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Estimating Structural Changes in the Vertical Price Relationships in U.S. Beef and Pork Markets

Brenda L. Boetel and Donald J. Liu

This paper examines structural breaks in the vertical price relationships in U.S. beef/cattle and pork/hog sectors using monthly data of the past 40 years. A major methodological issue addressed is how to estimate price relationships when data contain intermittent structural breaks with unknown break dates. The adopted procedures endogenously search for structural break dates while explicitly accounting for this search in statistical inferences. Four breaks for the beef/cattle price relationship and three breaks for the pork/hog price relationship are identified. The estimation results further confirm the importance of allowing for structural breaks in the analysis of vertical price relationships.

Key Words: farm cattle and hog prices, long-run price relationship, retail beef and pork prices, structural breaks of unknown timing, structural changes, vertical price relationship

Introduction

Perhaps one of the most powerful insights economists have offered is the idea that prices in a well-functioning market convey sufficient information to allow efficient resource allocation. Of particular interest to agricultural economists are long-run price relationships among major agricultural products and food items across various market dimensions such as time and space, and how prices adjust in the short run to move toward long-run equilibrium. Given that price analysis often involves long time series, analysts must be mindful of structural changes in the data-generating process. Omitting structural breaks in the estimation may result in biased estimates of the price relationships. In this paper, we reexamine retail-wholesale-farm price relationships in U.S. beef and pork markets, paying attention to the incorporation of structural breaks in the estimation.

The U.S. beef and pork sectors have undergone significant changes in the past several decades, including changes in consumer preferences for red meat, changes in vertical firm relationships through the increased use of production and marketing contracts, changes in the extent of asset fixity (especially within various stages of hog production), and changes in the ways producers promote consumption of their products (e.g., via check-off programs). Theory suggests these and other changes in the beef and pork sectors can affect how prices relate to one another at the various stages of the supply chain (Gardner, 1975). This structural change in the retail-wholesale-farm price relationship has welfare implications for consumers, processors, and producers, all of whom are concerned with obtaining their fair share of the economic surplus.

Brenda L. Boetel is associate professor, Department of Agricultural Economics, University of Wisconsin-River Falls, and Livestock Marketing Specialist for the University of Wisconsin Extension System. Donald J. Liu is professor and Morse-Alumni Distinguished Teaching Professor, Department of Applied Economics, University of Minnesota. Funding for this project was provided in part by the Minnesota Agricultural Experiment Station, Grant No. MIN-14-057. The authors are grateful to Kenji Adachi for invaluable advice on this research. We also thank Molly Wieland for her editorial remarks. Finally, the authors express appreciation to editor Gary Brester for helpful suggestions and comments.

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A major methodological issue addressed in this paper is how to estimate price relationships when data contain intermittent structural breaks with unknown break dates. The conventional method of specifying, a priori, a known structural break date to account for a change in the economic relationship in question (i.e., Chow test) has been replaced by procedures involving direct estimation of break timing (Hansen, 2001). The motivation for this innovation is two-fold. First, the subtlety of a structural break may lead to uncertainty about its timing. Second, even when there is convincing evidence that a structural change has occurred at a specific time, the effects of data mining on the legitimacy of subsequent statistical inferences still must be considered. Our procedures endogenously search for structural break dates while explicitly accounting for this search in statistical inferences.

Literature Review

Previous price transmission analyses address how price shocks in one market transmit asymmetrically to prices in other markets along the supply chain or in other geographical regions, depending on the magnitude of the shock (Kinnucan and Forker, 1987; Boyd and Brorsen, 1988; Bailey and Brorsen, 1989). Analysts have increasingly responded to concerns about price series nonstationarity by employing the cointegration framework for price transmission analysis (Ardeni, 1989; Goodwin and Schroeder, 1991); more recent studies allow for asymmetry in short-run adjustments toward the long-run price relationship (Abdulai, 2002; Goodwin and Piggott, 2001; Goodwin and Harper, 2000; Goodwin and Holt, 1999). Meyer and von Cramon-Taubadel (2004) provide a helpful review of price asymmetry studies in agricultural and commodity markets.

Researchers have also focused their attention on structural changes in price relationships. Tiffin and Dawson (2000) test for cointegration between retail and farm lamb prices in the United Kingdom, with an allowance for a single abrupt structural break in the price relationship. Goodwin (1992) employs a gradual switching vector autoregression to investigate structural change in short-run price dynamics among cattle, hog, and other commodity prices, allowing the parameter to gradually transition between regimes in accordance with a hyperbolic tangent function. Within a more general time-varying smooth transition autoregressive framework, Holt and Craig (2006) and Balagtas and Holt (2009) investigate, respectively, structural changes in the U.S. hog-corn price relationship and the price relationships among selected primary commodities and manufactured goods.

The most comprehensive studies of vertical price relationships in the U.S. beef and pork industries can be attributed to Goodwin and Holt (1999) and Goodwin and Harper (2000). Using weekly data, the authors investigate retail, wholesale, and farm price relationships by estimating a long-run price transmission equation and a short-run price dynamic system among the cointegrated price series. The authors focus on the asymmetry of the error correction adjustment brought about by an unknown threshold and conclude that the degree of price adjustment depends on shock size. In addition, they also find asymmetry in price causality; shocks are transmitted from farm to wholesale to retail markets, but not in the opposite direction. Recognizing asymmetry in price transmission allows analysts to gain additional insights into how prices interact in different regimes. However, the above studies do not acknowledge potential structural breaks in the price relationship.

This study expands the contributions of Goodwin and Holt (1999), Goodwin and Harper (2000), and other authors (e.g., Boyd and Brorsen, 1988; Hahn, 1990) in several ways. Utilizing recent econometric innovations (Perron, 1997; Gregory and Hansen, 1996; Bai and

Perron, 1998; Kejriwal and Perron, 2008, 2009), multiple structural changes of unknown timing are entertained explicitly in the estimation of long-run price relationships as well as in the unit root analysis and cointegration tests. We also adopt a dynamic least squares procedure introduced by Phillips and Loretan (1991) to ensure the unbiasedness of the estimated long-run price transmission parameter (which is of intrinsic interest to this study).[1] Finally, our analysis provides additional insights over the results reported by Goodwin and Holt and Goodwin and Harper, as we utilize monthly data spanning an almost 40-year period.[2] It is essential for sound policy analysis that the estimated long-run price transmission parameter be unbiased and reflective of the mutability of the underlying price relationship in question.

Estimating a Long-Run Price Relationship with Structural Breaks of Unknown Timing

We consider an equilibrium economic relationship among a set of stationary or cointegrated nonstationary variables characterized by a linear model of $y_t = a + \mathbf{b}\mathbf{x}_t + \varepsilon_t$. The evolution of this long-run relationship can be parsimoniously captured by occasional structural breaks where all or part of the parameters are permanently changed to a different level at the break dates. Bai and Perron (1998) propose a procedure to estimate multiple break dates globally (simultaneously) for the case in which both y and $\mathbf{x}$ are stationary. In particular, the regression coefficients can be obtained by OLS for each choice of possible break date configurations, and the optimal break dates are the break date configuration that yields the smallest sum of squared residuals.[3] To ease the computation burden, especially when there are a high number of breaks, Bai and Perron (2003a) present an efficient estimation algorithm based on the principle of dynamic programming.

Bai and Perron (1998) demonstrate that the estimated break fractions are super-consistent in that they converge to the true fraction at a fast rate of T (with T being the sample size),[4] making it possible to obtain the standard root-T consistency and asymptotic normality for the estimates of other parameters in the regression. As to the limiting distribution of the break date estimates, the authors defer to Bai (1997), who estimates multiple breaks one at a time. In testing for the number of breaks, Bai and Perron (1998) consider sup Wald tests for: (*a*) the null hypothesis of no structural change versus the alternative hypothesis of a prespecified number of breaks, and (*b*) a specific-to-general sequential procedure of testing for k versus $k + 1$ breaks to consistently determine the number of breaks in the data. Asymptotic critical values for the test statistics are reported in Bai and Perron (1998, 2003b), and Gauss code for the numerical implementation of Bai and Perron's (2003a) dynamic-programming based procedure can be found on Perron's website (Perron, 2004).

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[1] Stock (1987) demonstrates that the conventional ordinary least squares (OLS) procedure results in a biased estimate of the slope parameter in a cointegration equation. However, the author points out that the OLS estimate is super-consistent in that it converges to its true value at a very fast rate of T (with T being the sample size). The usual consistency requires only a convergence rate of $T^{1/2}$.

[2] Goodwin and Holt (1999) use data from January 1981 to the first week of March 1998, while Goodwin and Harper's (2000) data range from January 1987 to the first week of January 1999.

[3] Consider the case of m breaks in which an a priori trimming rule dictates the minimum length of each of the $m + 1$ segments, and hence possible candidates for the m break dates. For each allowable configuration of the m breaks, the regression coefficients are obtained by minimizing the sum of squared residuals across all $m + 1$ segments. The chosen configuration of the m break dates is the one that has the lowest minimized sum of squared errors among all the allowable configurations.

[4] The ith break fraction is defined as T_i / T, where T_i is the ith break date (e.g., 50th observation out of a sample size of 200).

Kejriwal and Perron (2008) extend Bai and Perron's (1998, 2003a) multiple breaks estimation procedure to allow for the case in which y and some or all of the variables in $\mathbf{x}$ are nonstationary but cointegrated. As in Bai and Perron (2003a), the estimation procedure involves an algorithm based on the principle of dynamic programming. However, the distributions of the break date estimates and the sup Wald statistics are different due to the nonstationarity of the time series. Asymptotic critical values for the test statistics are reported in Kejriwal and Perron (2009).

Unit Root and Cointegration Tests with a Structural Break of Unknown Timing

This study follows the procedures suggested by Perron (1997) and Gregory and Hansen (1996) to conduct unit root and cointegration tests. Lest analysts confuse a structural break with nonstationarity, Perron extends the conventional augmented Dickey-Fuller (ADF) unit root test to allow for a structural break in the trend function of a time series in the following manner:

$$(1) \qquad y_t = \left\{ \alpha_1 + \beta_1 t + \rho y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + e_t \right\} + (\alpha_2 + \beta_2 t) D_t(\tau) + \delta B_t(\tau),$$

where the braced terms pertain to the ADF test, with $\alpha_1 + \beta_1 t$ being the trend function and e_t the error term. The null hypothesis of a unit root is rejected if ρ, the coefficient of the lagged dependent variable, is found to be less than unity. $B_t(\tau)$ is a dummy variable (equal to 1 when $t = \tau + 1$, and 0 otherwise) representing a one-period bump in the drift rate under the null of $\rho = 1$; $D_t(\tau)$ is a dummy variable (equal to 1 when $t > \tau$, and 0 otherwise) which captures a shift in the trend function at time τ under the alternative hypothesis that $\rho < 1$.[5]

Subject to trimming of observations at both ends of the sample, the break date in (1) can be obtained via a search procedure based on the criterion of minimizing the t-statistic associated with the parameter of the lagged dependent variable (so as to maximize the power of rejecting the null). For any given break date candidate, the model is linear, and thus can be estimated by OLS. Perron (1997) derives asymptotic distributions of the test statistic and simulates critical values for alternative structural break specifications, such as allowing for only a shift in the intercept or a shift in both the intercept and slope of the trend function.

In a similar manner, Gregory and Hansen (1996) are concerned about analysts confusing a shift in long-run relationship with a lack of relationship when conducting cointegration tests among nonstationary variables. The authors extend Engle and Granger's (1987) conventional test to allow for a single structural break of unknown timing under both the null and alternative hypotheses:

$$(2) \qquad y_t = \left\{ a_1 + \mathbf{b}_1 \mathbf{x}_t + \mu_t \right\} + \left(a_2 + \mathbf{b}_2 \mathbf{x}_t \right) D_t(\tau),$$

where the braced terms correspond to the conventional cointegration equation, whose intercept and slope are subject to a structural break at time τ. The null hypothesis that y_t and $\mathbf{x}_t$ are not cointegrated is rejected if the residual series $\{\mu_t\}$ does not contain a unit root. For each possible break date candidate (subject to endpoint trimming), the authors estimate (2)

[5] Specifically, under the null hypothesis of $\rho = 1$: α_1 is the drift rate, which is subject to a one-time bump of the magnitude δ, and $\alpha_2 = \beta_1 = \beta_2 = 0$. Alternatively, when $\rho < 1$: α_1 and β_1 are the intercept and slope of the trend function, which are subject to permanent changes (of the magnitude α_2 and β_2, respectively), and δ is zero.

and conduct cointegration tests via checking for nonstationarity of the residuals, involving a specification similar to the one in the braced terms in (1).[6] In choosing the break date, the authors minimize the *t*-statistic associated with the lagged residual coefficient in the unit root equation and obtain approximate asymptotic critical values for the ADF test statistics via simulation.

The rejection of the null hypothesis of no cointegration in Gregory and Hansen (1996) provides evidence in favor of the alternative specification that the variables are cointegrated with a structural break. Because the alternative hypothesis includes the standard model of cointegration with no structural break as a special case, a rejection of the null hypothesis by itself does not provide much evidence regarding whether a structural change has indeed occurred.[7] Further, since the break date in Gregory and Hansen is chosen to maximize the power of the test, it need not be a consistent estimate of the true break date, in contrast to the estimates of Bai and Perron (1998) and Kejriwal and Perron (2008).

Results of Unit Root and Cointegration Tests

This study analyzes the retail-wholesale-farm price relationship in U.S. beef and pork markets by employing monthly data from January 1970 through February 2008. Retail and wholesale beef and pork prices are obtained from the USDA/Economic Research Service (2008); farm cattle and hog prices are taken from the Livestock Marketing Information Center (2008). Retail beef and pork prices are an average value of selected cuts of grocery store meat, measured in cents per pound of retail weight.[8] Wholesale beef and pork prices are composite prices of wholesale cuts in a cents-per-pound retail equivalency. Farm cattle and hog prices are the weighted Nebraska direct price for 1,100- to 1,300-pound steers and the Iowa/ Minnesota hog carcass price, respectively.[9] All price variables are expressed in natural logarithms.

Visual inspection of the nominal price series presented in figure 1 suggests beef and cattle price series are nonstationary at all levels of the vertical chain, given the upward trend of the series.[10] On the other hand, wholesale pork and farm hog prices appear to be stationary with the exception of the initial four to seven years; the retail pork price series seems to be nonstationary. Perron (1997) cautions that seemingly nonstationary series may prove to be stationary once a break in the trend function is accounted for.

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[6] Since the variable in question (μ_t) is a residual series, Gregory and Hansen (1996) omit the trend term $\beta_1 t$ from the braced terms in equation (1) when checking for unit root.

[7] In other words, Gregory and Hansen's (1996) cointegration test is not a test for structural change, per se. Rather, the focus is to improve the power of conventional cointegration tests by allowing for a structural break.

[8] The retail beef value calculated by the USDA/Economic Research Service (2008) includes the average price of various roasts, ground beef, steaks, and other products, as reported by the Bureau of Labor Statistics. The retail pork value includes the average price of bacon, chops, ham, and other pork products.

[9] There exist "quasi"-national cattle and hog price data. For example, the five-market steer price series by the USDA/ Agricultural Marketing Service (AMS) is often used to represent a national cattle price, and AMS reported a 51%–52% lean hog price, based on the now-defunct six-market hog price series. Those quasi-national farm price series are not utilized in this study, as they are only available beginning in January 1989. Retail and wholesale price series begin in January 1970. [See Boetel and Liu (2008) for more details on the local farm price data used in the analysis.]

[10] As pointed out by the editor, the unit root test result may differ depending on whether the time series being tested is specified in nominal or real terms. For example, upon removing inflation as the common trend, the resulting real series may exhibit stationarity even though the nominal series contains a unit root. Likewise, as discussed in Wang and Tomek (2007), a casual inspection of nominal price graphs may suggest stationarity, but the test may conclude that the real price series contains a unit root upon deflating by an upward-trending price index. Regardless, the ADF tests used in the study address this dilemma by considering specifications both with and without a deterministic trend in the price series.

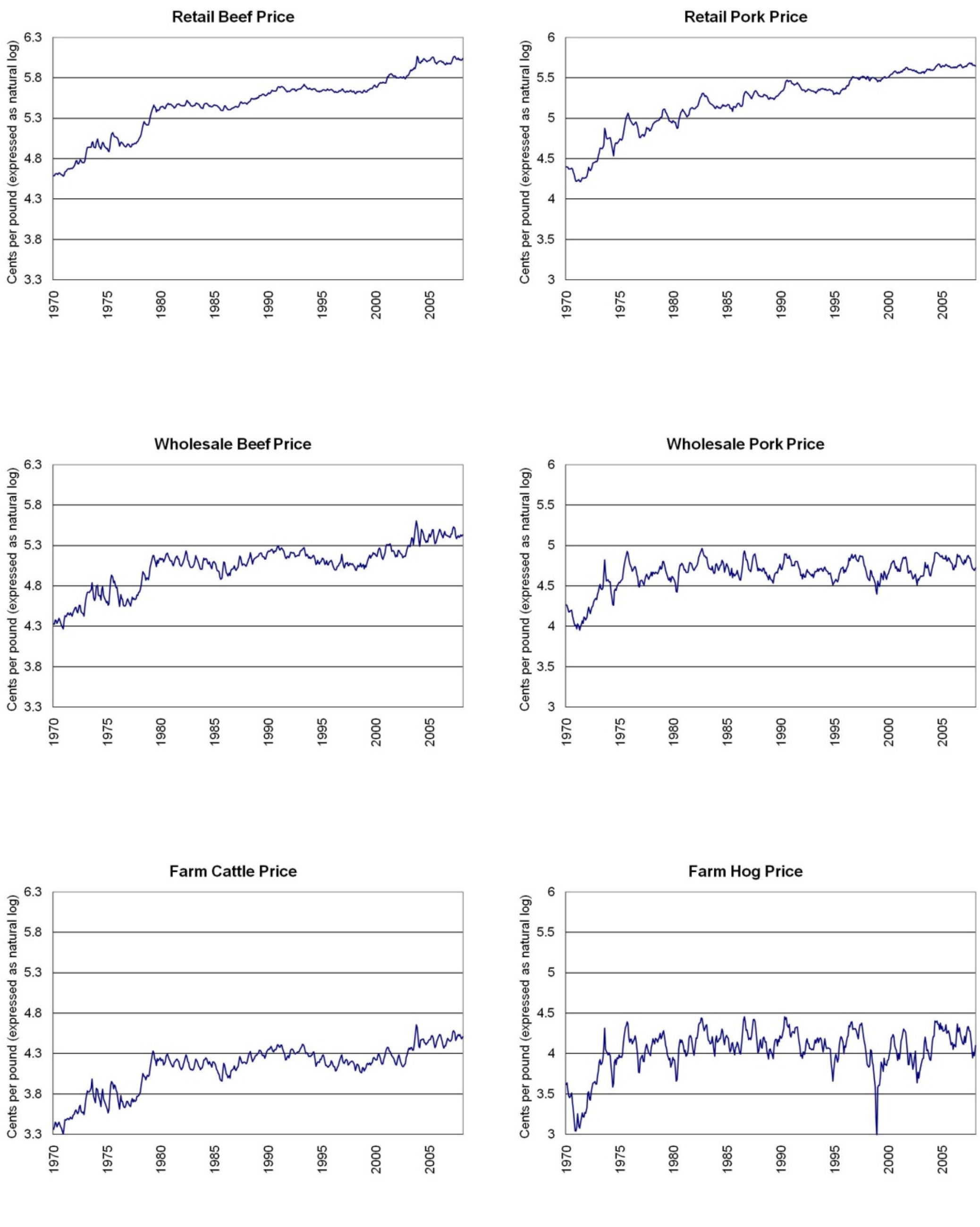

Figure 1. Nominal price series

Table 1. Preliminary Data Analysis

PANEL A. Unit Root Test[a] (*t*-statistics on the coefficient of the lagged dependent variable)

	ADF Test		Perron's Unit Root Test (allowing for a breaking trend)	
Price Series	With a Drift	With a Drift and a Trend	H_A: With Intercept Shift	H_A: With Intercept and Slope Shifts
Retail Beef	−1.19	−2.39	−4.00	−4.02
Wholesale Beef	−2.34	−3.00	−4.26	−4.11
Farm Cattle	−2.41	−2.93	−4.36	−4.27
Retail Pork	−1.88	−2.74	−5.33**	−6.09**
Wholesale Pork	−3.36**	−3.69**	−5.88***	−6.29***
Farm Hog	−3.46***	−3.56**	−5.73***	−6.12**

PANEL B. Cointegration Test[b] (*t*-statistic on the coefficient of the lagged residual)

		Gregory & Hansen Procedure	
Price Series	Engle & Granger Procedure	With Intercept Shift	With Intercept and Slope Shifts
Retail Beef on Constant, Wholesale Beef & Farm Cattle	−3.21**	−5.03**	−4.88

Notes: Double and triple asterisks (**,***) denote statistical significance at the 5% and 1% levels, respectively. Price variables are measured in logarithm form.

[a] The null hypothesis is that the price series contains a unit root.

[b] The null hypothesis is that the residual series in the cointegration equation contains a unit root.

Panel A of table 1 reports results from the conventional ADF and Perron's (1997) breaking trend unit root tests. Two specifications of the trend function are considered for the conventional ADF test. One includes only the drift term, while the other includes both drift and trend terms. Likewise, two specifications are considered for Perron's breaking trend models. The first allows for an intercept shift only, and the second for both intercept and slope shifts. The lag length for lagged first-difference terms of the price series in equation (1) is selected via Perron's general-to-specific *F*-sig criterion, beginning with a maximum lag of six months. We remove seasonality from each price series by conducting the unit root test on the residual series derived from regressing the price variable in question on 12 monthly dummy variables. Finally, in selecting the break date under Perron's method, a 15% trimming is used, dictating that the break date cannot occur within the first or last 15% of the sample observations.

As reported, the conventional ADF test clearly suggests retail and wholesale beef and farm cattle prices are nonstationary, consistent with intuition from visual inspection. Further, unit roots are found in the above beef/cattle price series even after allowing for a structural break in the trend function using Perron's procedure. For the pork/hog prices, the ADF test results indicate that retail pork price is nonstationary, while the wholesale pork and farm hog prices are stationary. However, the nonstationarity result pertaining to the retail pork price is overturned at the 95% confidence level when a breaking trend (at the level or at both the level and slope) is allowed, validating Perron's admonition. The stationarity conclusions pertaining to the wholesale pork and farm hog price series continue to hold under Perron's breaking trend specification. In sum, the results suggest that beef and cattle prices are nonstationary, while pork and hog prices are stationary if a break in the trend function is allowed.

Cointegration tests of the confirmed nonstationary beef/cattle prices are performed using Engle and Granger's (1987) conventional procedure and Gregory and Hansen's (1996) recent procedure, which allows for a structural break in the cointegration equation. A cointegration equation is estimated (with 15% trimming at both ends of the sample) by regressing the retail beef price on a constant, the wholesale beef price, and farm cattle price.[11] Panel B of table 1 demonstrates that the three beef/cattle price series are cointegrated at the 5% significance level under Engle and Granger's procedure and continue to be significant when we allow for a structural break in the intercept using the procedure of Gregory and Hansen. However, the power of Gregory and Hansen's test suffers when both the intercept and slope coefficients are subject to structural change; the null hypothesis of no cointegration cannot be rejected in this latter specification.

Results on Structural Breaks in Retail-Wholesale-Farm Price Relationships

Bai and Perron's (1998, 2003a) procedure is used to estimate structural breaks in the long-run relationship among the three stationary pork/hog price series, while Kejriwal and Perron's (2008, 2009) procedure is used for the cointegrated nonstationary beef/cattle prices. Following Goodwin and Holt (1999) and Goodwin and Harper (2000), the long-run price linkage equation expresses retail price as a function of wholesale and farm prices. This specific way of normalizing the long-run equation could be justified by arguing that farm markets are more competitive than wholesale and retail markets, and farm prices are therefore more or less determined in a competitive manner based on costs.[12]

To account for the time it takes the farm animal to reach the consumer, the retail beef price is specified as a function of contemporaneous wholesale beef price and one-period-lagged farm cattle price. The retail pork price is specified as a function of one-period-lagged wholesale pork price and two-period-lagged farm hog price using the same logic. The importance of allowing for lags in retail-wholesale-farm price transmission analysis was pointed out by Parham and Duewer (1980).[13] Both intercept and slope shifts are considered as dummy variables for the pork/hog price transmission equation, while only intercept shifts are allowed for the beef/cattle price linkage equation due to the finding that beef/cattle prices are non-stationary. Kejriwal and Perron (2008) note that if the parameters of nonstationary regressors are subject to structural breaks, the confidence intervals for the break date estimates are correlated, and consequently cumbersome to compute.[14] The estimation involves a hetero-skedasticity and autocorrelation consistent covariance matrix estimator. Finally, a maximum of

[11] Seasonality is not removed from the price variables (as done in the unit root test) because it is reasonable to assume that all three prices are subject to the same seasonal pattern.

[12] If farm prices are not determined exogenously based on costs, the issue of simultaneity in the estimation of the long-run static equation arises. The Phillips and Loretan (1991) procedure used in this study purges this contemporaneous endogeneity. Note that the order of normalization in the long-run equation does not change the specification of the ensuing short-run analysis, because the reduced-form equations express each price as a function of past observations of the three prices.

[13] The required lag length depends, in part, on the amount of processing and the time required to physically move the product. Parham and Duewer (1980) state that it takes about one week for a farm cattle price change to be transmitted to the carcass level and another two to three weeks to the retail level, while the time required for the pork/hog price transmission is one week from farm to wholesale level and another four weeks or longer to the retail level, depending on the amount of processing involved. Given that those lags are structural in nature, dictated by processing technology, the lags should persist in the long run.

[14] To gauge the extent of slope changes in the beef/cattle equation, slope dummies are considered in an alternative specification. This alternative specification results in similar break date estimates, although many of the coefficients of the slope dummy variables in the price transmission equation are insignificant.

Table 2. Structural Break Tests and Estimated Break Dates

PANEL A. Beef/Cattle Price Transmission Equation

Sup$F$(1 | 0)	**Sup$F$(2 | 0)**	**Sup$F$(3 | 0)**	**Sup$F$(4 | 0)**	**Sup$F$(5 | 0)**
810.87***	660.25***	578.13***	681.14***	608.23***
Sup$F$(1 | 0)	**Sup$F$(2 | 1)**	**Sup$F$(3 | 2)**	**Sup$F$(4 | 3)**	
810.87***	278.85***	183.34***	117.00***	

Break Dates	**95% Lower Bound**	**95% Upper Bound**
November 1975	September 1975	January 1976
July 1981	May 1981	September 1981
May 1993	April 1993	June 1993
April 2001	February 2001	June 2001

PANEL B. Pork/Hog Price Transmission Equation

Sup$F$(1 | 0)	**Sup$F$(2 | 0)**	**Sup$F$(3 | 0)**	**Sup$F$(4 | 0)**	**Sup$F$(5 | 0)**
84.57***	148.65***	356.17***	277.00***	223.16***
Sup$F$(1 | 0)	**Sup$F$(2 | 1)**	**Sup$F$(3 | 2)**		
84.57***	84.54***	40.22***		

Break Dates	**95% Lower Bound**	**95% Upper Bound**
October 1978	May 1978	March 1979
September 1987	May 1987	January 1988
October 1997	June 1997	February 1998

Notes: Triple asterisks (***) denote statistical significance at the 1% level. Sup$F(k\,|\,k-1)$ is the sup Wald test statistic for the null hypothesis of $k-1$ structural breaks versus the alternative hypothesis of k breaks.

five breaks for the study period are specified and a 15% trimming rate is adopted, dictating that the minimum length of each segment of the regression be no less than 15% of the total number of observations.[15]

Table 2 reports the test statistics of no break versus a prespecified number of breaks ranging from one to five [Sup$F(k\,|\,0)$, $k = 1, 2, \ldots, 5$] and the sequential tests of $k-1$ versus k breaks [Sup$F(k\,|\,k-1)$]. The Sup$F(k\,|\,0)$ statistics clearly reject the null hypothesis of no structural break for each of the two price relationships. Further, based on the sequential test statistics of Sup$F(k\,|\,k-1)$, we identify four breaks for the beef/cattle price transmission equation and three breaks for the pork/hog price linkage equation. As reported in table 2, the estimated break dates for the beef/cattle price transmission equation are November 1975, July 1981, May 1993, and April 2001, while the break date estimates for the pork/hog price relationship are October 1978, September 1987, and October 1997.

Our objective in estimating the break dates is to obtain unbiased price transmission elasticity estimates, accounting for the fact that structural breaks in the long-run price relationship have indeed occurred. While it is beyond the scope of the current model to provide explanations

[15] Bai and Perron (2003b) discuss size distortions associated with heteroskedasticity and autocorrelation consistent covariance matrix estimators when a low trimming rate is used. They suggest a rate of 20% for a sample size of 120 while acknowledging that this rate can be reduced somewhat if a larger sample size is involved. Regardless of the above theoretical consideration, the break date estimates obtained under the 15% trimming are found to be similar to those obtained with a lower trimming rate (e.g., 10%), although additional breaks often emerge in the latter specification. Some of the coefficients associated with these additional breaks in the price transmission equation, however, are not statistically different from zero. Importantly, the estimated wholesale and farm price transmission parameters are robust regardless of whether or not those additional breaks are included in the price transmission equation.

for why breaks occur at specific dates, the following conjectures may be insightful. Regarding the beef/cattle price relationship, the 1975 and 1981 breaks could be reflections of the 1973 and 1978 energy crises (after accounting for cattle production lags of 18 to 30 months), which might fundamentally shift the cost structure of producing and marketing beef products. The 1993 break may be a manifestation of a change in trade regime dictating the export of U.S. beef to Japan. In 1988, Japan signed an agreement with the United States to phase out the Japanese beef import quota, culminating in a complete liberalization of imports by April 1991. The 2001 break may capture the beginning of the decrease in the Atkins phenomenon popularity. It was estimated that media information supporting Atkins-type diets peaked in the early 2000s, and this information boosted beef demand from 1998 to 2003 by 2% (Tonsor, Mintert, and Schroeder, 2009).

The 1978 break in the pork/hog price relationship may reflect the energy price hike during that period, and the 1987 break may capture the effect of the national pork check-off program, which commenced in 1986. Finally, the 1997 break may reflect the dramatic increase in vertical contracting that occurred in the early 1990s.[16]

Estimated Long-Run Retail-Wholesale-Farm Price Relationship

The procedures used in estimating break dates give a consistent estimate of the price transmission parameter **b** in $y_t = a + \mathbf{b}\,\mathbf{x}_t + \varepsilon_t$ (Bai and Perron, 1998; Kejriwal and Perron, 2008). Because this parameter is of intrinsic interest, it is important to delve further into other properties of this regression estimate.

First, while super-consistent, Stock (1987) and Phillips and Loretan (1991) show that the OLS estimate of **b** is asymptotically biased when **x** and *y* are cointegrated due to contemporaneous correlations between the variables in the static equation. Second, Banerjee et al. (1993) find that the estimation of the above static long-run equation, which ignores the dynamics of the data-generating process, can result in persistent and substantial finite sample bias in the long-run coefficient as well. Third, the estimation bias of **b** will be carried over to the estimation of the short-term price dynamics if the residual in the long-run equation is utilized as an error correction term. To address this problem, we adopt the dynamic least squares procedure developed by Phillips and Loretan (1991), which includes *m* leading and *n* lagged terms of first differences of the regressors and *q* lagged terms of ε_t from the static price transmission equation:

$$(3) \qquad y_t = a + \mathbf{b}\,\mathbf{x}_t + \sum_{i=-m}^{n} \varsigma_i \Delta \mathbf{x}_{t-i} + \sum_{i=1}^{q} \xi_i \left(y_{t-i} - a - \mathbf{b}\,\mathbf{x}_{t-i} \right) + e_t \, .$$

Phillips and Loretan (1991) show that the inclusion of the leading and lagged terms purges the contemporaneous correlations between the variables; thus the estimate of **b** is asymptotically unbiased and normal. This suggests that the distributions of the parameter estimates can be obtained in the usual way.[17] While equation (3) can be estimated by nonlinear least squares, an alternative two-step procedure estimates the static components of the equation

[16] About 40% of hog sales were coordinated by contracts and integrated operations in 1998, compared with only 11% in 1993 and 3% in 1980 (Martinez, 1999).

[17] The inclusion of the leading terms of the first difference of the regressors is to account for feedback from the left-hand-side variable (y) to the right-hand-side variable (**x**), purging potential contemporaneous endogeneity. The inclusion of the lagged terms of the first difference of the regressors and the lagged terms of the residuals is to account for contemporaneous correlation between ε_t and the residuals that drives the nonstationary regressors (i.e., ν_t in $\mathbf{x}_t = \mathbf{x}_{t-1} + \nu_t$).

(i.e., $y_t = a + \mathbf{b}\,\mathbf{x}_t + \varepsilon_t$) and then estimates equation (3) with the lagged residual terms from the first step.

Retail price is again specified as a function of wholesale and farm prices, where the farm cattle price is lagged one period and wholesale pork and farm hog prices are lagged one period and two periods, respectively. Based on the break date estimation results in table 2, four structural dummy variables are included in the beef/cattle price transmission equation to allow for intercept shifts, and three structural dummy variables are specified for the pork/hog price equation to allow for intercept and slope shifts.[18] As to the leading and lagged terms, preliminary estimations suggest that including only one lagged and one leading term of the first difference of **x** (i.e., the wholesale and farm prices) and one lagged term of the static residuals (i.e., ε_t) in the price transmission equations is sufficient for capturing the dynamics of the data-generating process.[19] The equations are estimated using the two-step procedure discussed previously.

Table 3 presents estimation results. The intercept for the base period of regime 1 (spanning from the first observation of the sample to the first estimated break date) is 1.25 and 0.55 for the beef/cattle and pork/hog equations, respectively. The regime coefficients for the intercept other than the base regime are calculated as the base regime intercept plus the estimated parameter for the respective intercept dummy variable, and are reported in the columns titled "Intercept/Price Transmission Elasticity." The importance of allowing for intercept shifts in the estimation of price transmission equations is corroborated by the highly significant *t*-statistics associated with the estimated parameters for the intercept dummy variables. Pairwise tests (not reported in table 3), which check for equality of the intercept terms in two different regimes, confirm that intercepts vary from regime to regime.

With respect to the slope coefficients, the estimated wholesale and farm price transmission elasticities for the beef/cattle equation are 0.47 and 0.40, respectively. The regime-dependent price transmission elasticities for the pork/hog equation are calculated as the base regime slope coefficient plus the coefficient for the respective slope dummy variable. The *t*-statistics indicate that the wholesale pork and farm hog price transmission elasticities for the various regimes are different from those for the base regime.[20] With the exception of the farm price coefficient in the third and fourth regimes, pairwise tests (not reported) reject the null hypotheses that the slope terms in two different regimes are equal, providing further evidence of the importance of allowing for slope shift in the pork/hog equation.

Depending on regimes, the estimated price transmission elasticities range between 0.38 and 0.79 for the wholesale pork price, and between 0.02 and 0.26 for the farm hog price. When comparing the slope coefficients across regimes, we find that the wholesale pork price transmission elasticity of 0.38 for the most recent decade (i.e., regime 4) is notably smaller than those of the earlier years. Similarly, the farm hog price transmission elasticity for the earlier regimes is much larger than those pertaining to the latter two regimes (which were found to be not statistically different from each other in the pairwise test). These results

[18] The structural dummy variable for each break date takes the value of 1 for the observations between the break date in question and the subsequent break date (or the end date of the sample for the last structural dummy variable), and 0 otherwise.

[19] Including additional lead/lagged terms does not change the regression coefficients and key statistics (e.g., adjusted R^2 and Durbin-Watson) in a noticeable manner. Further, the coefficients for additional lead/lag terms are not statistically different from zero in most cases.

[20] The importance of including structural dummy variables (for intercept and slope terms) in the specification of the pork/hog price transmission equation is also manifested by a much lower coefficient of determination and Durbin-Watson statistic when those dummy variables are excluded.

Table 3. Estimation Results of Long-Run Price Transmission Equations (in log) [dependent variables = *Retail Beef Price* and *Retail Pork Price*]

	Retail Beef Price			***Retail Pork Price***		
Description	Estimated Parameter	*t*-Statistic	Intercept/ Price Transmission Elasticity	Estimated Parameter	*t*-Statistic	Intercept/ Price Transmission Elasticity
Intercept:						
Constant (R1) [a]	1.25	33.07	1.25	0.55	6.19	0.55
Dum for R2	0.11	35.32	1.36	1.16	8.28	1.71
Dum for R3	0.21	56.55	1.46	1.67	11.39	2.22
Dum for R4	0.35	91.95	1.60	3.05	19.29	3.59
Dum for R5	0.45	85.27	1.70			
Wholesale Price:						
WP (R1) [b]	0.47	17.80	0.47	0.79	16.90	0.79
WP $*$ Dum for R2				−0.30	−5.14	0.50
WP $*$ Dum for R3				−0.14	−2.34	0.65
WP $*$ Dum for R4				−0.41	−7.20	0.38
Farm Price:						
FP (R1) [b]	0.40	16.33	0.40	0.15	4.55	0.15
FP $*$ Dum for R2				0.11	2.69	0.26
FP $*$ Dum for R3				−0.13	−3.16	0.02
FP $*$ Dum for R4				−0.11	−3.16	0.05
Phillips & Loretan (1, 1, 1) Terms: [c]						
$\Delta WP_{contemp}$	−0.40	−10.83		0.25	9.96	
ΔWP_{lead}	0.08	2.21		−0.46	−9.65	
ΔWP_{lag}	0.05	1.24		0.03	1.08	
$\Delta FP_{contemp}$	0.18	5.12		−0.06	−4.16	
ΔFP_{lead}	−0.30	−8.28		−0.02	−1.16	
ΔFP_{lag}	−0.07	−1.88		0.08	3.02	
Lagged residuals	0.87	41.49		0.86	35.61	
Adjusted R^2	0.998			0.996		
Durbin-Watson Statistic	1.81			2.01		

[a] R1 denotes Regime 1; R2, R3, R4, and R5 are defined similarly. Dum stands for structural change dummy variable. The break dates defining the four dummy variables in the beef/cattle equation are 11/75, 7/81, 5/93, and 4/01. The break dates defining the three dummy variables in the pork/hog equation are 10/78, 9/87, and 10/97.

[b] Farm price is lagged one month in the beef/cattle equation and two months in the pork/hog equation, while wholesale price is contemporaneous in the beef/cattle equation and lagged one month in the pork/hog equation (see text footnote 13).

[c] The dynamic specification of PL (1, 1, 1) entails contemporaneous, one-period leading and lagged terms of the first difference of the wholesale and farm prices (allowing for lags as discussed in footnote b above), and one-period-lag residuals from the static price transmission equation estimated without the Phillips and Loretan terms.

suggest the relationship between pork wholesale and retail prices, as well as the relationship between hog farm and pork retail prices, has weakened over the study period.[21]

The lower portion of table 3 reports the coefficients pertaining to Phillips and Loretan's (1991) dynamic specification. With the exceptions of the lagged first difference of wholesale beef and pork prices and the leading first difference of the farm cattle price, all the terms are significant. Lagged residuals are highly significant in both equations, suggesting the importance of error correction in the maintenance of the long-run price relationship. Excluding the Phillips and Loretan terms lowers the Durbin-Watson statistics to 0.35 and 0.42 for the beef/cattle and pork/hog equations, respectively, highlighting the relevance of considering the dynamics of the data-generating process when estimating price transmission equations.

Short-Run Price Dynamics

Error correction vector autoregressive models provide insight into how retail, wholesale, and farm prices adjust in the short run (facing shocks) and then return to the long-run price relationship. We use residuals from previously estimated long-run price equations as the error correction terms in the price dynamic system, and then use the short-run price dynamic system to gain insights into the causal relationships among the retail, wholesale, and farm prices in question. In the case of beef/cattle prices, which we found to be cointegrated non-stationary variables, the approach can be justified by invoking Granger's representation theorem (Engle and Granger, 1987), stipulating that in a cointegrated system there exists an error correction mechanism whereby deviations from the long-run equilibrium can be reflected in the short-run dynamics to ensure the upkeep of the long-run condition. In the case of pork/hog prices, we justify the use of an error correction model by acknowledging the model is a special case of Hendry's autoregressive distributed lag (ADL) model, which is appropriate for stationary time series (Hendry, 1995).[22]

We estimate two error correction systems (beef/cattle and pork/hog), with each system consisting of three equations (retail, wholesale, and farm prices). In each of the three equations, the contemporaneous price change is expressed as a function of past changes (up to a lag truncation) of all three prices and a one-period-lag error correction term reflecting deviations from the long-run equilibrium. The lag truncation is specified as 13, reflecting both the monthly and annual effects of the data-generating process on price formation. To investigate the asymmetrical aspects of error adjustment, the error correction term is decomposed into positive and negative components (i.e., treating 0 as the threshold) and tests are conducted on the equality of the associated coefficients. Finally, the short-run price dynamic systems are estimated by OLS.[23]

Table 4 reports several key results from the estimation of the short-run price dynamics. First, the hypothesis that adjustment speeds for positive and negative deviations from the long-run equilibrium price relationship are symmetrical is rejected at conventional confidence levels for all three equations in the beef/cattle system. The asymmetrical adjustment speed in

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[21] Adachi and Liu (2009) also find a weakening relationship between retail and farm prices in the Japanese pork/hog market.

[22] In ADL(1, 1), the left-hand-side variable is expressed as a function of its first-period lag, as well as contemporaneous and one-period lagged terms of the regressors. Even with this very simple model, with one autoregressive term of the dependent variable and one distributed lagged term of the regressors, Hendry (1995) shows that it encompasses nine commonly used models as special cases, including univariate time series, distributed lags, partial adjustment, common factor, and error correction.

[23] Stock (1987) shows that the OLS estimator of an error correction system converges to limiting normal random variables at the usual rate of $T^{1/2}$, and because of the fast rate of convergence of the estimated parameters in the long-run price transmission equation, the short-run estimates are asymptotically independent of the long-run estimates.

Table 4. Estimation Results of Short-Run Price Dynamics

	Beef/Cattle System			Pork/Hog System		
	ΔRP	ΔWP	ΔFP	ΔRP	ΔWP	ΔFP
***p*-Values for Hypothesis that the Adjustment Speed is Symmetrical:**						
	0.03	0.00	0.00	0.20	0.31	0.92
Estimated Adjustment Speed:						
Impose Symmetry				−0.06	0.05	0.12
				(−1.19)	(0.47)	(0.64)
Allow Asymmetry						
Positive Deviation	0.40	1.94	2.29			
	(2.45)	(4.64)	(5.41)			
Negative Deviation	0.03	0.03	0.12			
	(0.38)	(0.16)	(0.69)			
***p*-Values for Hypothesis that the Coefficients Are Jointly Zero:**						
Lag Terms of ΔRP	0.00	0.00	0.00	0.02	0.09	0.01
Lag Terms of ΔWP	0.00	0.03	0.05	0.00	0.35	0.01
Lag Terms of ΔFP	0.00	0.00	0.00	0.76	0.01	0.00
Durbin-Watson Statistic	1.96	1.98	1.97	2.01	2.01	1.99
Adjusted R^2	0.47	0.26	0.26	0.38	0.15	0.19

Notes: The notations ΔRP, ΔWP, and ΔFP represent the first-difference terms of retail, wholesale, and farm prices, respectively. In each equation, the left-hand-side variable is the first difference of the price in question, while the right-hand-side variables are lagged (up to the 13th lag) first differences of the retail, wholesale, and farm prices, and the one-period lagged residuals from the long-run price linkage equation. Values in parentheses are *t*-statistics.

the retail beef equation is consistent with the notion that menu costs are asymmetrical depending on whether retailers raise or lower prices (Meyer and von Cramon-Taubadel, 2004). The asymmetrical results pertaining to the wholesale beef and farm cattle price equations are in congruence with the finding of Bailey and Brorsen (1989) that adjustments to price increases and decreases occur at different speeds in spatially related fed cattle markets in the United States. Contrary to the beef/cattle results, symmetry in adjustment speed is not rejected in any of the three pork/hog equations, consistent with a later result that the coefficients of the error correction terms in those three equations are not significantly different from zero.

Second, the estimated adjustment speeds are reported in table 4, with symmetry imposed in the pork/hog equations. Regarding the beef/cattle system, the coefficients for the positive deviations are all positive and statistically different from zero, while the coefficients for the negative deviations are all insignificant. These findings reveal that short-run adjustments in the retail, wholesale, and farm price occur only when retail beef price is too low from the perspective of the long-run equilibrium vertical price relationship. As to the adjustment speeds in the three pork/hog equations, all the coefficients are found to be insignificant, even with symmetry imposed.

Third, the hypothesis that the coefficients of the lagged terms of a price change are jointly zero is rejected at the 90% or higher confidence level in all equations (with the exceptions of lagged farm hog price changes in the retail pork price equation and lagged wholesale pork price changes in the wholesale pork price equation), suggesting a vigorous short-term dynamic relationship among the retail, wholesale, and farm prices. To gain further insight into the effect on short-run price formation of demand and supply shocks in markets across the vertical chain, impulse responses (60 steps) of the price system to a one-time shock in an innovation are simulated, with the size of the shock being the standard error of the estimate of the equation in question. Results from the impulse response confirm the existence of bi-directional feedback within the beef/cattle and the pork/hog systems; significant price responses are detected in all three markets regardless of the market in which the shock occurs.[24]

The finding that retail price shocks significantly affect wholesale and farm prices differs from the conclusions reported by Goodwin and Holt (1999) and Goodwin and Harper (2000) that retail shocks are mainly contained in the retail market. Similarly, the finding that wholesale price shocks significantly affect farm price is not in agreement with the contention of the above two studies that wholesale shocks affect only wholesale and retail prices. The difference in the short-run price transmission direction may be due to the fact that the current study incorporates Phillips and Loretan's (1991) dynamic lead/lag specifications, which explicitly account for feedbacks of various directions. Additionally, the discrepancy may be explained by the monthly data used in the current study, which enable the model to capture the type of feedback that takes more than several weeks to materialize.

Summary and Conclusion

The conventional econometric method of specifying, a priori, a definitive structural break date (e.g., Chow tests) to account for change is not always feasible given the subtlety of transformation timing. Further, even when convincing evidence exists to show that a structural change has occurred at a specific time, it is important to remain mindful of data mining issues, because the evidence has most likely been obtained by a series of formal and informal pretests of the data by the analysts or their predecessors. This argument has led to recent econometric developments of identifying structural change of unknown timing, while constructing distribution theory to account explicitly for the search of the break. Subscribing to the above modeling philosophy, this paper examines structural breaks in the vertical price relationships in U.S. beef/cattle and pork/hog sectors with the break dates being estimated endogenously.

Specifically, consistent estimates of break dates for the price linkage equation are obtained by the procedures developed by Bai and Perron (1998, 2003a) and Kejriwal and Perron (2008, 2009), depending on whether the variables in the price equation are stationary or nonstationary but cointegrated. With the identified break dates, Phillips and Loretan's (1991) dynamic least squares framework is adopted for the estimation of the long-run price linkage equation to ensure the unbiasedness of the estimated price transmission parameters. The estimated residual series in the long-run price linkage equation is then utilized as the error correction term in the estimation of the short-run price dynamic system. Finally, simulations are conducted to gain insights about the response of price variables to a shock in the innovation of the system.

Results from the break date analysis suggest there are four breaks in the beef/cattle price linkage equation, possibly reflecting the energy crises of the late 1970s and early 1980s, changes in trade policy in Japan in the early 1990s, and the phenomenon of the Atkins diet in the early 2000s. Three structural breaks for the pork/hog price linkage equation are identified,

[24] Detailed results on the impulse response analysis are available from the authors upon request.

explainable by the energy crisis of the late 1970s, the beginning of the national pork check-off program in the mid-1980s, and the dramatic increase in vertical contracting since the early 1990s.

In studying price linkages, retail price is specified as a logarithmic function of the wholesale and farm prices. With respect to the retail beef price equation, the long-run price transmission elasticity is about 0.47 and 0.40 for the wholesale beef and farm cattle prices, respectively. In the retail pork price equation, the long-run price transmission elasticity ranges between 0.38 and 0.79 for the wholesale pork price and 0.02 and 0.26 for the farm hog price, depending on the regimes defined by the identified structural break dates. Results from the analysis of short-run price dynamics confirm that adjustment speeds are asymmetrical in the beef/cattle system, but statistically insignificant in the pork/hog system. While previous studies argue that shocks are transmitted mainly in the direction of farm to wholesale to retail, the impulse response of the current analysis suggests a more versatile feedback mechanism that is bi-directional in nature.

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References

Abdulai, A. "Using Threshold Cointegration to Estimate Asymmetric Price Transmission in the Swiss Pork Market." *Appl. Econ.* 34,6(2002):679–687.

Adachi, K., and D. J. Liu. "Estimating Long-Run Price Relationship with Structural Change of Unknown Timing: An Application to the Japanese Pork Market." *Amer. J. Agr. Econ.* 91,4(2009):1440–1447.

Ardeni, P. G. "Does the Law of One Price Really Hold for Commodity Prices?" *Amer. J. Agr. Econ.* 71,3(1989):661–669.

Bai, J. "Estimating Multiple Breaks One at a Time." *Econometric Theory* 13,3(1997):315–352.

Bai, J., and P. Perron. "Estimating and Testing Linear Models with Multiple Structural Changes." *Econometrica* 66,1(1998):47–78.

———. "Computation and Analysis of Multiple Structural Change Models." *J. Appl. Econometrics* 18,1(2003a):1–22.

———. "Critical Values for Multiple Structural Change Tests." *Econometrics J.* 6(2003b):72–78.

Bailey, D., and B. W. Brorsen. "Price Asymmetry in Spatial Fed Cattle Markets." *West. J. Agr. Econ.* 14,2(1989):246–252.

Balagtas, J. V., and M. T. Holt. "The Commodity Terms of Trade, Unit Roots, and Nonlinear Alternatives: A Smooth Transition Approach." *Amer. J. Agr. Econ.* 91,1(2009):87–105.

Banerjee, A., J. J. Dolado, J. W. Galbraith, and D. F. Hendry. *Cointegration, Error-Correction and the Econometric Analysis of Nonstationary Data.* London: Oxford University Press, 1993.

Boetel, B. L., and D. J. Liu. "Incorporating Structural Changes in Agricultural and Food Price Analysis: An Application to the U.S. Beef and Pork Sectors." Work. Pap. No. 2008-02, Food Industry Center, University of Minnesota, 2008.

Boyd, M. S., and B. W. Brorsen. "Price Asymmetry in the U.S. Pork Marketing Channel." *N. Cent. J. Agr. Econ.* 10,1(1988):103–110.

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

Gardner, B. "The Farm-Retail Price Spread in a Competitive Food Industry." *Amer. J. Agr. Econ.* 57,3(1975): 399–409.

Goodwin, B. K. "Forecasting Cattle Prices in the Presence of Structural Change." *S. J. Agr. Econ.* 24,2(1992): 11–22.

Goodwin, B. K., and D. C. Harper. "Price Transmission, Threshold Behavior, and Asymmetric Adjustment in the U.S. Pork Sector." *J. Agr. and Appl. Econ.* 32,3(2000):543–553.

Goodwin, B. K., and M. T. Holt. "Price Transmission and Asymmetric Adjustment in the U.S. Beef Sector." *Amer. J. Agr. Econ.* 81,3(1999):630–637.

Goodwin, B. K., and N. E. Piggott. "Spatial Market Integration in the Presence of Threshold Effects." *Amer. J. Agr. Econ.* 83,2(2001):302–317.

Goodwin, B. K., and T. C. Schroeder. "Cointegration Tests and Spatial Price Linkages in Regional Cattle Markets." *Amer. J. Agr. Econ.* 73,2(1991):452–464.

Gregory, A. W., and B. E. Hansen. "Residual-Based Tests for Cointegration in Models with Regime Shifts." *J. Econometrics* 70,1(1996):99–126.

Hahn, W. F. "Price Transmission Asymmetry in Pork and Beef Markets." *J. Agr. Econ. Res.* 42,4(1990):21–30.

Hansen, B. E. "The New Econometrics of Structural Change: Dating Breaks in U.S. Labor Productivity." *J. Econ. Perspec.* 15,4(2001):117–128.

Hendry, D. F. "Interpreting Linear Models." In *Dynamic Econometrics: Advanced Texts in Econometrics*, pp. 195–229. London: Oxford University Press, 1995.

Holt, M. T., and L. A. Craig. "Nonlinear Dynamics and Structural Change in the U.S. Hog-Corn Cycle: A Time-Varying STAR Approach." *Amer. J. Agr. Econ.* 88,1(2006):215–233.

Kejriwal, M., and P. Perron. "The Limit Distribution of the Estimates in Cointegrated Regression Models with Multiple Structural Changes." *J. Econometrics* 146,1(2008):59–73.

———. "Testing for Multiple Structural Changes in Cointegrated Regression Models." *J. Bus and Econ. Statis.* (2009). Online at doi: 10.1198/jbes.2009.07220.

Kinnucan, H. W., and O. D. Forker. "Asymmetry in Farm-Retail Price Transmission for Major Dairy Products." *Amer. J. Agr. Econ.* 69,2(1987):285–292.

Livestock Marketing Information Center. Farm cattle and hog price data, 2008. Online. Available at http://www.lmic.org. [Accessed March 2008.]

Martinez, S. W. "Vertical Coordination in the Pork and Broiler Industries: Implications for Pork and Chicken Products." Agr. Econ. Rep. No. 777, USDA/ Economic Research Service, Washington, DC, April 1999.

Meyer, J., and S. von Cramon-Taubadel. "Asymmetric Price Transmission: A Survey." *J. Agr. Econ.* 55,3(2004):581–611.

Parham, K. D., and L. A. Duewer. "An Evaluation of Price Spread and Cost Component Estimates for Choice Beef and Pork." Economics and Statistics Service Staff Rep. No. AGES-S801215, USDA/National Economics Division, Washington, DC, December 1980.

Perron, P. "Further Evidence on Breaking Trend Functions in Macroeconomic Variables." *J. Econometrics* 80,2(1997):335–385.

———. "Computation and Hypothesis Testing in Models with Multiple Structural Changes." 2004. Online. Available at http://people.bu.edu/perron/code.html. [Accessed February 2010.]

Phillips, P. C. B., and M. Loretan. "Estimating Long-Run Economic Equilibria." *Rev. Econ. Stud.* [Special Issue: *The Econometrics of Financial Markets*] 58,3(1991):407–436.

Stock, J. H. "Asymptotic Properties of Least Squares Estimators of Cointegrating Vectors." *Econometrica* 55,5(1987):1035–1056.

Tiffin, R., and P. J. Dawson. "Structural Breaks, Cointegration, and the Farm-Retail Price Spread for Lamb." *Appl. Econ.* 32,10(2000):1281–1286.

Tonsor, G., J. Mintert, and T. Schroeder. "U.S. Beef Demand Drivers and Enhancement Opportunities." Pub. No. MF-2893, Kansas State University Agr. Exp. Sta. and Coop. Ext. Ser., June 2009.

U.S. Department of Agriculture, Economic Research Service. "Meat Price Spreads." USDA/ERS, Washington, DC, 2008. Online. Available at www.ers.usda.gov/data/meatpricespreads. [Accessed March 2008.]

Wang, D., and W. G. Tomek. "Commodity Prices and Unit Root Tests." *Amer. J. Agr. Econ.* 89,4(2007): 873–889.