Corn	max eigenvalue test: $r = 0$	36.09**
Candor–Williamston	trace test: $r = 0$	37.9**
	max eigenvalue and trace test: $r = 1$	1.81
	ADF test of nonstationary differential	–3.95**
	LR test of $\alpha = 0, \beta = 1$	33.12**
Corn	max eigenvalue test: $r = 0$	45.50**
Cofield–Williamston	trace test: $r = 0$	47.73**
	max eigenvalue and trace test: $r = 1$	2.23
	ADF test of nonstationary differential	–5.17**
	LR test of $\alpha = 0, \beta = 1$	27.54**
Corn	max eigenvalue test: $r = 0$	12.63*
Kinston–Williamston	trace test: $r = 0$	14.44
	max eigenvalue and trace test: $r = 1$	1.81
	ADF test of nonstationary differential	–2.07
	LR test of $\alpha = 0, \beta = 1$	8.96**
Soybeans	max eigenvalue test: $r = 0$	49.35**
Raleigh–Fayetteville	trace test: $r = 0$	51.42**
	max eigenvalue and trace test: $r = 1$	2.08
	ADF test of nonstationary differential	–4.58**
	LR test of $\alpha = 0, \beta = 1$	9.2**
Soybeans	max eigenvalue test: $r = 0$	28.51**
Greenville–Fayetteville	trace test: $r = 0$	33.09**
	max eigenvalue and trace test: $r = 1$	1.96
	ADF test of nonstationary differential	–3.52**
	LR test of $\alpha = 0, \beta = 1$	25.18**
Soybeans	max eigenvalue test: $r = 0$	31.14**
Kinston–Fayetteville	trace test: $r = 0$	33.09**
	max eigenvalue and trace test: $r = 1$	1.96
	ADF test of nonstationary differential	–3.43**
	LR test of $\alpha = 0, \beta = 1$	27.49**

[a] The maximum eigenvalue and trace tests are Johansen cointegration tests. The ADF test is an Augmented Dickey–Fuller test of the stationarity of the logarithmic price difference. The LR test is a likelihood ratio test of $\alpha = 0, \beta = 1$ in the cointegrating regression. Single and double asterisks indicate statistical significance at the $\alpha = 0.10$ and $\alpha = 0.05$ levels, respectively. Critical values taken from Johansen and Nielsen.

We also utilized the maximum likelihood testing procedures of Johansen and Juselius to make inferences about the cointegrating vector implicit in the Johansen estimates. Though, as is reflected in table 1, the estimates were very close to the hypothesized values, the large number of observations and resulting low standard errors resulted in large likelihood ratio test statistics (table 3), thus rejecting the null hypothesis that $\alpha = 0$ and $\beta = 1$. Of course, the assumption that $\alpha = 0$ and $\beta = 1$ implies proportional transactions costs that are invariant with respect to price, thereby ruling out things such as nonconstant tariffs, risk premia, brokerage fees, etc.[17] Thus, rejecting the null hypothesis that $\alpha = 0$ and $\beta = 1$ is not surprising and should not be taken as evidence against integration.

Standard and threshold autoregressive (error correction) models of the price differentials and error correction terms for pairs of market prices were considered next. The specific form of the autoregressive model is

$$(5) \qquad \Delta e_t = \lambda e_{t-1} + v_t$$

where v_t represents a white-noise residual. As Obstfeld and Taylor note, this model has been used "countless times" in analyses of the law of one price. We allow λ to vary according to whether e_{t-1} is within (i.e., $|e_{t-1}| \leq c$) or outside (i.e., $|e_{t-1}| > c$) of a symmetric band

[17] We are grateful to Chris Barrett for pointing this out.

Table 4. Threshold Autoregression Model ($\Delta y_t = \lambda y_{t-1}$): Estimates and Test Results

Markets	AR λ	Half Life (Days)	TAR λ_{in}	TAR λ_{out}	Half-Life (Days)	TAR c	Tsay's Test	Hansen's Test
			Corn: Price Differential					
Candor–Williamston	−0.0085 (0.0031)[a]	81.41	−0.0003 (0.0038)	−0.0267 (0.0056)	25.59	0.0493	23.21**	15.27**
Cofield–Williamston	−0.0250 (0.0054)	27.40	−0.0153 (0.0055)	−0.1531 (0.0200)	4.17	0.0152	29.75**	44.02**
Kinston–Williamston	−0.0668 (0.0085)	10.02	−0.0202 (0.0084)	−0.4363 (0.0237)	1.21	0.0184	19.89**	273.18**
			Corn: Error Correction Term					
Candor–Williamston	−0.0778 (0.0092)	8.56	−0.0492 (0.0101)	−0.2002 (0.0209)	3.10	0.1263	22.56**	42.54**
Cofield–Williamston	−0.0562 (0.0079)	11.99	−0.0173 (0.0093)	−0.1496 (0.0143)	4.28	0.0132	29.75**	60.13**
Kinston–Williamston	−0.1868 (0.0138)	3.35	−0.0294 (0.0131)	−0.8075 (0.0260)	0.42	0.0294	19.89**	715.63**
			Soybeans: Price Differential					
Raleigh–Fayetteville	−0.2498 (0.0156)	2.41	−0.0537 (0.0168)	−0.6890 (0.0252)	0.59	0.0090	34.93**	439.61**
Greenville–Fayetteville	−0.0160 (0.0043)	42.92	−0.0261 (0.0069)	−0.0413 (0.0054)	16.43	0.0383	23.11**	58.60**
Kinston–Fayetteville	−0.0217 (0.0050)	31.54	−0.0033 (0.0052)	−0.1470 (0.0136)	4.36	0.0167	36.00**	97.36**
			Soybeans: Error Correction Term					
Raleigh–Fayetteville	−0.2952 (0.0168)	1.98	−0.0858 (0.0192)	−0.6789 (0.0259)	0.61	0.0090	70.74**	338.54**
Greenville–Fayetteville	−0.1152 (0.0043)	5.66	−0.0755 (0.0069)	−0.1761 (0.0054)	3.58	0.0273	23.04**	19.69**
Kinston–Fayetteville	−0.1510 (0.0126)	4.24	−0.0796 (0.0155)	−0.2762 (0.0206)	2.14	0.0217	36.56**	58.18**

[a] "AR" and "TAR" refer to autoregressive and threshold autoregressive models, respectively. λ_{in} and λ_{out} are autoregressive parameters within and outside of the neutral band, respectively. c is the estimated threshold parameter. Half-lives are the days required for one-half of a deviation from equilibrium to be eliminated.

[b] Numbers in parentheses are standard errors.

defined by the threshold c. The threshold autoregression models are estimated for both the price differential, which assumes $\alpha = 0$ and $\beta = 1$, and for the OLS residuals, representing error correction terms. In light of the OLS parameter estimates, we expect similar results for both models.

Tsay's test was conducted for the price differentials and the OLS residuals. Test results are presented in table 4. In every case, Tsay's test rejects the null hypothesis of no thresholds and thus suggests that threshold behavior characterizes spatial price linkages among the regional markets. Estimates of standard autoregressive models and threshold autoregressive models allowing for differential effects when lagged price differences (or error correction terms) lie outside of a threshold are presented in table 4. Deviation half-lives, given by $\ln(0.5)/\ln(1+\lambda)$, represent the period of time (in days) required for one-half of a deviation from equilibrium to be eliminated. Half-lives are presented for each model in table 4. Finally, the sum of squared error minimizing thresholds are presented in table 4. For $\beta = 1$ in the cointegrating regression (a condition that is guaranteed in the price differential models and is strongly supported in five of the six the OLS residual models), the thresholds represent the amount that proportional price differences must exceed to cross the threshold and trigger the "outside-band" regime adjustments.

In every case, the threshold models suggest much faster adjustments in response to deviations from equilibrium than is the case when threshold behavior is ignored. In particular, the threshold models indicate adjustments that are generally about twice as fast as those implied by standard autoregressive models. The error correction term (OLS residual) models yield similar results, though the residual models imply faster adjustments in many cases. The price differential models imply deviation half-lives ranging from 25.6 days for corn price linkages between Candor and Williamston to 0.6 days for soybean price linkages between Raleigh and Fayetteville. Likewise, the models based upon error correction terms calculated from OLS residuals imply half-lives ranging from 4.3 days for Cofield–Williamston corn price linkages to 0.4 days for Kinston–Williamston corn price linkages. In three of the four cases, the largest half-lives occur for those market pairs that are the most distant from one another. This result is consistent with expectations that greater physical separation between pairs of spatial markets should be correlated with slower adjustments to deviations from equilibrium. The estimated thresholds, which are analogous to transactions costs which define a neutral band, are also consistent with the patterns one would expect—the greater distance between markets corresponds to larger thresholds (transactions costs) and thus wider neutral bands.[18] In the price differential models, the largest threshold parameters are obtained for market combinations that are the most distant—Candor and Williamston (240 miles apart) for corn and Greenville–Fayetteville (136 miles apart) for soybeans. In particular, the thresholds imply that Candor and Williamston corn prices would have to be at least 4.9% different to exceed the threshold and trigger the faster adjustment. Likewise, Greenville and Fayetteville soybean prices would have to differ by at least 3.8% to trigger the faster "outside-band" adjustment.

The statistical significance of the alternative regimes indicated by the thresholds was evaluated using Hansen's modification of standard Chow-type tests. Two-hundred replications of the null test statistic were used to determine approximate probability values associated with the nonstandard test statistics.[19] In every case, the test statistics exceeded the implied critical values at the $\alpha = 0.01$ or smaller level. In all, the estimates confirm that significant thresholds exist and imply that these thresholds have an important role in characterizing regional price adjustments. In particular, adjustments to deviations from conditions generally considered to represent market integration (i.e., price parity) are found to be much faster when the deviations exceed a threshold. Small deviations that are not large enough to result in differentials (or errors) that exceed the threshold imply very slow adjustments, as is reflected in the autoregressive parameters that are close to zero.[20]

A richer threshold vector error correction model was also considered to investigate short-run dynamics and to allow for asymmetric adjustments in response to positive and negative price shocks. Commodity flows between markets often occur primarily in one direction, suggesting some markets are usually exporters and other markets are typically importers. Thus, transportation infrastructure and handling facilities may be better suited for commodity flows in one direction. In such a case, thresholds may be asymmetric as shipment in one direction is more costly than shipments in the opposite direction.

A two-variable error correction model of the form given by equation (3) was estimated using the two-dimensional grid search procedure described above. Estimation and testing results are presented in table 5. Hansen's modified Chow test of the significance of the parametric differences across regimes is again highly significant in every case, confirming the significance of the threshold effects. Tsay's nonlinearity tests for the error correction terms implied by the OLS residuals (table 4) are again relevant to these results.

[18] It should be noted that the threshold for the OLS residual model of Candor–Williamston corn market linkages is much larger than the thresholds implied by other models. This reflects the unusually small price transmission parameter $\beta = 0.90$.

[19] The number of replications was necessarily limited by the long computing time required for a single iteration. The grid search was modified by searching for each threshold conditional upon the other. This was repeated, alternating between each threshold, until estimates converged. A comparison to the full two-dimensional grid search revealed identical threshold estimates.

[20] In fact, the autoregressive terms, λ, within the thresholds are often very close to zero, suggesting a lack of cointegration between the two prices within the neutral band implied by the threshold.

Table 5. Threshold Estimation and Testing Results

Markets	Test	Test Statistic[a]
Corn	Regime I (no. obs.)	$-\infty < \nu_{t-1} \leq -0.0255$ (295)
Candor–Williamston	Regime II (no. obs.)	$-0.0255 < \nu_{t-1} \leq 0.0073$ (761)
	Regime III (no. obs.)	$0.0073 < \nu_{t-1} < \infty$ (716)
	Hansen's threshold test	77.54**
Corn	Regime I (no. obs.)	$-\infty < \nu_{t-1} \leq -0.0572$ (69)
Cofield–Williamston	Regime II (no. obs.)	$-0.0572 < \nu_{t-1} \leq 0.0594$ (1669)
	Regime III (no. obs.)	$0.0594 < \nu_{t-1} < \infty$ (35)
	Hansen's threshold test	48.56**
Corn	Regime I (no. obs.)	$-\infty < \nu_{t-1} \leq -0.0125$ (249)
Kinston–Williamston	Regime II (no. obs.)	$-0.0125 < \nu_{t-1} \leq 0.0178$ (1469)
	Regime III (no. obs.)	$0.0178 < \nu_{t-1} < \infty$ (55)
	Hansen's threshold test	65.67**
Soybeans	Regime I (no. obs.)	$-\infty < \nu_{t-1} \leq -0.0060$ (166)
Raleigh–Fayetteville	Regime II (no. obs.)	$-0.0060 < \nu_{t-1} \leq 0.0103$ (1559)
	Regime III (no. obs.)	$0.0103 < \nu_{t-1} < \infty$ (47)
	Hansen's threshold test	62.28**
Soybeans	Regime I (no. obs.)	$-\infty < \nu_{t-1} \leq -0.0102$ (410)
Greenville–Fayetteville	Regime II (no. obs.)	$-0.0102 < \nu_{t-1} \leq 0.0216$ (1292)
	Regime III (no. obs.)	$0.0216 < \nu_{t-1} < \infty$ (70)
	Hansen's Threshold Test	61.84**
Soybeans	Regime I (no. obs.)	$-\infty < \nu_{t-1} \leq -0.0060$ (544)
Kinston–Fayetteville	Regime II (no. obs.)	$-0.0060 < \nu_{t-1} \leq 0.0070$ (508)
	Regime III (no. obs.)	$0.0070 < \nu_{t-1} < \infty$ (721)
	Hansen's threshold test	63.73**

[a] Single and double asterisks indicate statistical significance at the $\alpha = 0.10$ and $\alpha = 0.05$ levels, respectively.

In particular, significant threshold behavior in the error correction terms of these models is implied by Tsay's test.

The thresholds are reasonably symmetric about zero. In five of the six cases, the regime defined by being between the two thresholds dominates in terms of the frequency of observations. Such a result is to be expected, implying deviations from equilibrium (as represented by the error correction terms) are seldom large enough to exceed the neutral band implied by the thresholds. The neutral band implied by the thresholds is widest for the Cofield–Williamston corn markets and for the Greenville–Fayetteville soybean markets. This is somewhat surprising given the Cofield and Williamston markets are close to one another, while the Greenville and Fayetteville markets are the most separated of the pairs of soybean markets considered. This result is consistent with the threshold autoregressive models for soybeans. The threshold for Cofield and Williamston corn price linkages is much wider in the bivariate threshold error correction model and may reflect other market characteristics influencing price transmission and dynamics. As expected, in most cases the thresholds indicate that the distribution of the error correction terms is heavily skewed toward one side of zero. This reflects a typical basis differential where one market's prices are usually above (or below) another market's prices.[21] The central market against which comparisons are made typically has a lower price. Parameter estimates (not presented here) indicated significant dynamic relationships among the price series. Error correction terms were especially significant, confirming the cointegrated nature of the price series.

Comparison of the estimated thresholds as one moves from the symmetric TAR models to the asymmetric threshold error correction models reveals some interesting results worthy of further discussion. In four of six cases,

[21] Of course, the error correction terms are constructed from OLS residuals and thus must have a mean of zero. The error correction terms only represent price differences in cases where $\alpha = 0$ and $\beta = 1$ in the cointegrating regression.

the thresholds for the asymmetric bands lie within those implied for the symmetric bands. This tendency does not appear to hold for the Cofield–Williamston corn market relationship and the Raleigh–Fayetteville soybean market comparison (though the bands are very close in the latter case). This may suggest that restricting the thresholds to be symmetric may be unnecessarily restrictive, causing the neutral-band sizes to be overstated. However, in most cases, the differences in the bands are slight. An exception exists for the Candor–Williamston relationship for corn. In this case, the asymmetric model revealed a much smaller neutral band. This may be related to the relatively low (0.90) price transmission elasticity estimate (β in the cointegrating relationship), which may imply a nonconstant relationship between prices and proportional transactions costs. It is again relevant to point out that this pair of markets is separated by considerably more distance than any other pair considered.

Figure 1 illustrates the timing of switching between regimes implied by the asymmetric threshold error correction model. Recall that the model consists of three regimes (denoted I, II, and III). A particular regime characterizes the relationship among the prices at a point in time. Each observation, depending upon the value of the error correction term, falls into one of the three regimes. The figure illustrates which of the three regimes each observation falls into and thus the timing of switching among the regimes. Recall that Regime II corresponds to observations falling between the two thresholds and thus suggests price differences that are within the neutral bands. A long series of observations falling outside of the implied neutral band (i.e., in Regimes I or III) suggests price differences that persistently exceeded transactions costs. Strongly integrated markets would be expected to not display such persistent deviations from equilibrium.

The greatest degree of switching among regimes and most frequent occurrence of price differences exceeding the neutral bands is realized generally by those markets that are the most distant. This is particularly the case for the Candor and Williamston corn markets. Candor is in a different geographic region of the state and is located away from major producing and consuming areas. Given its role as a "satellite" market, it is not surprising to see more frequent switching among regimes being implied by price differences outside of the neutral band. Considerable regime switching in the relationships between Fayetteville and the Greenville and Kinston soybean markets is also reflected. Finally, it is interesting to note that between mid- 1998 and 1999 the Kinston–Williamston corn market linkage was characterized almost exclusively by the first regime. Recall that this market pair was the only one for which standard cointegration tests failed to support a stable equilibrium. This may reflect a structural change or other shift in market infrastructure that affected price linkages between these markets.

Interpretation of the dynamic interrelationships among prices at alternative markets is best pursued through a consideration of impulse response functions. In contrast to the linear model case, the response to a shock is dependent upon the history of the series and the possibility of asymmetric responses implies that the size and sign of the shock will influence the nature of the response. Consequently, there are many different possible impulse response functions. We chose a single observation (observation 1773, corresponding to the last observation in our data, 4 March, 1999) to evaluate responses to one-half standard deviation positive and negative shocks. We adopt the nonlinear impulse response function approach of Potter and Koop, Pesaran, and Potter, which defines responses (denoted I_{t+k}) on the basis of observed data $(z_t, z_{t-1}, \dots)$ and a shock (υ) as

$$(6) \qquad I_{t+k}(\upsilon, Z_t, Z_{t-1}, \dots) = E[Z_{t+k} \mid Z_t = z_t + \upsilon, Z_{t-1} = z_{t-1}, \dots] - E[Z_{t+k} \mid Z_t = z_t, Z_{t-1} = z_{t-1}, \dots].$$

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.

Figures 2 and 3 illustrate responses to one-half standard deviation positive and negative shocks in each of the markets. Note that each panel illustrates responses to both positive and negative price shocks (denoted by (+) and (–), respectively). To the extent that asymmetries exist in alternative regimes, the potential exists for asymmetric responses. Figure 2 illustrates responses

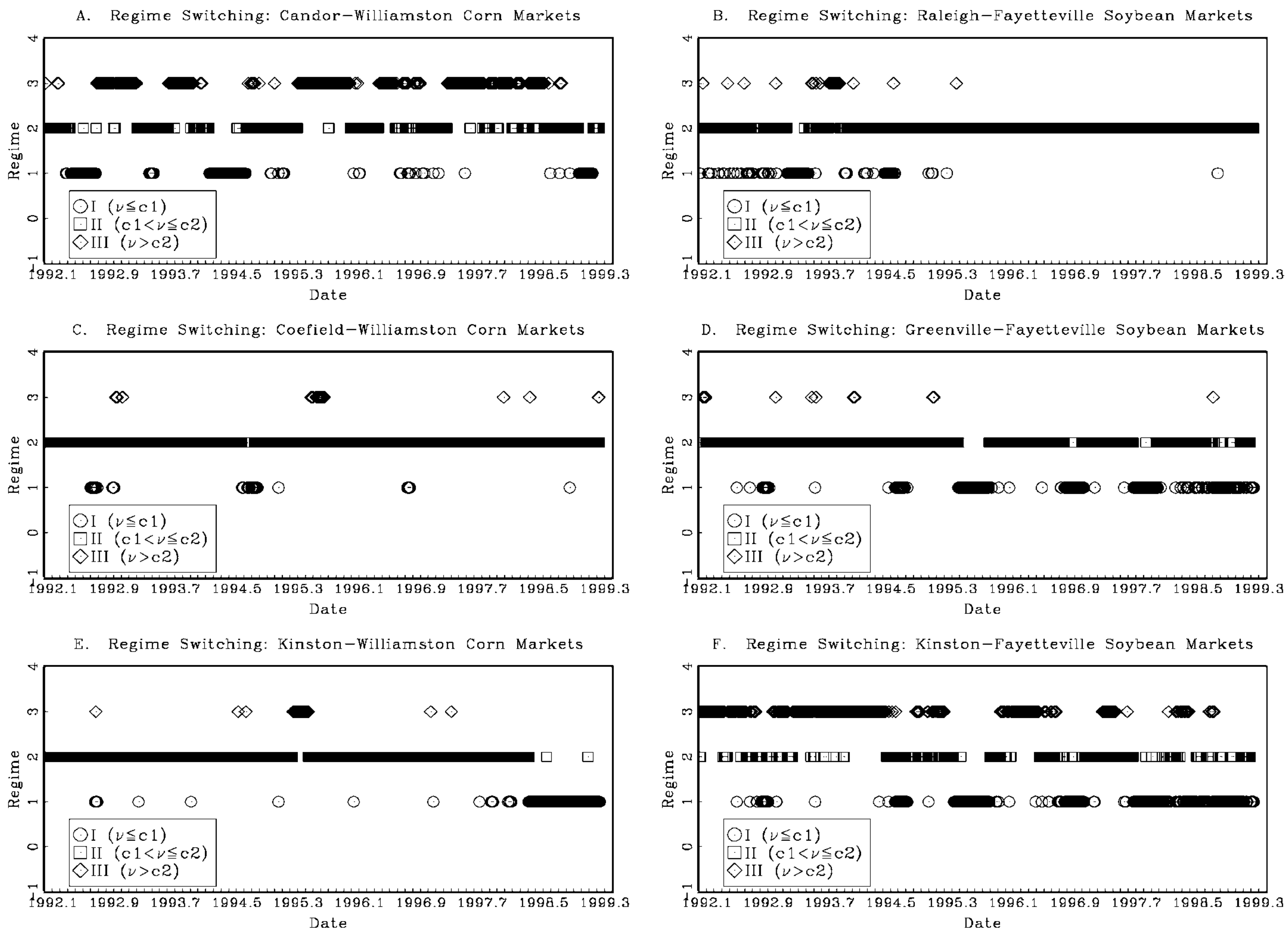

Figure 1. The timing of switching between regimes

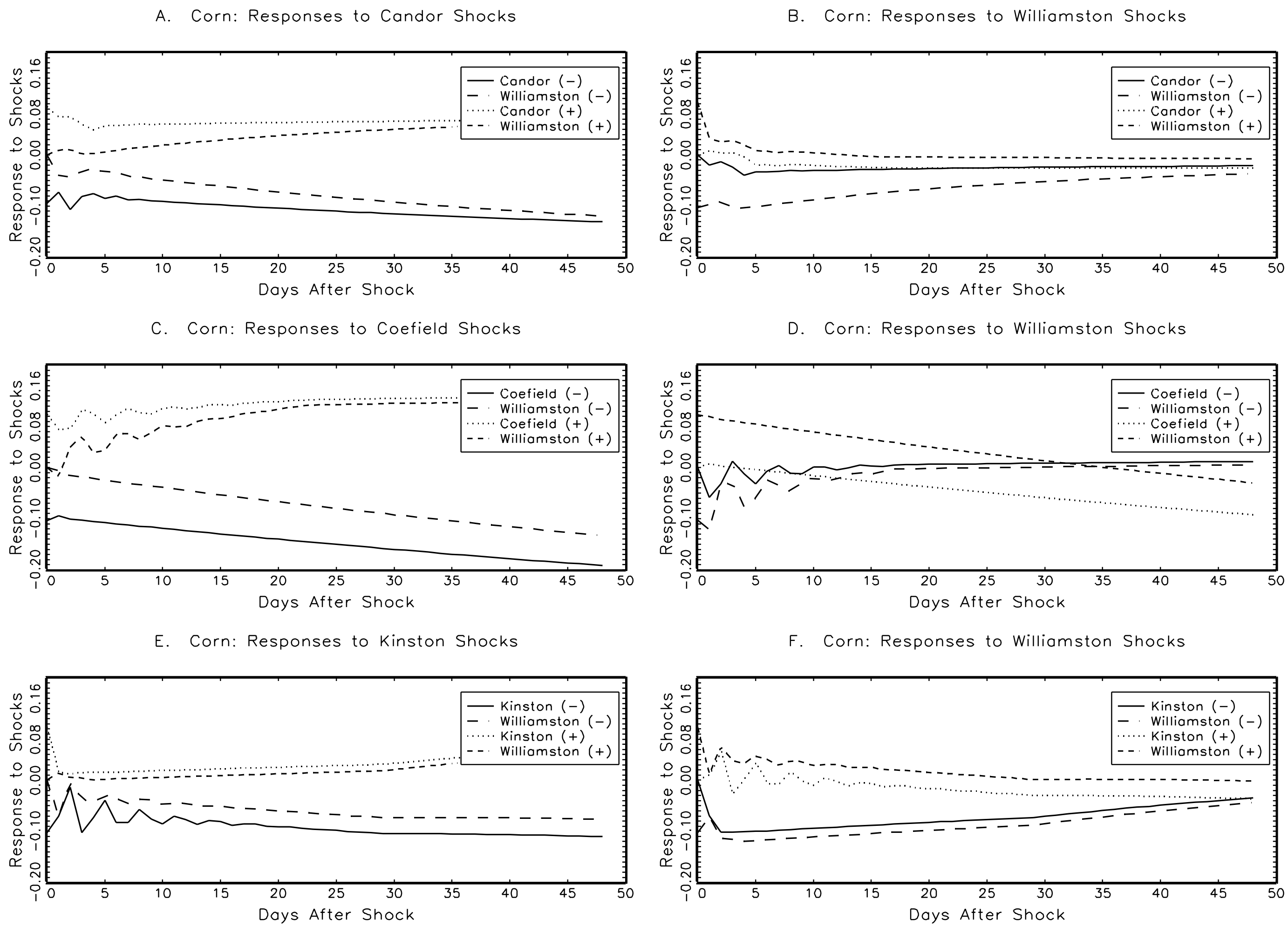

Figure 2. Nonlinear impulse response functions (observation 1773): corn

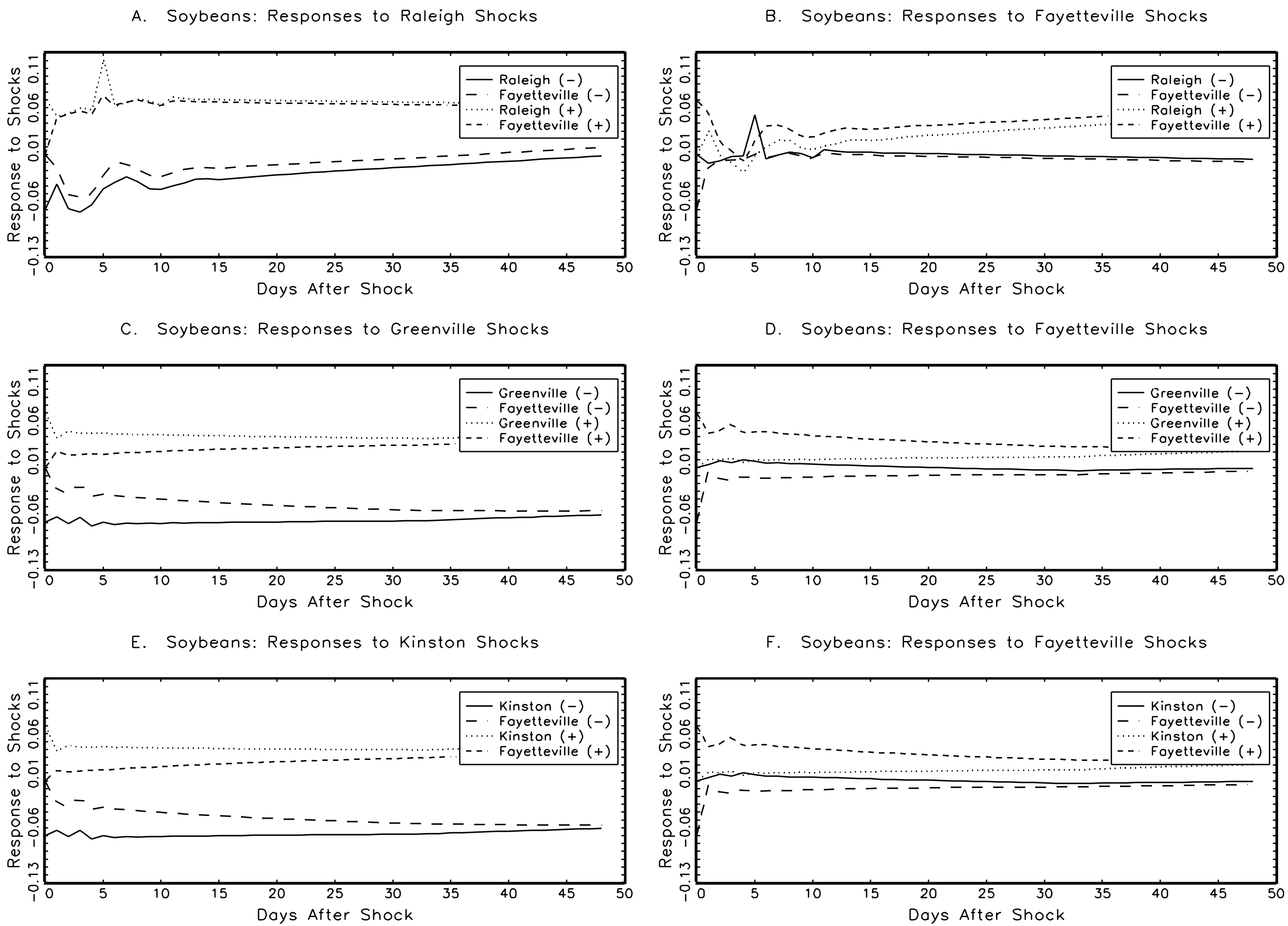

Figure 3. Nonlinear impulse response functions (observation 1773): soybeans

to shocks in corn markets while figure 3 illustrates responses to shocks in soybean markets. The responses are highly consistent with long-run market integration. Isolated shocks in one market provoke responses in other markets that eventually lead to a tendency for prices to equalize. For example, a positive shock to the Candor corn price (Panel A) evokes an equilibrating response in the Williamston price that eventually (approximately twenty days later) leads to price convergence. A negative shock (also in Panel A) leads to a similar pattern of convergence, with both prices exhibiting a tendency to be permanently decreased as a result of the negative shock. In most cases, the shocks result in permanent shifts in the price series, reflecting the nonstationary nature of the price data. Thus, positive shocks often lead to permanent price increases while negative shocks often lead to permanent price decreases. The responses suggest that, after some short run dynamic adjustments, the prices converge to one another over the long run (i.e., generally after ten to fifteen days following the shock). Evidence of asymmetries in price adjustments is limited. In most cases the responses to negative shocks, though naturally of an opposite sign, are quite similar to the corresponding responses to positive shocks.

Perhaps most important for market integration is the finding that prices quickly converge following an isolated, exogenous shock to one of the price series. Although the prices may not converge toward absolute equality for many periods, the impulses do reflect behavior consistent with price convergence. It should be noted that one-half standard deviation shocks are relatively large and are not likely to be observed in day to day price movements. Responses to very small shocks (i.e., shocks that are not large enough to push the price differential outside of the neutral band between the thresholds) resulted in much different responses.[22] In many cases, these responses were not consistent with stable price adjustments. Correlation between prices within the neutral band may be largely spurious and thus may not imply equilibrating behavior.

In all, the impulse responses are in agreement with expectations and provide strong support for integrated markets. Responses to market shocks are generally complete after fifteen days. Responses are generally as one would expect, with positive shocks eliciting positive responses and negative shocks eliciting negative responses. Most shocks result in permanent shifts in prices, reflecting the nonstationary nature of the price data.

Concluding Remarks

We have evaluated spatial price linkages and daily price dynamics among regional corn and soybean markets in North Carolina utilizing asymmetric, threshold autoregression and error correction models. Our results confirm that such markets are tightly integrated. Our analysis confirms the significance of threshold effects and suggests that their presence may significantly influence spatial price linkages. Models that explicitly recognize threshold behavior imply much faster adjustments to deviations from equilibrium than is the case when thresholds are ignored. We utilize nonlinear impulse response functions to evaluate the dynamic paths of adjustment to exogenous, localized shocks. The responses confirm equilibrating responses consistent with price equalization and integration of markets. Adjustments are generally complete after fifteen days. In every case, the threshold models suggest much faster adjustments in response to deviations from spatial equilibrium than is the case when threshold behavior is ignored. While modest asymmetries are revealed, positive and negative shocks generally yield symmetric responses.

Our analysis was motivated in part by unobservable transactions costs and the important influence that their presence may exert on equilibrium spatial price relationships. In particular, we noted that transactions costs may result in a neutral band within which markets are well integrated even though their prices are not directly linked. As has been widely recognized in recent years, the large literature which has applied regression-based tests of market integration may provide misleading inferences when the role of transactions costs is ignored. Our analysis addresses this shortcoming, albeit in a very restrictive way. In particular, we allow for regime switching which may occur when shocks result in price differences exceeding transactions costs and thus permit profitable trade. It is important to note, however, that, like all other studies attempting to account

[22] Results for alternative periods and shocks are available from the authors on request.

for unobservable transactions costs, our analysis relies on several restrictive assumptions. In particular, our thresholds are estimated as constants, which implies that the neutral band is fixed (in proportional terms). This assumption may be reasonable for the relatively short period of time spanned by our data. It is possible, however, that nonstationary transactions costs could result in a nonconstant neutral band. To explicitly address such a situation, one needs direct observations of actual transactions costs.[23] Other modeling procedures which allow thresholds to vary according to some other external criteria are certainly conceivable and serve as an important direction for future research.

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References

Ardeni, P.G. "Does the Law of One Price Really Hold for Commodity Prices?" *Amer. J. Agr. Econ.* 71(August 1989):661–69.

Baffes, J. "The Law of One Price Still Holds." *Amer. J. Agr. Econ.* 73(November 1991): 1264–73.

Balke, N.S., and T.B. Fomby. "Threshold Cointegration." *Int. Econ. Rev.* 38(August 1997):627–45.

Barrett, C.B. "Market Analysis Methods: Are Our Enriched Toolkits Well-Suited to Enlivened Markets?" *Amer. J. Agr. Econ.* 78(August 1996):825–29.

Baulch, B. "Transfer Costs, Spatial Arbitrage, and Testing For Food Market Integration." *Amer. J. Agr. Econ.* 79(May 1997):477–87.

Chow, G. "Tests of Equality Between Sets of Coefficients in Two Linear Regressions." *Econometrica* 28(July 1960):591–605.

Dickey, D., and W. Fuller. "Distribution of the Estimators for Autoregressive Time Series with a Unit Root." *J. Amer. Statist. Assoc.* 74(June 1979):427–31.

Enders, W., and C.W.J. Granger. "Unit Root Tests and Asymmetric Adjustment with an Example Using the Term Structure of Interest Rates." *J. Bus. and Econ. Statist.* 16(July 1998): 304–12.

Engle, R.F., and C.W.J. Granger. "Cointegration and Error Correction: Representation, Estimation, and Testing." *Econometrica* 55(March 1987):251–76.

[23] For an example of threshold modeling with observable transactions costs, see Goodwin and Grennes.

Fackler, P., and B.K. Goodwin. "Spatial Market Integration." *Handbook of Agricultural Economics*. Amsterdam: North Holland, in press.

Goodwin, B.K., and T.C. Schroeder "Cointegration Tests and Spatial Price Linkages in Regional Cattle Markets." *Amer. J. Agr. Econ.* 73(May 1991):452–64.

Goodwin, B.K., and T.J. Grennes. "Tsarist Russia and the World Wheat Market." *Explorations Econ. Hist.* 35(1998):405–30.

Hansen, B.E. "Inference in TAR Models." *Stud. Nonlinear Dynamics Econometrics* 2(April 1997) (online).

Heckscher, E.F. "Växelkursens Grundval vid Pappersmyntfot." *Ekonomisk Tidskrift* 18(October 1916):309–12.

Johansen, S. "Statistical Analysis of Cointegration Vectors." *J. Econ. Dynamics Control* 12(1988): 231–54.

Johansen, S., and K. Juselius. "Testing Structural Hypotheses in A Multivariate Cointegration Analysis of the PPP and the UIP for UK." *J. Econometrics* 53(1992):211–44.

Johansen, S., and B. Nielsen. "Asymptotics for Cointegration Rank Tests in the Presence of Intervention Dummies—User's Manual for the Simulation Program DISCO." Unpublished manuscript, Institute of Mathematical Statistics, University of Copenhagen.

Koop, G., M.H. Pesaran, and S.M. Potter. "Impulse Response Analysis in Nonlinear Multivariate Models." *J Econometrics* 74(1996):119–47.

Li, J.R., and C.B. Barrett. "Distinguishing Between Equilibrium and Integration in Markets Analysis." Unpublished manuscript, Cornell University, June 1999.

Martens, M., P. Kofman, and T.C.F. Vorst. "A Threshold Error-Correction Model for Intraday Futures and Index Returns." *J. Appl. Econometrics* 13(May–June 1998):245–63.

McNew, K., and P. Fackler. "Testing Market Equilibrium: Is Cointegration Informative?" *J. Agr. Resour. Econ.* 22(December 1997): 191–207.

Obstfeld, M., and A.M. Taylor. "Nonlinear Aspects of Goods-Market Arbitrage and Adjustment; Heckscher's Commodity Points Revisited." *J. Japanese Int. Econ.* 11(December 1997): 441–79.

Potter, S.M. "A Nonlinear Approach to US GNP." *J. Appl. Econometrics* 10(April–June 1995): 109–25.

Sexton, R., C. Kling, and H. Carman. "Market Integration, Efficiency of Arbitrage, and Imperfect Competition: Methodology and

Application to U.S. Celery Markets." *Amer. J. Agr. Econ.* 73(1991):568–80.

Siklos, P.L., and C.W.J. Granger. "Regime Sensitive Cointegration: With an Application to Interest Rate Parity." *Macroecon. Dynamics* 1, no. 3 (1997) (online journal: http://wueconb.wustl.edu/barnett/MD.html).

Spiller, P., and Wood R.O. "The Estimation of Transaction Costs in Arbitrage Models." *J. Econometrics* 39(1988):309–326.

Tong, H. *On a Threshold Model in Pattern Recognition and Signal Processing.* C. Chen, ed. Amsterdam: Sijhoff and Noonhoff, 1978.

Tsay, R.S. "Testing and Modeling Threshold Autoregressive Processes." *J. Amer. Stat. Assoc.* 84(March 1989):231–40.