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Incorporating Structural Changes in Agricultural and Food Price Analysis: An Application to the U.S. Beef and Pork Sectors

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Dedicated to Neva on the date of her birth, **October 3,** 2008 in Hudson, Wisconsin.

Dear Talia:

Your upcoming birth was the motivation needed to complete this paper on a timely basis. Your occasional gentle nudge in my stomach has proven to be quite effective. This paper is about change and many questions still remain unanswered but one thing that is certain is that you will change my life, enriching it beyond what I can imagine. I wish you a wonderful, happy and productive life.

Love,

Your Mom (Brenda)

Executive Summary

While change is the only constant in life, one may not always be aware or certain of the timing of the transformation due to its subtlety and/or due to complications associated with concomitant changes. For this reason the conventional method of specifying, ***a priori,*** a definitive structural break date to account for changes in the estimation of an economic relationship is not always feasible. Further, even if one is convinced through examination of the data that a structural change has occurred at a specific time, one still has to be mindful of the issues of data mining and the subsequent undesirable implications on the legitimacy of statistical inferences on the estimates. This argument has led to recent methodological developments of identifying structural break dates with unknown timing while constructing statistical inference procedures to explicitly account for the fact that a search for the break date has been conducted.

Subscribing to the philosophy that it is important to account for structural change in conducting agricultural and food price analysis, this paper proposes a protocol that explicitly entertains structural changes when analyzing long-term price relationship and short-term price dynamics. The proposed protocol also pays attention to the time series properties of the price variables, such as nonstationarity (e.g., the price series do not have a constant mean and variance over time), and addresses the issue of a cointegration relationship among nonstationary price variables and possible structural change in that relationship. The concept of cointegration is important for understanding short-term price movements because it ascertains whether a group of nonstationary price variables maintain an equilibrium relationship over time, even though individually the price series do not have constant means and variances.

The proposed protocol is applied to the U.S. pork/hog and beef/cattle sectors to investigate how the retail, wholesale, and farm prices interact in the market place, both in the short-run and in the long-run, and how those price relationships have changed over a period of 38 years. Attention is also given to the asymmetry aspects of price transmission, in that the magnitude of the pass-through of a change in one price to another price depends on whether the price change is positive or negative. Additionally, the analysis investigates whether the speed of short-run price adjustments are the same when faced with a positive or negative deviation from the long-run price relationship.

The empirical results strongly suggest that retail beef, wholesale beef, and farm cattle prices do not maintain constant means and variances over the study period (i.e., the prices are nonstationary), while there is ample evidence to support the conclusion of constant means and variances for retail pork, wholesale pork and farm hog prices. Even though retail beef, wholesale beef, and farm cattle prices are found to be nonstationary, the results suggest that there exists an equilibrium relationship among the three prices. Similarly, the results suggest that there is an equilibrium relationship among retail pork, wholesale pork and farm hog prices. The equilibrium price relationship tying together the pork/hog prices and the relationship tying together the beef/cattle prices, however, are found to have experienced structural breaks over the study period. Specifically, three structural breaks are identified for the pork/hog price relationship, reflecting the energy crisis of the late 1970's, the beginning of vertical contracting in the mid 1980's, and the period of extremely low hog prices in the late 1990's. As to the

beef/cattle price relationship, two break dates are found, which capture the energy crisis of the late 1970's and severe weather patterns in the U.S. in the early to mid 1990's.

Vis-à-vis retail pork price, the long-run price pass-through elasticities are found to be about 0.80 and 0.20 for the wholesale pork and farm hog prices, respectively, implying the percentage mark-up in the retail pork price to be about 23% of the wholesale pork price and 123% of the farm hog price. Likewise, given the estimated long-run price pass-through elasticities of 0.85 and 0.15 for the wholesale beef and farm cattle prices, the mark-up contained in the retail beef price is calculated as 36% and 193% of the wholesale beef and farm cattle prices, respectively.

Asymmetry in pass-through is confirmed in two cases. The effect of wholesale pork price change on the retail pork price movement depends on whether the wholesale pork price is increasing or decreasing, while the pass-through rate of the farm cattle price change on retail beef price movement depends on whether the former is increasing or decreasing. For practical purposes, however, it is important to note that the difference in the magnitude of the pass-through rates are trivial, statistically significant notwithstanding.

The short-run adjustment speed toward a long-run equilibrium price relationship is found to be asymmetrical for retail pork prices, retail beef prices, wholesale beef prices, and farm cattle prices, depending on whether the price shock is positive or negative. On the other hand, the adjustment speed for wholesale pork and farm hog prices are found to be symmetrical, regardless of whether the short-run deviation from the long-run price relationship is positive or negative. The asymmetrical adjustment speed for retail pork and beef prices is consistent with the notion that menu costs are asymmetrical depending on whether retailers raise or lower prices, whereas the symmetrical results pertaining to the wholesale pork and farm hog prices may be attributable to the high level of integration between the pork processing and hog production sub-sectors. The finding of asymmetry in speed suggests that price adjustments facing a shock follow different paths depending on whether the shock is positive or negative. Finally, while previous studies utilizing weekly data find that shocks are transmitted mainly in the direction of farm to wholesale to retail, the analysis of short-run price dynamics of the current study, using monthly data, suggest a more versatile feedback mechanism that is multi-directional in nature.

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1. Introduction

The idea that prices convey sufficient information regarding efficient resource allocation is perhaps one of the most powerful insights that economists have offered. Of particular interest to agricultural and food economists is the long-run price relationship of major agricultural and food items across various dimensions such as time and space. Also of interest are the issues pertaining to the short-run adjustment toward the equilibrium price relationship. The importance of having a good understanding of food price relationship, both long-run and short-run, cannot be overstressed. First, by understanding how prices behave, one gains insights into the robustness of the underlying market. Second, the manner prices adjust to shocks in the short-run and the ability prices move toward the long-run equilibrium have clear welfare implications for consumers, producers and other market participants. Third, an enhanced understanding of food price evolution facilitates the arrival of sound food policy which, in some cases, is tantamount to saving lives and alleviating malnutrition. The recent food price hikes, especially in some developing countries, illustrate the importance of basing production, marketing, and food policy decisions on sound analysis of price behavior and relationship.

A significant portion of earlier agricultural and food price analysis results is aptly discussed and summarized in Tomek and Robinson (1990), wherein the focus lies on price variations over time, space, product forms and quality, as well as on the effect of pricing institutions and government interventions on the evolution of prices. Issues that are less discussed in Tomek and Robinson are the time series properties of the price series, such as unit root, cointegration, and error correction adjustment, and the implications on the econometric procedures appropriate for the analysis. In this regard, there are three important articles that fundamentally affect the way our profession has conducted price analysis during the recent two decades. Granger and Newbold (1974) point out the pitfalls of spurious regression when variables involved are nonstationary. Nelson and Plosser (1982) elicit a great deal of interest in unit root tests after demonstrating that most of the major macroeconomic time series are integrated of order one. Engle and Granger (1987) show that if nonstationary variables are cointegrated the analysis is immune from the problem of spurious regression, and provide a two step residual based procedure to test for cointegration under which there exists an error correction representation characterizing the short-run dynamics of the price series. Taken together, these three articles can be viewed as what might be called the unit root revolution, under which price analysts follow a common protocol of testing for unit root (e.g., augmented Dickey-Fuller, ADF, test, 1979), estimating and testing for long-run cointegration relationship (e.g., Engle and Granger's test, or Johansen's (1988, 1991) maximum likelihood method), and then estimating and analyzing the associated short-run price dynamics allowing for error corrections.

The unit root protocol, as described above, was quickly adapted by agricultural economists at its inception (e.g., Ardeni, 1989; Goodwin and Schroeder, 1991; Liu, Chung, Meyers, 1993).

Further, the procedure has been extended to allow for threshold effects in error correction, accounting for such factors as transaction costs and asymmetry in adjustment speed toward a long-run equilibrium relationship (e.g., Goodwin and Holt, 1999; Goodwin and Harper, 2000; Goodwin and Piggott, 2001; Abdulai, 2002). Less heeded by the agricultural economics profession has been the structural change counter revolution, which is discussed below.

While the unit root revolution was still gaining momentum, analysts began to have concerns about potential confusion between structural change and nonstationarity in the economic time series being studied. Extending the conventional ADF test to allow for exogenously specified structural dummy variables in both the null and alternative hypotheses, Perron (1989) is able to overturn many of Nelson and Plosser's unit root conclusions. Agonizing over the issue of data mining, subsequent authors, including Perron (1997), have developed procedures and distribution theories to endogenously specify structural break dates while testing for unit root.[1] Lest analysts should confuse a structural break with a lack of long-run relationship, Gregory and Hansen (1996) extend Engle and Granger's procedure to allow for structural change of unknown timing in the long-run equation when testing for cointegration.

While the above mentioned structural change counter revolution has motivated a plethora of empirical research in general economics (see Hansen, 2001, for a review), studies in this area that are related to agricultural and food products are not as abundant. Wang and Tomek (2007) question the frequent findings of nonstationarity in agricultural prices, and conduct unit root tests for several U.S. agricultural commodities, including corn, soybeans, barrows and gilts, and milk, allowing for an exogenously specified break date. The authors find that a shift in the intercept is an important consideration in unit root tests, and caution analyst to have a healthy skepticism toward the previous results based on conventional tests. Regarding cointegration tests, Tiffin and Dawson (2000) adapt Gregory and Hansen's procedure and find that there is a structural break in the early 1990's in the farm-retail price spread for lamb in the United Kingdom. Most recently, Frank and Garcia (2008) employ a nonparametric method to endogenously identify structural breaks in unit roots tests, and find that accounting for a structural break in the 1970's affects the conclusion pertaining to the presence of a time varying risk premium in agricultural futures markets.[2]

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[1] The conventional approach of specifying a structural break exogenously (e.g., Chow test) is criticized as data mining because usually either the current or previous analysts have examined the data and developed a pre-conceived notion of the timing of the structural break. The endogenous structural change identification procedures of recent literature search for break dates among possible candidates, but explicitly account for this search in the development of distribution theory and critical values.

[2] Much earlier, Dorfman (1993) utilized a nonparametric Bayesian framework to test for unit root in corn and soybean futures prices. The author examines the issue of structural change in market efficiency by, *a priori*, splitting the sample into subsamples, only to find no systematic difference in the test results.

The purpose of this paper is two-fold. First, given the limited attention the agricultural economics profession has paid to the recent time series econometric development in structural change of unknown timing, we will modify the conventional unit root protocol to include various endogenous break date identification, estimation, and testing procedures when conducting unit root tests, cointegration tests, and short-run price dynamic analysis. Second, we will apply this modified protocol, which may be called the "structural change protocol", to the analysis of price relationship in U.S. beef/cattle and pork/hog sectors using monthly data. The empirical analysis provides additional insights beyond Goodwin and Holt and Goodwin and Harper, which employ weekly data and (other than splitting the sample into earlier and later periods in their impulse response analysis of short-term price dynamics) do not entertain structural change.

The paper is organized in the following manner. Section two provides a brief discussion of the recent structural change methodologies used in the empirical sections, including Perron's unit root test, Gregory & Hansen's cointegration test, and the break point identification procedures of Bai and Perron (1998, 2003a) and Kejriwal and Perron (2008a, 2008b). Note that although Perron's and Gregory and Hansen's procedures allow for a structural break while testing for unit root and cointegration, respectively, they are not tests of structural change, per se. Bai and Perron and Kejriwal and Perron discuss estimation and testing procedures for multiple break dates of unknown timing, respectively, for stationary and nonstationary but cointegrated variables within a linear regression framework. Section three discusses the data. Section four contains empirical results pertaining to the time series properties of the price variables and the estimation of the long-run price linkage equations, followed by the estimation and simulations of the short-term price dynamics in section five. The paper is concluded with a reiteration of the procedure for the proposed structural change protocol for price analysis and a summary of major empirical results and implications.

2. Structural Change of Unknown Timing

In his seminal work, Perron (1989, 1997) establishes that many macroeconomic time series are trend-stationary if one allows a change in the intercept and/or the slope of the trend function at an exogenously specified break date (1989) or at an endogenously estimated break date (1997). This paper subscribes to Perron's modeling philosophy that it is important to allow for structural break when testing for unit root so that analysts would not confuse a structural break in the deterministic trend function with a unit root, biasing the results toward a false acceptance of a unit root.[3]

The current study adapts the three models put forth by Perron in 1997, which entertain a single break date in the trend function characterizing the time series. Specifically, the author considers the case where structural change happens gradually (innovation outlier model) and where it occurs instantaneously (additive outlier model). In the gradual change models of innovation outlier, Perron entertains structural change in the intercept and in both the intercept and slope of

[3] For a brief discussion of other unit root test procedures allowing for structural break see Wang and Tomek.

the trend function, whereas in the additive outlier model of sudden change the structural break occurs only in the slope coefficient.

For the innovation outlier models, Perron nests the null and alternative specifications into a composite model, in which the univariate time series is expressed as a function of its own lag, lagged terms of the first-difference, an intercept, a trend term, and dummy variable(s) pertaining to structural change in the trend function. The lagged terms of the first-difference of the variable are to purge autocorrelation in the residuals and to account for Perron's assumption that structural change occurs gradually.[4] A test for unit root is carried out by examining whether the coefficient for the lagged dependent variable is statistically different from one. In the additive outlier model, the unit root null is a special case of the breaking trend alternative and, hence, Perron invokes a two-step procedure: (i) removing the deterministic trend from the univariate time series by regressing the variable on an intercept, a trend term, and a dummy variable capturing the shift in the slope coefficient of the trend function as specified in the alternative model and (ii) applying the conventional ADF test by regressing the de-trended series on its own lag and lagged terms of the first difference, and checking for the coefficient of the lagged dependent variable. See Appendix for details on Perron's three models.

Subject to trimming of observations at both ends of the sample, the break date is obtained via a search procedure based on various criteria, such as minimizing the t-statistic associated with the parameter of the lagged dependent variable, or maximizing the absolute value of the t-statistic associated with a parameter involving structural break.[5] For any given break date candidate, the model is linear and thus can be estimated by ordinary least squares method. Based on asymptotic distributions of the test statistics, Perron simulates large sample critical values for the three models under various break date selection criteria and lag length selection methods for the first-differenced terms. The author also simulates small sample critical values.

Gregory and Hansen provide a multivariate extension of Perron's univariate tests by allowing for a structural break of unknown timing in the intercept, or the intercept and slope, of the conventional cointegration equation. For each possible break date candidate (subject to endpoint trimming), the authors estimate the cointegration equation with the structural break dummy variable included and conduct cointegration tests via checking for nonstationarity in the residuals using various methods, including the ADF test. In the ADF case, the break date is chosen by minimizing the t-statistic associated with the coefficient of the lagged residual in the unit root

[4] Perron assumes that the economy responds to a shock to the trend function the same way as it reacts to any other shock; thus the lagged terms on the first-difference will capture the dynamics in the residual, as well as the dynamics in the gradual structural change. Subject to a pre-specified lag truncation, Perron proposed two general to specific recursive lag length selection methods for the first-differenced lagged terms: a t-test on the significance of the last included lag and an F-test on the joint significance of the included lags.

[5] The rationale for the former criterion is to choose the break date that maximizes the power of rejecting the unit root null, whereas the second criterion focuses on lending the maximum support for the confirmation of the significance of the structural dummy coefficient in question. Note that the t statistic on the coefficient of the lagged dependent variable is a negative number.

equation. The authors derive the asymptotic distributions of the test statistics, which do not depend on any nuisance parameters other than the number of regressors. Approximate asymptotic critical values for the ADF test are then obtained via simulation for various specifications of structural change and up to four regressors.

In sum, the rejection of the null hypothesis of no cointegration in Gregory and Hansen provides evidence in favor of the alternative specification that the variables are cointegrated with a structural break. The authors emphasize that this hypothesis test, however, does not provide much evidence regarding whether a structural change has indeed occurred, because the alternative hypothesis includes the standard model of cointegration with no structural break as a special case. In other words, Gregory and Hansen's cointegration tests are not a test for structural change, per se. By the same token, Perron's unit root tests are not tests for structural change. Further, since the chosen break dates are identified via various criteria in the above unit root and cointegration tests, they need not be consistent estimates of the true break dates. The underlying motivation of Perron's and Gregory and Hansen's methods is not the estimation of the break dates; rather the focus is to improve the power of the conventional unit root and cointegration tests by allowing for a structural break. To estimate and test for break dates within a linear regression framework, this study adapts the procedures by Bai and Perron, and Kejriwal and Perron.

Bai and Perron (1998) consider a linear model of y on x, where both y and x are stationary. The procedure is to estimate multiple structural breaks of unknown timing by minimizing the sum of squared residuals, allowing for the case in which not all the parameters are subject to shifts. 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 global sum of squared residuals across all the m+1 segments. The chosen configuration of the m break dates is the one that has the lowest minimized residual sum of squared errors among all the allowable configurations. To ease the computation burden, especially when the number of breaks is large, Bai and Perron (2003a) present an efficient algorithm to obtain the global minimizers of the sum of squared residuals based on the principle of dynamic programming.

Bai and Perron (1998) demonstrate that the break fractions obtained in the manner above are consistent.[6] Further, the estimated break fractions converge to the true fraction at a fast rate of T, allowing one to obtain the standard root-T consistency and asymptotic normality for the estimated regression parameters. As to the limiting distribution of the break date estimates, the authors defer to Bai (1997), which estimates multiple breaks one at a time. Regarding tests for the number of breaks, Bai and Perron (1998) consider a sup Wald type test for the null hypothesis of no structural change versus the alternative hypothesis of a pre-specified number of breaks or of an unknown number of breaks, and a specific to general modeling strategy of sequentially testing for L breaks versus the alternative of L+1 breaks to consistently determine the appropriate number of breaks in the data. Asymptotic critical values for the test statistics under 5% trimming are reported in Bai and Perron (1998) for up to ten regression coefficients

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[6] The i^{th} break fraction is defined as T_i / T, where T_i is the i^{th} break date (e.g., 50^{th} observation) and T is the sample size (e.g., 200).

that are subject to structural change, and up to nine breaks. Critical values for trimming other than 5% are reported in Bai and Perron (2003b). Gauss code for the numerical implementation of Bai and Perron's 2003 dynamic-programming based procedure can be found on Perron's website at http://people.bu.edu/perron/code.html.

Extending Bai and Perron's procedure to allow for nonstationary data, Kejriwal and Perron (2008a, 2008b) discuss issues related to estimation, inference and testing of multiple structural changes of unknown timing in regression models involving nonstationary but cointegrated variables. It is worthwhile to note that Kejriwal and Perron's framework also allows for the inclusion of stationary regressors. The global minimization procedure of the break fractions is the same as that in Bai and Perron, involving an algorithm based on the principle of dynamic programming. However, the distribution of the break fraction estimates and the distribution of the sup Wald type and sequential test statistics are different due to the nonstationarity of the time series. Asymptotic critical values for the test statistics under 15% trimming are reported in Kejriwal and Perron (2008b) for various combinations of number of nonstationary regressors and numbers of nonstationary and stationary regressors subject to a maximum of five structural breaks. Critical values for trimming other than 15% are reported in the staff paper section on Perron's website.

3. Data and Motivation

The following sections contain an empirical analysis of the price transmission in the U.S. beef/cattle and pork/hog sectors. Within a threshold error correction framework, Goodwin and Holt, and Goodwin and Harper, GHH henceforth, investigate the retail, wholesale and farm price relationships in U.S. beef/cattle and pork/hog industries, respectively, by estimating long-run price linkage equations and short-run price dynamics among the cointegrated price series, focusing on the asymmetry of the error correction adjustment speed. The empirical application in the subsequent sections expands the contributions of GHH in several ways. First, structural change is entertained in unit root and cointegration tests and in the estimation of the long-run price linkage equation, using the methodology discussed in the previous section. Second, a dynamic least squares procedure is adapted that ensures unbiasedness in the estimated long-run price transmission parameters, taking into account the insights in Phillips and Loretan (1991) and Stock (1987). Third, asymmetry is entertained, not just in the short-run error correction adjustment speed, but also in the long-run price transmission coefficient. Fourth, the analysis may provide additional insights over GHH's weekly results, given that the currently analysis utilizes monthly data.

Retail and wholesale beef and pork prices are obtained from the USDA Economic Research Service for the period of January 1970 through February 2008. The retail beef price is a choice price reported in cents per pound, while the wholesale beef price is a composite price of wholesale cuts on a cent per pound retail equivalency. Likewise, the retail and wholesale pork prices are on a cent per pound basis and are in retail equivalency. As to farm prices, the 5-market steer price, from January 1989 through February 2008, is used to represent the national farm cattle price, while the 51-52% lean hog price, from January 1989 through February 2008, is used to represent the national farm hog price. To allow the farm price series to have the same number of observations as the retail and wholesale prices, the analysis also considers local farm cattle and hog data. Specifically, the weighted Nebraska direct price for 1,100 to 1,300 pound steers and the Iowa/Minnesota hog carcass price data are obtained from the Livestock Marketing Information Center (LMIC) for the period of January 1970 through February 2008. The original LMIC data for Iowa/Minnesota hog carcass price data are available only from January 1973 through February 2008. To impute the observations for the periods of January 1970 through December 1972, the relationship between the Iowa/Minnesota price series and an LMIC Omaha slaughter price series for 230-250 pound barrows and gilts (January 1970 through December 1979) is utilized. The LMIC data for the weighted Nebraska direct price for steers also have missing observations (from January 1970 through December 1991), and an LMIC Omaha average price series for 1,000 to 1,100 pound steers, ranging from January 1970 through December 1998 is used to impute the missing observations for the Nebraska series.[7]

4. Long-run Price Linkage

Results from the conventional ADF and Perron's structural change unit root tests are first reported and compared, followed by the results of cointegration tests using Engle and Granger's procedure and the structural change procedure of Gregory and Hansen. Depending on the test results regarding the time series property of the price variables, the procedure of Bai and Perron or Kejriwal and Perron is used to identify potential structural break dates in the retail/wholesale/farm price linkage. These break dates are then included in the estimation of the price linkage equation, allowing for asymmetry in the price transmission coefficients, using Phillip and Loretan's dynamic least squares procedure to ensure unbiasedness of the pass-through coefficient.

4.1 Unit Root Tests

Figure 1 presents the eight nominal price series (in logarithms) used in the analysis. Upon visual inspection, the wholesale price and the farm price, for both beef/cattle and pork/hog, appear to follow a similar pattern over time, albeit a mark-up exists for the wholesale price. Further, the beef prices and retail pork price appear to be suspect of nonstationarity, while wholesale pork and local (Iowa/Minnesota) farm hog prices appear to be stationary. Two specifications are considered for the conventional ADF tests; one includes drift and trend terms, and the other includes only the drift term. As to the structural change in the trend function, the specifications

[7] To impute missing observations for the Iowa/Minnesota carcass hog price series, the Omaha live price series is converted into a carcass based price by using the conversion factor of 74%, and then the following procedure is employed. First, monthly premiums/discounts between the two series are computed for the overlapping periods of January 1973 through December 1979, deflated by the producer price index for farm products. Second, to allow for seasonality, the computed real premiums/discounts are averaged to obtain a monthly average figure for each of the 12 months. Third, for each month of the missing period, the Omaha price is used to impute the Iowa/Minnesota price using the respective computed monthly premiums/discounts (after converting the monthly premium back into nominal terms using the PPI for that month). The same three step procedure is used to impute the missing values for the Nebraska cattle price series.

include a gradual intercept shift, gradual intercept and slope shifts, and a sudden slope shift.[8] The lag length for lagged terms of the first-difference of the series in the equation for unit root test is selected via Perron's general to specific F-sig criterion, starting with a lag truncation (maximum lag) of six months. As reported in Table 1, the conventional Dickey-Fuller test clearly suggests that beef/cattle prices are nonstationary, consistent with Goodwin and Holt's weekly data result, which uses ADF without the inclusion of drift and trend terms. Contrary to Goodwin and Harper's unit root conclusion (again employing an ADF test without drift and trend terms), the current ADF test results for the pork/hog series are mixed. Specifically, the retail pork and national farm hog prices are found to be nonstationary, while the wholesale pork and local farm hog prices are stationary at the 5% significance level. When structural change is allowed, using Perron's procedure, one continues to reject the unit root hypothesis for wholesale pork and local farm hog prices. As to the previous six variables that the ADF tests deem nonstationary, two are found to be stationary under Perron's tests: national farm cattle price and retail pork price. Although not all the nonstationary conclusions of ADF are overturned by Perron's tests, the two overturning cases do indicate the importance of entertaining structural change when conducting unit root tests.

4.2 Cointegration Tests

With few exceptions, the previous unit root test results strongly suggest that beef/cattle prices are nonstationary. On the other hand, one would tend to conclude that pork/hog prices are stationary, even though the results are more mixed, depending on the price variable in question and the unit root test employed. The analysis continues with cointegration tests of the beef/cattle prices, and to be on the safe side, cointegration of the pork/hog series is also checked for. A double log price transmission equation is estimated by regressing retail price on a constant, wholesale price and farm price for both the beef/cattle and pork/hog variables. Tests are then performed to ascertain the null hypothesis that the residual contains a unit root, and hence the price series are not cointegrated. Both the conventional procedure of Engle and Granger and the structural change procedure of Gregory and Hansen are employed. In both cointegration tests the unit root equation precludes drift and trend terms, because the variable involved in the test is a residual series. The structural change specification in the price transmission equation includes an intercept shift, with or without a structural break in the price transmission coefficient (i.e., slope shift).

The cointegration results are presented in Table 2. While Engle and Granger's test suggests that retail beef, wholesale beef, and local farm cattle prices are cointegrated, no such evidence is found when the national farm cattle price series is involved in the trivariate cointegration analysis. Similar conclusions can be drawn when using Gregory and Hansen's procedure, especially in the case where only the intercept is subject to structural change.[9] To determine the

[8] The series are first filtered through 12 monthly dummy variables to remove seasonality.
[9] Comparing the last two columns of Table 2, the rejection of the null hypothesis of no cointegration is much stronger in the case in which only the intercept is subject to structural break than the case where both the intercept and slope coefficients are allowed to break, suggesting a reduction in power due to over-parameterization in the latter specification.

missing link of the trivariate beef/cattle price relationship, cointegration tests are conducted for the three possible bivariate price relationships. The results of a lack of cointegration between the bivariate price variables still prevail when Engle and Granger's procedure is used, but become more diverse under Gregory and Hansen. Specifically, upon allowing for structural break, retail beef price is found to be cointegrated with wholesale beef price, and the latter is cointegrated with the national farm cattle price, but the link stops there because the national farm cattle price is not found to be cointegrated with retail beef price. As to the cointegration test results for the pork/hog price variables, both Engle and Granger and Gregory and Hansen's procedures provide evidence of cointegration among the retail/wholesale/farm prices regardless of whether national or local farm hog prices are used.

4.3 Identifying Break Dates in Price Transmission

The national farm cattle price and national farm hog price are precluded in the subsequent estimation of price transmission and price dynamics because the series contain 20 years less observations than the local farm price series. Previously, cointegration is found to exist among retail beef, wholesale beef, and local farm cattle prices, which are nonstationary. Additionally, retail pork, wholesale pork, and local farm hog prices are found to be cointegrated, though the cointegration test may not be necessary because there is sufficient evidence suggesting that the three pork/hog price series are stationary.

In identifying the break dates for the long-run price linkage, Kejriwal and Perron's procedure is used for the cointegrated, nonstationary beef/cattle prices, and Bai and Perron's procedure is used for the stationary pork/hog prices. As before, the variables are in logarithms, and in the pork/hog price equation, farm hog price is lagged two periods to avoid a negative sign on the coefficient of the farm hog price.[10] A maximum of three breaks for the study period is specified and a 20% trimming rate is adapted dictating that the minimum length of each segment of the regression be no less than 20% of the total number of observations. A result in Kejriwal and Perron relevant to the estimation of the break dates in the beef/cattle price linkage is that if the parameters of nonstationary regressors are subject to structural break, the confidence intervals for the break date estimates are correlated and, hence, more cumbersome to compute. For this reason, only intercept shifts are entertained in the beef/cattle price transmission equation. Since the correlation of the break date estimates is not an issue for stationary regressors, both intercept and slope shifts are allowed in the pork/hog price transmission equation.

The break date estimates and their 95% confidence intervals are reported in Table 3. Also included in Table 3 are the test statistics for the number of breaks using the sup Wald type tests of testing for no break versus a pre-specified number of breaks and the specific to general sequential test of L versus L+1 breaks. The sup F statistics in Table 3 clearly reject the null hypothesis of no structural break for each of the two price relationships. Further, the sequential tests identify two breaks for the beef/cattle price relationship, and three breaks for the pork/hog

[10] For the recent 20 years, the increasing prevalence of contracting can be used to justify the lagged specification, as producer contract prices for the current period are typically based on weighted averages of the past several periods, which can be parsimoniously captured by two lags.

price relationship. The estimated break dates for the beef/cattle price relationship are September 1980 and June 1993, whereas the estimated break dates for the pork/hog price transmission equation are October 1978, July 1987, and July 1997. The confidence intervals for the break date estimates are very tight, even though they become wider if heterogeneity in the residuals is allowed to vary across regimes (not reported in Table 3).[11] The estimated break dates for the pork/hog price equation are more readily explained: the 1978 break reflects the energy price hike which might fundamentally shift the cost structure, the 1987 break captures the beginning of vertical contracting, and the 1997 break signifies the beginning of an extended period of extremely low farm hog prices. As to the break date estimates for the beef/cattle price equation, the 1980 break reflects the 1978 energy crisis, incorporating cattle production lags, and the 1993 break may be a reflection of severe flooding and drought in certain parts of the nation hiking corn price to a then historical high in 1996.

4.4 Estimating Long-run Price Transmission Equations using Dynamic OLS

The asymptotic properties of the ordinary least squares (OLS) estimators of a static cointegration equation ($y_t = \alpha + \beta\, x_t + \mu_t$, where x and y are nonstationary and μ is stationary) are derived in Stock, which shows that OLS estimates of nonstationary regressor parameters, β, are asymptotically biased and non-normal, albeit super-consistent.[12] Further, Banerjee et al. (1993) find that the estimation of the static cointegration equations, which ignore the dynamics of the data generating process, can result in persistent and substantial finite sample bias in the long-run coefficient. The biasness and the non-normality of the cointegration parameter are problematic in the current context of price transmission analysis for several regards. First, the magnitude of the long-term price transmission parameter, β, is of intrinsic interest, and one would like to be able to obtain an unbiased estimate of this parameter and test if the coefficient is indeed equal to a certain value, such as unity in the case of perfect price transmission hypothesis. Second, as pointed out by Liu, Margaritis and Tourani-Rad (2007), the non-normality of the estimate of β is non-conducive for structural change analysis; one can only test for a level shift but not for a change in the degree of pass-through because the t-ratio is asymptotically normal only for the coefficient of the stationary regressor. Third, the bias in the estimation of β will be carried over to the estimation and analysis of the short-term dynamics, because the residual in the long-run equation will be utilized in the short-term dynamics as an error correction term.

Moving beyond the static framework, Phillips and Loretan propose a dynamic cointegration equation, showing that the resulting least squares estimator is asymptotically unbiased and normal.[13] In Phillips and Loretan's specification of PL(m,n,q) the static cointegration equation

[11] The estimated break dates are the same regardless whether or not heterogeneity in the residual is allowed.

[12] The OLS estimator of the cointegration parameter β is consistent and converges in probability faster than $T^{1-\delta}$ for any positive δ, where T is the sample size. A comparison to the typical root-T convergence rate of the standard asymptotic theory gives rise to the term super-consistent.

[13] Similar dynamic cointegration regression specifications have been proposed by Saikkonen (1991) and others.

of $y_t = \alpha + \beta\ x_t + \mu_t$ is augmented by m leading and n lagged terms of first-differences of the regressors, and q terms of the lagged residuals of the static cointegration equation. That is,

$$(1)\quad y_t \ = a + b\ x_t + \sum_{i=-m}^{n} V_i\ \Delta x_{t-i} + \sum_{i=1}^{q} x_i\ (y_{t-i} - a - b\ x_{t-i}) + e_t .$$

As pointed out in Phillips and Loretan, the inclusion of the lead terms 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 residual and the lagged first-difference of the regressors is to account for contemporaneous correlation between the residual in the static cointegration equation (μ_t) and the residuals that drives the nonstationary regressors x_t, i.e., ν_t in $x_t = x_{t-1} + \nu_t$. Phillips and Loretan's specification is also desirable from a practical viewpoint, because by incorporating lead and lag terms of the first-difference of the regressors, one explicitly accounts for the impact on the estimated relationship of past surprises and future expectations in market conditions (Liu, Margaritis and Tourani-Rad). Phillips and Loretan suggest that equation (1) be estimated by non-linear least squares. Alternatively, a two-step procedure of estimating the static components of the equation (i.e., $y_t = \alpha + \beta\ x_t + \mu_t$) and then estimating equation (1) with the regressors in the second to the last term being replaced by the lagged residuals from the first step can be employed.

As in the previous section of break date identification, log retail price is specified as a function of log wholesale and log farm prices (with farm hog price lagged two periods). Structural dummy variables are constructed using the previously estimated structural break dates and included in the estimation of the price transmission equation. Specifically, intercept dummy variables are included in the beef/cattle equation, and both intercept and slope dummy variables are included in the pork/hog equation. To entertain asymmetry in the pass-through coefficient, both the wholesale price and the farm price are decomposed into an increasing and a decreasing series based on the sign of the first-difference of the price variable in question. Since preliminary estimations suggest that PL(1,1,1) is sufficient for capturing the dynamics of the data generating process, the price transmission equations include only one lag term and one lead term of the first-difference of the wholesale and farm prices and one lag term of the residuals (i.e., μ_t). The equation is estimated via the two-step procedure discussed above.

The estimated price transmission elasticities, along with the constant term, are presented in Table 4 for the symmetric and asymmetric cases. Consider first the symmetrical case, in which the right-hand side price variables are not disaggregated into increasing and decreasing components. The estimated elasticities in the pork/hog price transmission equation are 0.74 and 0.20 for the wholesale and farm prices, respectively. While Goodwin and Harper's weekly results suggest that the price transmission elasticity is one for wholesale pork price, the current results, using monthly data, clearly indicate that the wholesale price transmission elasticity is far from unity. Also note that the farm hog price transmission elasticity is estimated as a negative number (-0.32) in Goodwin and Harper, a result corroborated by our preliminary finding of a negative price transmission coefficient when the hog price is not lagged. With regard to the beef/cattle equation, the estimated wholesale and farm price transmission elasticities are 0.84 and 0.15, respectively, compared to Goodwin and Holt's weekly wholesale and farm price transmission elasticities of 0.30 and 0.30, respectively. Given the estimated price transmission elasticities reported in Table 4, the percentage mark-up, evaluated at the sample mean, can be computed as 23% and 123% for the wholesale pork and farm hog prices, respectively, and 36% and 193% for the wholesale beef and farm cattle prices.[14] The above spreads are in line with the fact that the retail, wholesale and farm pork/hog prices in the sample average 194.17, 107.06 and 58.87 cents per pounds, and the averages for the beef/cattle price series are 254.79, 161.00 and 64.15 cents per pound.

Turn to the asymmetrical price transmission elasticities in Table 4. In all cases, the estimated price transmission elasticities appear very similar for the increasing and decreasing components, and differ only by a small magnitude from the symmetric case. For example, the wholesale price transmission coefficient in the pork/hog equation is 0.8011 when the wholesale price is increasing and 0.8027 when the wholesale price is decreasing, while the transmission elasticity under the symmetrical case is 0.74. Tests for symmetry of the wholesale (or farm) price transmission coefficients are conducted, with the farm (or wholesale) coefficients being allowed to be asymmetrical in one case and imposed to be symmetrical in the other case. Based on the p-values reported in the lower half of Table 4, the wholesale price transmission coefficients are found to be asymmetrical in the pork/hog equation regardless of whether the farm/hog price is allowed to be asymmetrical or not, while one cannot reject the hypothesis of symmetry for the farm hog price transmission elasticity. The results pertaining to the beef equation indicate that symmetry cannot be rejected for the wholesale price transmission elasticity, while farm price pass-through elasticities are found to asymmetrical when the wholesale price is imposed to be symmetrical.

The coefficients pertaining to the lead and lag terms of Phillips and Loretan's dynamic specification and those of the structural dummy variables are not reported in Table 4. A majority of the lead and lag terms are highly significant, and with their exclusion the Durbin-Watson statistics drop to a low level of about 0.33 and 0.42 for the pork/hog and beef/cattle equations, respectively (from about 2.0 in Table 4). The coefficients for the intercept dummies in the pork/hog and beef/cattle equations are all highly significant with a p-value close to zero for each. Also, the slope dummy coefficients in the pork/hog equation are predominately highly significant. The importance of allowing for structural change in the estimation of the pork/hog price transmission equation is manifested not just by the statistical significance of the structural dummy coefficients, but also by the lower coefficients of determination when those variables are excluded from the equations, dropping from 0.996 to 0.74 in the symmetrical case, for example. Excluding structural dummies also results in negative signs for the farm price coefficients in both the pork/hog and beef/cattle equations.

To conclude, the estimation results confirm the importance of including variables capturing the structural change and the dynamics of the underlying data generating process. The price transmission elasticities are estimated at about 0.75 to 0.85 for the wholesale prices and 0.15 to 0.20 for the farm prices, with asymmetry confirmed for the wholesale price in the pork/hog equation, and for the farm price in the beef/cattle equation, even though the estimated pass-through rate appears very similar for the increasing and decreasing price phases.

[14] Consider a price transmission equation of $\ln R = \alpha_0 + \alpha_1 \ln X_1 + \alpha_2 \ln X_2$, where R is the retail price and X_1 and X_2 are the wholesale and farm prices. The percentage mark-up between R and X_i, holding X_j constant ($j \neq i$), can be expressed as: $\frac{R - X_i}{X_i} = e^{a_0 + a_j \ln X_j + (a_i - 1)\ln X_i} - 1$.

5. Short-run Price Dynamics

Error correction models are used for the estimation of the short-run dynamics of the retail/wholesale/farm price series. In the case of beef/cattle prices, which were found to be cointegrated nonstationary variables, the approach can be justified by invoking Granger's representation theorem (Engle and Granger), stipulating that in a cointegrated system there exists an error correction mechanism such that 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, one justifies the use of an error correction model by acknowledging that the model is a special case of Hendry's Autoregressive Distributed Lag (ADL) model, which is appropriate for stationary time series (1997).[15]

In each of the three equations in an error correction system of retail/wholesale/farm prices, the contemporaneous change in the price in question 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 the price dynamics. Values for the error correction terms are the residuals in the appropriate Philips and Loretan's long-run price transmission equations reflecting the previous asymmetry test results in pass-through coefficients.[16] As in the estimation of the long-run price relationship, the previously identified structural break dates are incorporated into the short-run price dynamics estimation via the inclusion of appropriate intercept and slope dummy variables. Further, the asymmetrical aspects of adjustment speed are investigated by decomposing the error correction term into positive and negative components, treating zero as the threshold, and testing for the equality of associated coefficients. Finally, the short-run error correction price dynamics are estimated by OLS. Stock shows that the OLS estimator converges to limiting normal random variables at the usual rate of $T^{1/2}$, and because of the fast rate of convergence of the estimators of long-run parameters in the price transmission equation, the short-run estimates are asymptotically independent of the long-run estimates.

5.1 Estimation Results

The estimation results of the short-run price dynamics are presented in Table 5, with panel A reporting the p-values for symmetry tests of the adjustment speed in each of the short-run price

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

[16] Specifically, the error correction term in the pork/hog equation system comes from the long-run price transmission equation with symmetry imposed on the coefficient of the farm hog price, whereas that in the beef/cattle equation system is the residual from the equation with symmetry imposed on the coefficient of the wholesale beef price.

equations. With the exceptions of wholesale pork and farm hog equations, symmetry in the adjustment speed is rejected at the 10% or lower significance level. The result that the adjustment speed is symmetrical for the wholesale pork and farm hog equations, regardless of the signs of the equilibrium deviation, can be explained by the greater degree of vertical coordination (such as contracting and integration) in the processing and production segments of the pork/hog operation.[17] In a similar vein, the result of asymmetrical adjustment speed in the wholesale beef and farm cattle equations can be explained by the relatively limited amount of vertical coordination in the beef/cattle industry. Finally, the asymmetrical adjustment speed in both the retail pork and beef equations 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 above test results regarding adjustment speed symmetry are incorporated in the estimation of the short-run price dynamics and are reflected in the adjustment speeds reported in Panel B of Table 5. Consider the adjustment speed in the pork/hog equations first. The importance of considering error correction in estimating short-run price dynamics is manifested by the fact that all the adjustment speed coefficients are statistically different from zero. Also, the negative adjustment speeds in the retail pork, wholesale pork, and farm hog equations suggest that the price series are mean reverting. On the other hand, the signs of the adjustment speed in the beef/cattle equations suggest that retail beef, wholesale beef, and farm cattle price series are not mean reverting, a finding not inconsistent with the previous results that beef/cattle prices are nonstationary. Also interesting in the beef/cattle case is that the estimated adjustment speeds, under the negative deviation, are not significantly different from zero in all three equations, suggesting that price adjustments occur only when the equilibrium deviations are positive.

The coefficients pertaining to the structural dummy variables and lagged terms of price change are not presented in Table 5. However, Panel C reports the p-values associated with the significance of the above variables in explaining the price change equations. The structural dummy variables are highly significant with p-values close to zero for the pork/hog equations, but only marginally significant (p-value around 10% to 13%) for the beef/cattle equation, suggesting that the effect of structural change on pork/hog price dynamics is much more prominent than that on beef/cattle prices. The lagged price changes also explain the movement of the price in question (with the exceptions of lagged farm hog price changes on retail pork price movement and lagged wholesale pork price changes on farm hog price movement), suggesting a vigorous short-term dynamic price relationship among the retail, wholesale, and farm prices.

5.2 Impulse Response Analysis

To further understand the dynamic interaction among the retail, wholesale and farm prices, the estimated error correction system is used to compute the responses of the system to a one time

[17] In their impulse response analysis, Goodwin and Harper found that asymmetry is notably less apparent in the later part of their sample, which is characterized by increased vertical coordination and integration.

shock to a component of the innovations of the system, with the magnitude of the shock being the standard error of the estimate of the equation in question.

Figure 2 presents the responses of retail pork, wholesale pork and farm hog prices to shocks in the retail, wholesale and farm innovations, with each response being normalized by the size of the shock. In the figure, each row represents the responses of the three prices to a specific shock, while each column represents the responses of a specific variable to the three shocks. Retail shock appears to have significant immediate impact (i.e., contemporaneous impact for own variable and one-period lag impact for other variables) on all three prices, eliciting a response in retail price of about 100%, a response in wholesale price of about 70%, and a response in farm price of about 50% of the size of the retail price shock. In addition, the subsequent response of each price to the retail shock reduces quickly, within two to four months, to the range of 10% and -10% of the size of the shock, converging to zero after approximately three years. The second row of Figure 2 indicates that wholesale pork price shock is more important to the farm hog market than to the retail pork market, with the immediate impact on retail, wholesale and farm price being about 30%, 70% and 50%, respectively, of the size of the wholesale price shock. The subsequent response drops to within a range of 10% and -10% after about two to four months, and lingers on for about 18 to 24 months before dying out to zero. Finally, the shock in the farm hog price has minimal impact on retail pork price (with the immediate impact being about 10% of the size of the farm price shock), moderate impact on the wholesale pork price (with the immediate impact being about 30%), and has a significant impact on the farm price (with the immediate impact being about 70%).

The normalized responses of retail beef, wholesale beef and farm cattle prices to shocks in the retail, wholesale and farm equations, are presented in Figure 3. In addition to causing significant response in its own price, retail price shock is found to significantly affect wholesale and farm prices, eliciting a response of about 50% of the size of the retail price shock in each of the two prices. Similar to the pork/hog case, the subsequent effect of the retail beef price shock lingers between 10% and -10%, but appears to die out at a slower rate than the pork/hog case. As to a shock in the wholesale beef price, the immediate response of wholesale beef price is about 90% of the shock, whereas the retail and farm price responses are of the same magnitude of around 70%, with the subsequent effects also sharing a similar dynamic pattern. Finally, the last row of Figure 3 indicates that farm cattle price shock has a minimal effect on retail and wholesale beef prices, drawing out a response in each of the two prices of about 20% of the size of the farm price shock, and that the response of farm cattle price to its own shock is about 50%.

The results that retail pork price shock significantly affects wholesale pork and farm hog prices (70% and 50%, respectively) and that wholesale pork price shock affects farm hog price (50%) are different from Goodwin and Harper's conclusion that retail shocks are mainly contained in the retail market and the wholesale price shock affects only wholesale and retail prices. Likewise Goodwin and Holt's beef/cattle study indicates that the direction of price transmission flows from farm to wholesale to retail markets, with little feedback from retail to wholesale and farm, and little feedback from wholesale to farm. On the other hand, the current beef/cattle results suggest that retail beef price shock affects wholesale beef and farm cattle prices (50% and 50%, respectively) and that wholesale beef price shock affects farm cattle price (70%). The difference in the direction of price transmission may be due to the fact that the current study includes Phillips and Loretan's dynamic lead/lag specifications, which explicitly account for feedbacks of various directions. Additionally, the discrepancy may be due to the monthly data frequency used in the current study, enabling the model to capture the type of feedbacks that take longer than several weeks to materialize.

6. Summary and Conclusion

While change is the only constant in life, one may not always be aware of the timing of the transformation due to its subtlety. As such, the conventional econometric method of specifying, *a priori*, a definitive structural break date to account for change is not always feasible. Further, even if one is convinced that a structural change has occurred at a specific time, one still has to be mindful of the issues of data mining because this information has most likely been obtained by a series of formal and informal pretests of the data by the analyst or her predecessors. This argument has led to recent econometric developments of identifying structural change with unknown timing, while constructing distribution theory to explicitly account for the search of the break date. Subscribing to the above modeling philosophy, this paper proposes a protocol for agricultural and food price analysis, which goes beyond the existing practices of focusing on the time series properties of the variables in question, such as the existence of a unit root and of a cointegrating relationship among nonstationary prices, by entertaining structural breaks when conducting those tests. In so doing, one hopes to increase the power of the tests, avoiding the pitfall of confusing structural change with a unit root and structural change with a lack of long-run cointegration relationship. In addition, the proposed "structural change protocol" includes a sequence of procedures for estimating the break dates, which are to be incorporated into the analysis of the long-run price relationship and the short-run price dynamics.

Specifically, this study applies Perron's unit root test and Gregory and Hansen's cointegration test (both of which allow for structural change of unknown timing) to the U.S. beef/cattle prices and pork/hog prices. To obtain consistent estimates of the break dates in the retail beef, wholesale beef and farm cattle price linkage equation and those in the retail pork, wholesale pork and farm hog price relationship, the study employs Bai and Perron's and Kejriwal and Perron's global sum of squares minimization procedure to search for the optimal break dates, utilizing the principle of dynamic programming to ease the computation burden. With the identified break dates, the dynamic least squares framework of Phillip and Loretan is adapted for the specification of the long-run price linkage equations to ensure the unbiasedness of the estimated price transmission parameters. The lead and lag terms of the regressors in Phillips and Loretan allow for feedbacks from the left-hand side price variable to the right-hand side regressors, while incorporating in the estimation past surprises and future expectations. Within an error correction framework, the estimated error terms in the long-run equations and the identified break dates are then utilized in the estimation of the short-run dynamics. Asymmetry in the long-run pass-through rate (depending on whether the price change in question is increasing or decreasing) and in the short-run adjustment speed (depending on whether the equilibrium deviation is positive or negative) are both entertained in the estimation.

The empirical results strongly suggest that beef/cattle prices are nonstationary, while there is ample evidence to support the conclusion that pork/hog prices are stationary. There are cases in which the conventional Dickey-Fuller test suggests the presence of a unit root in the price series in question, only to be overturned by Perron's test, reminiscent of Wang and Tomek's admonition of the need to be critical regarding the results from the conventional unit root tests. As to the existence of a long-run relationship, the retail beef, wholesale beef and local farm cattle prices are found to be cointegrated, as well as the retail pork, wholesale pork and local farm hog prices, even though testing for cointegration for the stationary pork/hog price series is not necessary. Results from the break date analysis suggest that there are three structural break dates in the pork/hog price linkage equation, reflecting the energy crisis of the late 1970's, the beginning of vertical contracting in the mid 1980's, and the period of extremely low hog prices in the late 1990's. Regarding structural breaks for the beef/cattle price linkage equation, two break dates are identified, corresponding to the above mentioned energy crisis and severe weather patterns in the U.S. in the early to mid 1990's. Given the break date estimates, the long-run price transmission parameters are estimated by regressing retail price on wholesale and farm prices and other terms. With respect to retail pork price equation, the long-run price transmission elasticity is about 0.80 and 0.20 for the wholesale pork and farm hog prices, respectively, and asymmetry in pass-through is confirmed for the wholesale pork price but not for the farm hog price. As to the retail beef price equation, the long-run price transmission elasticity is about 0.85 and 0.15 for the wholesale beef and farm cattle prices, respectively, and asymmetry in pass-through is confirmed for the farm cattle price but not for the wholesale beef price. Given the estimated price transmission elasticities, the percentage mark-up contained in the retail pork price is 23% and 123% of the wholesale pork and farm hog prices, respectively, while the percentage mark-up contained in the retail beef price is 36% and 193% of the wholesale beef and farm cattle prices. Results from the analysis of short-run price dynamics suggest that pork/hog prices are mean reverting while the beef/cattle prices are not. In addition, with the exceptions of the wholesale pork and farm hog price dynamics, adjustment speeds are found to be asymmetrical. Finally, while previous studies find 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 multi-directional in nature.

Figure 1: Nominal Price Series in Natural Logarithms

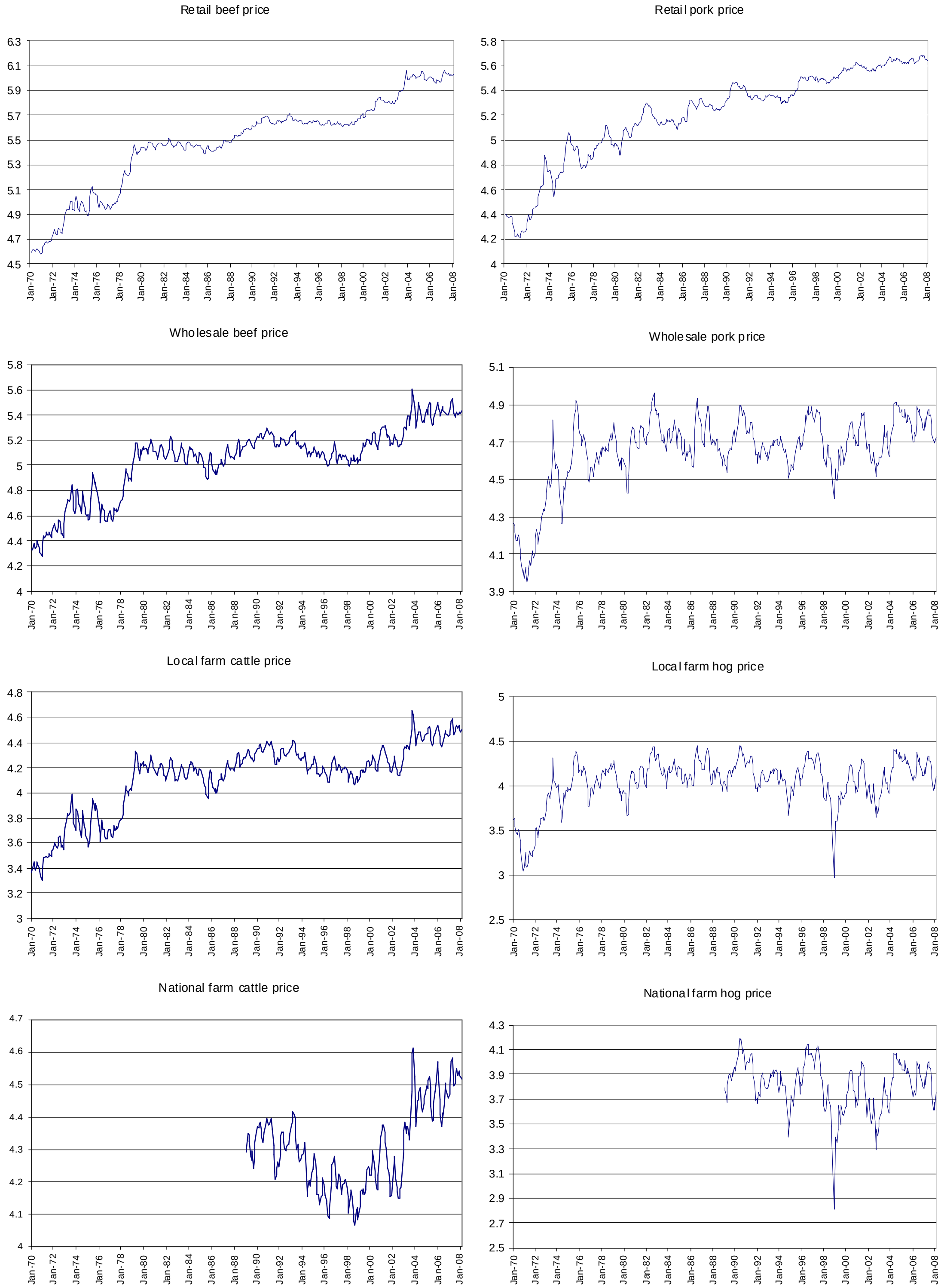

Figure 2: Standardized Impulse Responses of Pork/Hog Prices

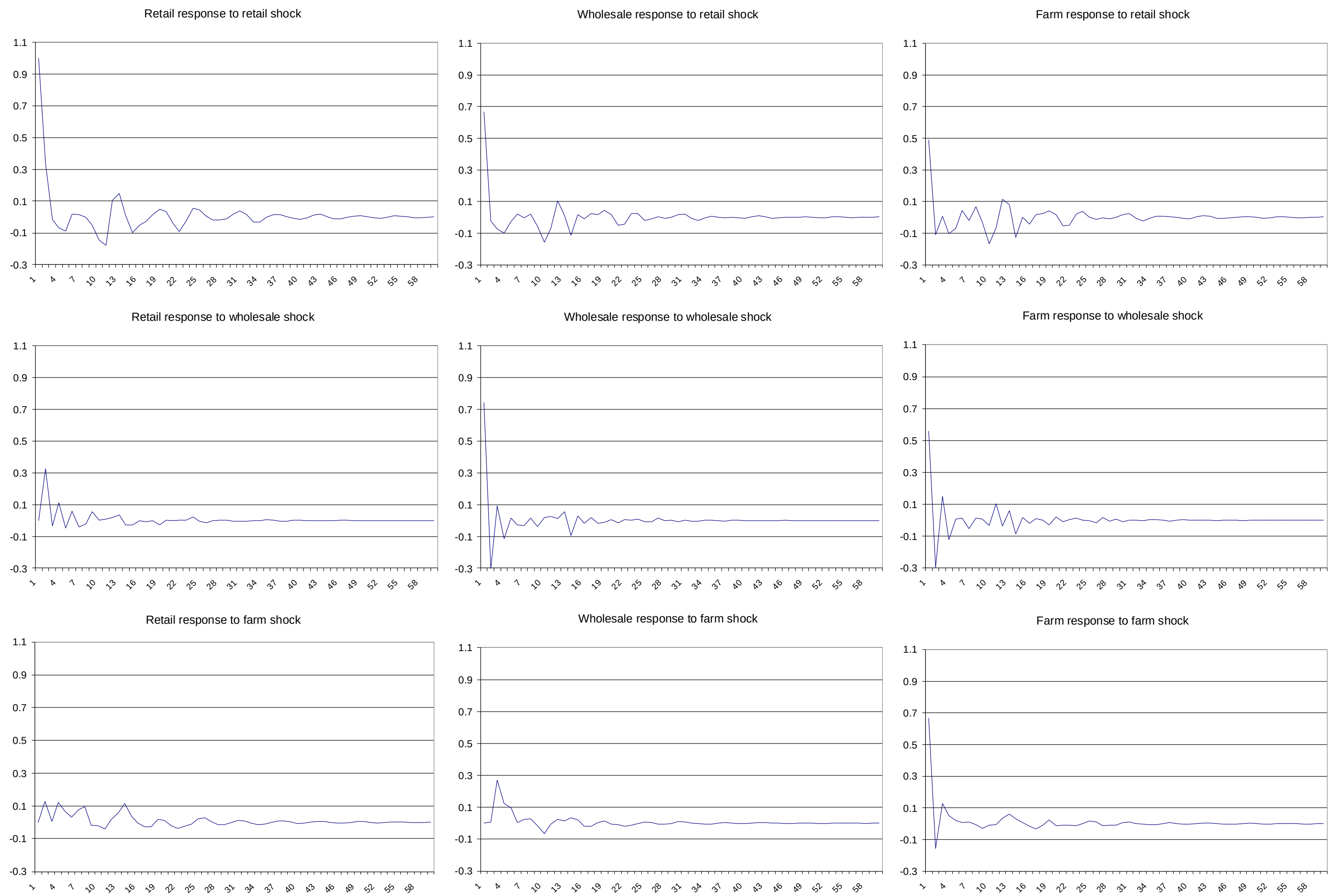

Figure 3: Standardized Impulse Responses of Beef/Cattle Prices

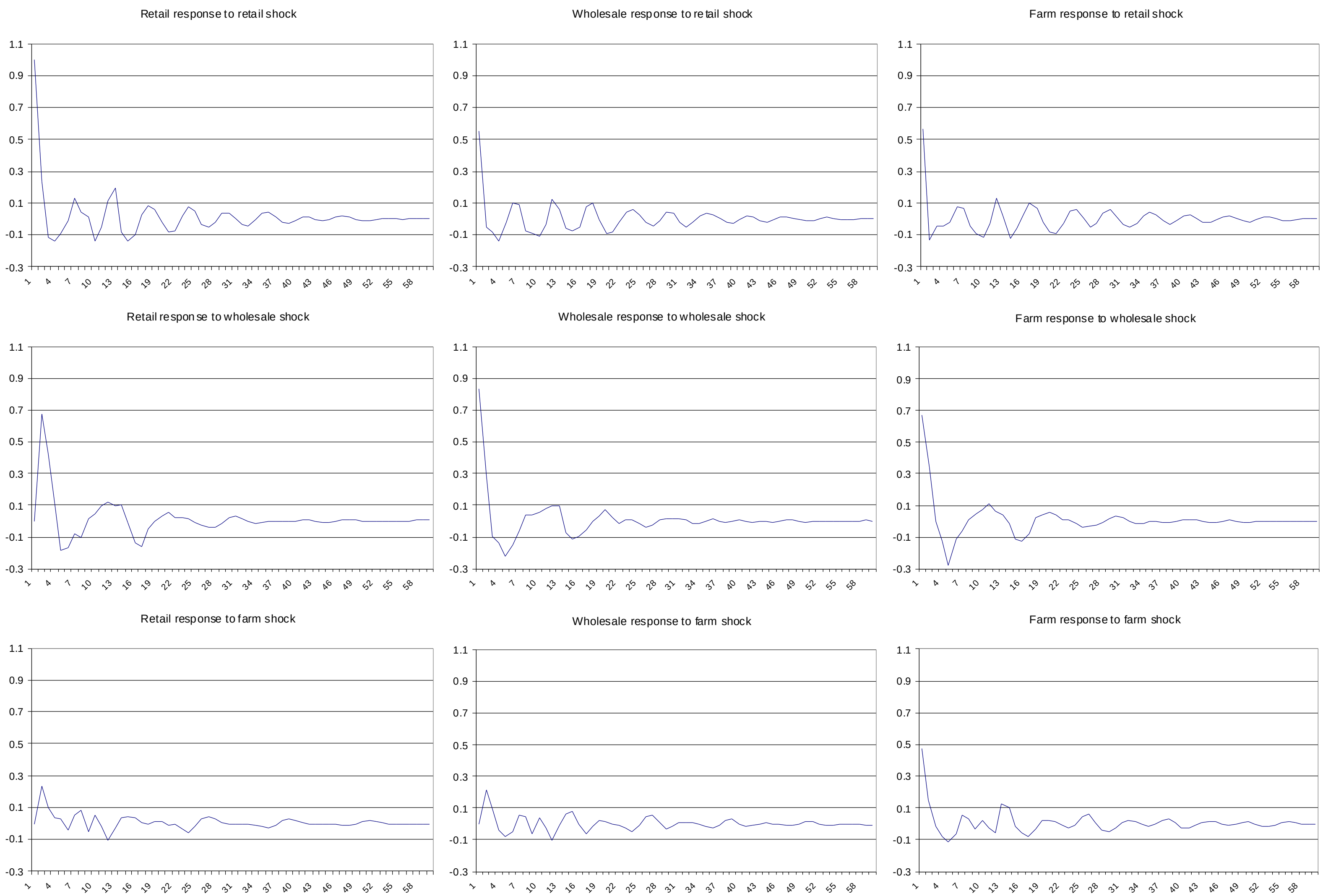

Table 1: Unit Root Tests [a, b]

	t-statistics on the lagged coefficient of the dependent variable				
	Dickey-Fuller w/drift and trend	**Dickey-Fuller w/drift**	**Perron w/intercept shift (gradual change)**	**Perron w/intercept and slope shifts (gradual change)**	**Perron w/slope shift (sudden change)**
Retail Beef	-2.39	-1.19	-4.00	-4.02	-2.80
Wholesale Beef	-3.00	-2.34	-4.26	-4.11	-3.32
Local Farm Cattle	-2.93	-2.41	-4.36	-4.27	-3.37
National Farm Cattle	-1.82	-1.25	-4.39	-6.3***(OT)	-5.29***(OT)
Retail Pork	-2.74	-1.88	-5.33**(OT)	-6.09**(OT)	-5.35***(OT)
Wholesale Pork	-3.69**	-3.36**	-5.88***	-6.29***	-5.85***
Local Farm Hog	-3.56**	-3.46***	-5.73***	-6.12**	-5.78***
National Farm Hog	-2.72	-2.65*	-3.58	-3.84	-4.21

[a] Single (*), double (**) and triple asterisks (***) indicate statistical significance at 10%, 5% and 1% levels, respectively.

[b] The notation OT indicates the overturn of the conclusion under the conventional procedure of not allowing for structural change.

Table 2: Cointegration Tests [a, b]

	t-statistics on the lagged coefficient of the residual		
	Engle & Granger	**Gregory & Hansen**	
		w/level shift	**w/level and slope shift**
Retail beef on constant wholesale beef, local farm beef	-3.21**	-5.03**	-4.88
Retail beef on constant, wholesale beef, national farm beef	-2.12	-3.45	-4.37
Retail beef on constant, wholesale beef	-3.14*	-4.84** (OT)	-4.84*
Retail beef on constant, national farm beef	-1.80	-3.36	-4.57
Wholesale beef on constant, national farm beef	-2.21	-5.05** (OT)	-4.89*
Retail pork on constant, wholesale pork, local farm pork	-3.82***	-5.71***	-5.45*
Retail pork on constant, wholesale pork, national farm pork	-5.18***	-5.18**	-5.27*

[a] Single (*), double (**) and triple asterisks (***) indicate statistical significance at 10%, 5% and 1% levels, respectively.

[b] The notation OT indicates the overturn of the conclusion under the conventional procedure of not allowing for structural change.

Table 3: Structural Break Tests and Estimated Break Dates [a]

Retail Beef on Constant, Wholesale Beef, Local Farm Cattle

SupF(1)	SupF(2)	SupF(3)	SupF(2|1)	SupF(3|2)
602.83***	519.58***	375.89***	171.53***	0.00

Break Dates	95% Lower Bound	95% Upper Bound
Sept. 1980	July 1980	Jan. 1981
June 1993	Apr. 1993	Aug. 1993

Retail Pork on Constant, Wholesale Pork, Lagged Local Farm Hog

SupF(1)	SupF(2)	SupF(3)	SupF(2|1)	SupF(3|2)
120.34***	216.69***	562.19***	100.46***	45.48***

Break Dates	95% Lower Bound	95% Upper Bound
Oct 1978	Sept. 1978	Nov. 1978
July 1987	June 1987	Aug. 1987
July 1997	June 1997	Aug. 1997

[a] Triple asterisks (***) indicate statistical significance at the 1% level.

Table 4: Estimated Long-Run Price Transmission Elasticities [a, b]

	Retail Pork			Retail Beef		
	Symmetric	Asymmetric		Symmetric	Asymmetric	
Constant	0.5942		0.4920	0.4433		0.3706
	(7.62)		(6.17)	(15.51)		(10.36)
Wholesale Price (WP)	0.7436	Increase	0.8011	0.8441	Increase	0.8683
	(18.34)		(19.30)	(39.49)		(32.16)
		Decrease	0.8027		Decrease	0.8689
			(19.36)			(32.21)
Farm Local Price (FP)	0.1982	Increase	0.1567	0.1526	Increase	0.1414
	(6.88)		(5.32)	(7.28)		(5.34)
		Decrease	0.1578		Decrease	0.1423
			(5.35)			(5.37)
Durbin-Watson	2.05		2.05	1.95		1.82
Adjusted R^2	0.996		0.996	0.998		0.997
P-Values for Symmetry Tests						
WP w/FP allowed to be asymmetrical			0.03			0.31
WP w/FP imposed to be symmetrical			0.03			0.28
FP w/WP allowed to be asymmetrical			0.22			0.22
FP w/WP imposed to be symmetrical			0.19			0.02

[a] The specification of the model is PL(1,1,1), where the retail price is regressed on wholesale price, farm price, one-period lag term and one-period lead term of the first-difference of the wholesale and farm prices, and one-period lag term of the residual of the static price transmission equation (i.e., retail price on wholesale and farm prices). The pork equation includes both intercept and slope structural changes, and the farm price is lagged two periods. The beef equation includes intercept structural change.

[b] Figures in parentheses are t-statistics.

Table 5: Estimation Results of Short-Run Price Dynamics [a, b]

	Pork			Beef		
	Retail	**Whole.**	**Farm**	**Retail**	**Whole.**	**Farm**
Panel A:						
P-values for Symmetry in Adjustment Speed						
	0.09	0.17	0.91	0.07	0.05	0.05
Panel B:						
Error Correction Coefficient						
Symmetry		-0.2919	-0.3287			
		(-3.32)	(-2.15)			
Asymmetry:	-0.3652			0.1912	0.6698	0.8368
Positive	(-3.19)			(1.98)	(2.64)	(3.25)
Asymmetry:	-0.1364			-0.0341	0.0311	0.1838
Negative	(-2.54)			(-0.55)	(0.19)	(1.10)
Panel C:						
P-values for the Significance of the Variables						
Intercept Dummies	0.00	0.00	0.00	0.13	0.10	0.10
Slope Dummies	0.00	0.00	0.00	na [c]	Na	na
Lagged Retail ΔP [d]	0.00	0.00	0.00	0.00	0.01	0.00
Lagged Whole. ΔP	0.05	0.00	0.13	0.00	0.02	0.05
Lagged Farm ΔP	0.30	0.00	0.00	0.00	0.00	0.00
Durbin-Watson	1.98	1.50	1.61	1.97	1.97	1.98
Adjusted R^2	0.45	0.46	0.51	0.47	0.24	0.23

[a] In each equation the left-hand side variable is the change in price and the right-hand side variables include past changes (up to 13 lags) of the retail, wholesale, and farm prices, a one-period-lag error correction term, and structural change dummy variables.

[b] Figures in parentheses are t-statistics.

[c] Slope dummies are not utilized for the beef/cattle equations because the variables are nonstationary.

[d] ΔP is the first-difference of the price in question.

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Appendix: Perron's Unit Root Tests

Innovation Outlier Model 1: Entertain a gradual change in the intercept of the trend function at an unknown break date τ under both the unit root null hypothesis and the breaking trend alternative hypothesis.

$$(null) \quad y_t \quad = m + a \; y_{t-1} + d \; D\_TEMP_t + e_t, \text{ with } a = 1$$
$$(alternative) \quad y_t \quad = m + b \; t + (m_2 - m) \, D\_PERM_t + e_t$$

where $D\_TEMP_t = 1$ for $t = \tau + 1$ and zero otherwise, $D\_PERM_t = 1$ for $t > \tau$ and zero otherwise. Perron arrives at a composite model by (i) nesting the models under the null and alternative hypotheses and (ii) including lagged terms of first-difference to reflect the assumption that the structural change is of a gradual type (Perron 1989).

$$(composite) \quad y_t \quad = m + a \; y_{t-1} + b \; t + d \; D\_TEMP_t + q \; D\_PERM_t + \sum\nolimits_{i=1}^{L} c_i \, \Delta y_{t-i} + e_t$$

where $\theta \equiv \mu_2 - \mu$. Under the null hypothesis of a unit root $\alpha = 1$.

Innovation Outlier Model 2: Entertain a gradual change in both the intercept and the slope of the trend function at an unknown break date τ.

$$(null) \quad y_t \quad = m + a \; y_{t-1} + d \; D\_TEMP_t + (m_2 - m) \, D\_PERM_t + e_t, \text{ with } a = 1$$
$$(alternative) \quad y_t \quad = m + b \; t + (m_2 - m) \, D\_PERM_t + (b_2 - b) \, D\_TREND_t + e_t$$

where $D\_TREND = t$ for $t > \tau$ and zero otherwise. As in model 1, the composite model is obtained by (i) nesting the models under the null and alternative hypotheses and (ii) including lagged terms of first-difference to reflect the assumption that the structural change is of a gradual type.

$$\begin{aligned}(composite) \quad y_t \quad &= m + a \; y_{t-1} + b \; t + d \; D\_TEMP_t + q \; D\_PERM_t \\ &+ g \; D\_TREND_t + \sum\nolimits_{i=1}^{L} c_i \, \Delta y_{t-i} + e_t\end{aligned}$$

where $\gamma \equiv \beta_2 - \beta$. Under the null hypothesis of a unit root $\alpha = 1$.

Additive Outlier Model: Entertain a sudden change in the slope of the trend function at an unknown break date τ.

$$(null) \quad y_t \quad = m + a \; y_{t-1} + (m_2 - m) \, D\_PERM_t + e_t, \text{ with } a = 1$$
$$(alternative) \quad y_t \quad = m + b \; t + (b_2 - b) \, D\_R\_TREND_t + e_t$$

where D_R_TREND is a rescaled trend variable and equals $t - \tau$ for $t > \tau$ and zero otherwise. Under this model, a change in the slope is allowed but both segments of the trend function are joined at the break date. The two-step procedure in Perron (1989, 1997) is as follows. First, remove the trend from the series by estimating the alternative model. Second, conduct an augmented Dickey-Fuller unit root test on the detrended series.

$$\text{(Step 1)} \quad y_t \quad = m + b\ t + g\ D\_R\_TREND_t + y_t^{\sim}$$

$$\text{(Step 2)} \quad y_t^{\sim} \quad = a\ h_{t-1} + \sum_{i=1}^{L} c_i\ \Delta n_{t-i} + e_t$$

where $y^{\sim}$ is the residual in the first step regression. Under the null hypothesis of a unit root $\alpha = 1$.