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Further Evidence of Price Transmission and Asymmetric Adjustment in the U.S. Beef and Pork Sectors

Brenda L. Boetel and Donald J. Liu

Brenda L. Boetel is an assistant professor in the Agricultural Economics Department at the University of Wisconsin-River Falls. Donald J. Liu is a professor in the Department of Applied Economics at the University of Minnesota. Senior authorship is equally shared.

Brenda L. Boetel can be reached at Brenda.boetel@uwrf.edu. Donald J. Liu can be reached at dliu@umn.edu.

Selected Paper prepared for presentation at the American Agricultural Economics Association Annual Meeting, Orlando, FL, July 27-29, 2008.

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The idea that prices convey sufficient information regarding efficient resource allocation is perhaps one of the most powerful insights that economists have offered. Under the assumption that the market operates competitively and efficiently, price properly reflects changes in the underlying determinants of the supply and demand. Of interest to economists is the magnitude and speed of price adjustments at one market level facing shocks at a different level within the market chain. Furthermore, it has been a great concern whether the price adjustment is asymmetric with respect to the nature of the shock (e.g., positive or negative, large or small). Clearly, the magnitude, speed, and asymmetry of price transmissions of shocks from one level to another are manifestations of the behavior of market players, reflecting the underlying economic environment in which they reside, and has welfare implications for consumers, wholesalers and producers.

The literature on price transmission is long-standing, where much of the earlier work focused on price asymmetry and was initiated by agricultural economists (e.g., Tweeten and Quance, 1969; Wolffram, 1971; Houck, 1977). Meyer and von Cramon-Taubadel (2004) aptly classify the econometric models for investigating price transmissions into two major categories, the so-called “pre-cointegration approaches” and the “cointegration based approaches”. The pre-cointegration approach descends from the earlier work of above mentioned agricultural economists, whereas the cointegration approach, which allows for partial adjustment type error correction, has become the mainstay of the method of investigation for the past two decades after the seminal work of Engle and Granger (1987).

The focus of the current study is on the price transmission in the U.S. cattle/beef and hog/pork industries. In the spirit of Engle and Granger, Goodwin and Holt (1999) and Goodwin and Harper (2000), GHH henceforth, investigate the retail, wholesale and farm price relationships in U.S. beef and pork industries, respectively. Specifically, GHH estimate the long run price linkage equations among the cointegrated price series while allowing for the correction of deviations from the long run equilibrium in their specification of the short term price dynamics. Further, GHH subscribe to Balke and Fomby's (1997) threshold specifications, entertaining different speeds for error correction based on the magnitude of the deviations. In fact, a major focus of GHH's studies is on the asymmetry of the error correction adjustment path.

The current study expands the contribution of GHH in three areas. First, in light of the recent advancement made by Perron (1997) in unit root tests, we re-examine in a more comprehensive manner GHH's conclusion that the weekly U.S. cattle/beef and hog/pork price series are nonstationary. This is significant because GHH's cointegration approach to price transmission builds on the premise that the time series are nonstationary, albeit cointegrated. Contrary to Perron (1997), the conventional augmented Dickey-Fuller (1979) test that GHH adopt has low power in discriminating against the unit root null because it does not entertain the possibility of a structure break in the deterministic trend function. Second, we examine more closely the estimation procedure surrounding the long run price linkage equation, taking into account the insights in Phillips and Loretan (1991) and Stock (1987) to ensure unbiasedness in the estimated long run price transmission parameter. Moreover, we entertain structural changes in the long run price linkage equation in light of the dramatic changes in the industries during the past three decades, such as the increased vertical integration, coordination and consolidation, and the increased adoption of the new information management technologies. Third, the current study employs two data sets with different frequencies, weekly data (as in GHH) and monthly data. Various theories have been proposed to explain asymmetry in price adjustment, such as shorter term problems associated with market information and longer term problems related to market power and price fixing. By employing data sets with different frequencies the researchers are in a better position to gain insight toward the underlying causes of price asymmetry.

The Pre-cointegration and Cointegration Approaches to Price Transmission

Initially motivated by the concern for supply irreversibility of agricultural commodities with respect to an increase/decrease in output price, the pre-cointegration method to price transmission has dwelt on the issue of asymmetry. Consider a linear price transmission equation of $y_t = \alpha + \beta x_t + \mu_t$, where x and y are two prices at different levels in the market chain, and μ_t is the error term. To account for asymmetry in the effect of x on y, Tweeten and Quance decomposed the explanatory price series x into an increasing and a decreasing component:

(1) $$y_t = a + b^+ D_t x_t + b^- (1 - D_t) x_t + m_t,$$

where D_t takes the value of one if $\triangle x_t$ (i.e., $x_t - x_{t-1}$) is positive and zero otherwise. A test of lack of asymmetry is then conducted via the conventional t statistic on $\beta^+ = \beta^-$. To account for persistent effects on y of past shocks in x, Wolffram modifies (1) by re-defining the increasing and decreasing components of x_t as the summation up to t of positive and negative $\triangle x_t$, respectively:

$$(2) \quad y_t = a + b^+(x_0 + \sum_{k=1}^{t} D_k \, \Delta x_k) + b^-(x_0 + \sum_{k=1}^{t} (1 - D_k) \, \Delta x_k) + m_t \, ,$$

where x_0 is the initial value of x_t at t = 0 and the indicator variable D is similarly defined. To streamline estimation, Houck takes the first difference of (2) while retaining the constant term and proposes:

$$(3) \quad \Delta y_t = a + b^+ \, D_t \, \Delta x_t + b^- \, (1 - D_t) \, \Delta x_t + e_t \, ,$$

where e_t is the transformed error term. To allow for distributed lag effects, Ward (1982) extends Houck's specification by including L lagged terms of $D_t \, \Delta x_t$ and $(1 - D_t) \, \Delta x_t$ in the price transmission equation:

$$(3') \quad \Delta y_t = a + \sum_{k=1}^{L} b_k^+ \, D_{t-k+1} \, \Delta x_{t-k+1} + \sum_{k=1}^{L} b_k^- \, (1 - D_{t-k+1}) \, \Delta x_{t-k+1} + e_t \, .$$

The above pre-cointegration models have been the workhorse over the past three decades for analyzing price transmission of various agricultural commodities. For example, Ward investigates price asymmetries for fresh vegetables, while Kinnucan and Forker (1987) examine major dairy products. Goodwin and Holt, and Goodwin and Harper (GHH) criticize the pre-cointegration models to price transmissions for their lack of attention to the time series properties of the data. Specifically, GHH are concerned with the omissions of error correction terms in the price adjustment process, and thus the incompatibility of the above models with the concept of long run equilibrium. GHH espouse the cointegration based approach, which estimates a long run cointegrating relationship between the prices in question and a short run price adjustment process with a built-in mechanism for error correction.

The cointegration literature concerns the long run relationship among nonstationary time series. Granger and Newbold (1974) point out the pitfall of spurious regression when variables are nonstationary. After Nelson and Plosser (1982) found that macroeconomic time series tend to contain unit roots there has been an abundance of research on ways of coping with the problem of spurious regression facing nonstationary data. Engle and Granger provide a two step procedure for estimating and testing for long run cointegration relationship between nonstationary variables. Nonstationary time series are said to be cointegrated if stationarity can be achieved via certain linear combinations of contemporaneous values of the variables. In this case, a cointegration equation of the variables such as:

(4) $y_t = \alpha + \beta x_t + \mu_t$,

is immune to the problem of spurious regression because the residual series is a stationary process, ensuring a meaningful long run equilibrium relationship between y and x. The stationarity of μ_t can be tested via the augmented Dickey-Fuller unit root test, using the critical values provided by Engle and Granger to account for the fact that the residual series has to be estimated from the regression. Further, the Granger representation theorem (in Engle and Granger) stipulates 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 current context of $y_t = \alpha + \beta x_t + \mu_t$ the short term dynamics can be expressed as:

$$(5a)\quad \Delta y_t = u_1\,(y_{t-1} - a - b\,x_{t-1}) + \sum\nolimits_{k=1}^{L} g_{1k}\,\Delta x_{t-k} + \sum\nolimits_{k=1}^{L} d_{1k}\,\Delta y_{t-k} + e_{1t}$$

$$(5b)\quad \Delta x_t = u_2\,(y_{t-1} - a - b\,x_{t-1}) + \sum\nolimits_{k=1}^{L} g_{2k}\,\Delta x_{t-k} + \sum\nolimits_{k=1}^{L} d_{2k}\,\Delta y_{t-k} + e_{2t}$$

where the term ($y_{t-1} - \alpha - \beta x_{t-1}$) is the deviation from the long run equilibrium at the previous period (i.e., μ_{t-1}) and ν_i is the adjustment speed of the error correction. Aside from asymmetry, note that the cointegration equation in (4) is reminiscent of equation (1) of the pre-cointegration approach, with the exception that (1) does not address the unit root properties (or lack thereof) of the residual series. Again, aside from asymmetry, the short run dynamic price adjustment process in (5a) can be regarded as a generalization of equation (3′) in that, while both include as regressors a distributed lag component (i.e., Δx_{t-k}'s), equation (5a) includes also the error correction term and an autoregressive component (i.e., Δy_{t-k}'s).

The asymptotic properties of the least squares estimators of (4) are derived in Stock, who shows that the ordinary least squares (OLS) estimators of α and β are super-consistent, converging to their probability limits at a rate faster than the one typically associated with the conventional asymptotic theory, $T^{½}$ with T being the sample size.[1]

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1. The intuition of super-consistency is given in Stock. In the case of stationary time series the sum of squares error (SSE) function obtains its minimum at the true value of the parameters, while the SSE has a larger but finite value for other values of the parameters; the estimators converge to their probability limits at the usual rate of $T^{½}$, where T is the sample size. In contrast, when the time series are nonstationary the linear combination of the variables, using the true α and β, is stationary and has a finite second moment, while all other linear combinations are nonstationary and have an infinite unconditional second moment. The OLS estimators of the cointegration parameters are

However, the OLS estimators of the parameters in (4) are not asymptotically normal and, surprisingly, are asymptotically biased. The behavior of the OLS estimators of the parameters in the short term dynamics in (5) is quite different. Stock shows that these estimators converge to limiting normal random variables at the usual rate of $T^{½}$, and because of the fast rate of convergence of the estimators of α and β in (4) the short run parameter estimators in (5) are asymptotically independent of the estimators in (4).

Motivated by adjustment cost considerations, GHH invoke Balke and Fomby's threshold specification to entertain the possibility that the short term dynamics behave in different manners depending on the magnitude of deviation from the equilibrium. For example, consider the case of a two regime threshold model where there exist two processes dictating the short term dynamics depend on whether the absolute magnitude of the equilibrium error is within a range defined by a threshold or not. In this case, equation (5a) and (5b) become:

$$(6a)\quad \Delta y_t = \begin{cases} u_1^{(1)}\, \hat{m}_{t-1} + \sum_{k=1}^{L} g_{1k}^{(1)}\, \Delta x_{t-k} + \sum_{k=1}^{L} d_{1k}^{(1)}\, \Delta y_{t-k} + e_{1t} & \text{if } \left| \hat{m}_{t-1} \right| \le h \\ u_1^{(2)}\, \hat{m}_{t-1} + \sum_{k=1}^{L} g_{1k}^{(2)}\, \Delta x_{t-k} + \sum_{k=1}^{L} d_{1k}^{(2)}\, \Delta y_{t-k} + e_{1t} & \text{if } \left| \hat{m}_{t-1} \right| > h \end{cases}$$

$$(6b)\quad \Delta x_t = \begin{cases} u_2^{(1)}\, \hat{m}_{t-1} + \sum_{k=1}^{L} g_{2k}^{(1)}\, \Delta x_{t-k} + \sum_{k=1}^{L} d_{2k}^{(1)}\, \Delta y_{t-k} + e_{2t} & \text{if } \left| \hat{m}_{t-1} \right| \le h \\ u_2^{(2)}\, \hat{m}_{t-1} + \sum_{k=1}^{L} g_{2k}^{(2)}\, \Delta x_{t-k} + \sum_{k=1}^{L} d_{2k}^{(2)}\, \Delta y_{t-k} + e_{2t} & \text{if } \left| \hat{m}_{t-1} \right| > h \end{cases}$$

where the superscripts (1) and (2) in (6a) and (6b) denote inner and outer regimes, respectively, and h is the threshold parameter, which can be estimated together with other

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consistent and converge in probability faster than $T^{1-\delta}$ for any positive δ, giving rise to the term super-consistent.

parameters in the system (e.g., Hansen, 1997). Note that the error correction term, $\mu^{\wedge}$, is the residual series from the cointegration equation (4).

GHH first test for the existence of a unit root in each of the price series (retail, wholesale and farm prices) employing the conventional Dickey-Fuller test. The authors fail to reject the unit root null for each of the series. Further, employing the maximum likelihood procedure of Johansen (1988), the authors conclude that the price series are cointegrated. Invoking Engle and Granger, the authors then estimate the cointegration equation (4), where y now is the retail price and x is a vector containing wholesale and farm prices. The estimated β is the long run price transmission coefficient. The residuals are then used in the estimation of a short run dynamic process similar to the one in equations (6a) and (6b), from which impulse response functions are simulated to aid interpretation of the dynamic interrelationships among the prices at alternative market levels. The authors find that shocks appear to be largely transmitted from farm to wholesale to retail markets but not in the opposite direction. In addition, they find that the price responsiveness to shocks has increased and become less asymmetrical in recent years.

The approach of current research

Building on GHH, this research adopts Perron's (1997) procedure of testing for unit root versus a breaking trend, to improve the power of the unit root test. Wang and Tomek (2007) question the frequent findings of nonstationarity in agricultural prices, contending that price theory suggests otherwise, and conduct unit root tests for several agricultural commodity prices allowing for an exogenous break date. Based on their empirical results, the authors caution analyst to have a healthy skepticism toward the unit root results of previous studies based on conventional unit root test. Unlike Wang and Tomek, the current study estimates the break date endogenously. We also use Phillips and Loretan's procedure to estimate the long term price linkage equation, to obtain an asymptotically unbiased long run price transmission coefficient. In addition, we entertain the possibility of a structure break of unknown date in the long run price linkage equation using the procedure by Bai (1997) and Bai and Perron (1998). Finally, we use Hendry, Pagan and Sargan's (1984) autoregressive distributed lag (ADL) model to fashion the short term price dynamics, which is applicable for both stationary and nonstationary data, and include GHH's error correction specification as a special case. The ADL model is useful in the current research as, unlike the affirmative unit root conclusion of GHH, the results from Perron's (1997) unit root test are mixed.

An important implication of Nelson and Plosser's revolution is that the unit root hypothesis stipulates that fluctuations are not transitory because random shocks have a permanent effect. After Nelson and Plosser, a flurry of empirical research has confirmed the prevalence of unit root in many interesting economic time series, eliciting skepticisms about the power of the conventional unit root tests in discriminating between a time series process driven stochastically by a unit root and that driven by a deterministic trend function with structural breaks. In Perron's (1997) breaking trend function hypothesis, the author seeks to establish that economic time series are predominately trend-stationary if one allows a change in the intercept and/or the slope of the trend function at an exogenously specified break date (Perron, 1989) or at an endogenously estimated break date (Perron, 1997). Upon allowing for structural breaks in the deterministic trend function, Perron is able to overturn many of the unit root conclusions in Nelson and Plosser. We subscribe to Perron's modeling philosophy that it is important to allow for structural breaks when conducting unit root tests so that analysts would not confuse a structural break in the deterministic trend function with a unit root.

The current study adopts the three models put forth by Perron (1997), which allow for endogenous estimations of a single break date both in the stochastic trend (i.e., unit root) null hypothesis and in the deterministic trend alternative. In Perron's first model, a shock in the constant term in both the unit root null and the breaking trend alternative specifications are specified. In his second model, the shocks pertain to both the constant and the slope parameters, while in the third model only slope parameters are allowed to change. For the first two models, Perron nests the null and the alternative specifications and includes lagged terms of the first-difference of the variable to account for his assumption that the structural change occurs gradually. These gradual change models are termed as the innovational outlier models, which are discussed in great detail in Perron's 1989 article. For the third model, the slope change is assumed to occur instantaneously and hence both segments of the trend function are joined at the time of break.[2] In this so-called 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) estimating the alternative model to remove the trend from the series and (ii) applying the conventional Dickey-Fuller test to the de-trended series to ascertain the existence of a unit root. For each

2. Perron (1989) discusses the problem associated with this model if one assumes that the structural change occurs gradually, as in the first two models.

specified break date, the model is linear and can be estimated by ordinary least squares method. In terms of break date estimation, one method proposed by Perron (1997) is to choose the date such that the t statistic associated with the null hypothesis of unit root (i.e., the coefficient associated with the lagged dependent variable) is as small as possible.[3] See Appendix A for details on Perron's three models.

Our second contribution to the literature beyond GHH is the estimation of the long run price linkage equation in (4). As mentioned, the OLS estimators of the parameters in the cointegration equation (4) are asymptotically biased and non-normal, albeit super-consistent. 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 coefficients. The biasness and the non-normality of the cointegration parameter are problematic in the current context of price transmission analysis, because the magnitude of the long term price transmission coefficient β is of intrinsic interest. One would like to be able to obtain an unbiased estimate of the transmission coefficient and test if the coefficient is indeed equal to a certain value, such as unity in the case of perfect price transmission hypothesis. Following Liu, Margaritis and Tourani-Rad (2007), the current study addresses the above mentioned problem by adopting Phillips and Loretan's (PL) approach.

Phillip and Loretan consider the following triangular system of equations that encompasses the static cointegration equation in (4):

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3. This method maximizes the power of the unit root test. Note that the t statistic of α=1 in the regression of $y_t = \alpha\, y_{t-1} + \epsilon_t$ is a negative number.

$$(7a) \quad y_t = a + b\,x_t + m_{1t}$$
$$(7b) \quad x_t = x_{t-1} + m_{2t}$$

where μ_{1t} and μ_{2t} are stationary processes. Equation (7b) explicitly recognize that x_t is nonstationary while the assumption that μ_{1t} is stationary acknowledges that x and y are cointegrated. In Phillips and Loretan's specification of PL(M,N,Q) the cointegration equation in (7a) is augmented by M leading and N lagged terms of Δx_t, and Q terms of the lagged residuals in the following format:

$$(8) \quad y_t = a + b\ x_t + \sum_{i=-M}^{N} V_i\ \Delta x_{t-i} + \sum_{i=1}^{Q} x_i\ (y_{t-i} - a - b\ x_{t-i}) + e_t\,.$$

The inclusion of the lead and lag terms of Δx_t can be thought of as projecting μ_{1t} in (7a) against the leads and lags of μ_{2t} in (7b), rendering the error e_t in (8) uncorrelated with μ_{2t}. The inclusion of the lagged error correction terms $(y_{t-i} - \alpha - \beta\ x_{t-i})$ removes potential autocorrelation in e_t. Phillips and Loretan propose estimating (8) by non-linear least squares method, and show that the estimated long run parameters are asymptotically unbiased and normal and performs well in finite samples.

Work Plan

Below, we discuss what has been estimated and what work still needs to be completed. This paper will be updated prior to the AAEA annual conference, and will included detailed results.

<u>Completed tasks</u>

Using weekly and monthly data sets, we conduct various versions of the conventional Dickey-Fuller tests and Perron's unit root tests which allows for structural break. Unlike GHH, which found unit root for all price series, our results are mixed with one-half of the cases suggesting stationarity and the other half confirming unit root. This set of tests is currently under refinement. Second, we have conducted Granger causality tests on the price data. The preliminary results generally confirm the findings of previous studies that the causality runs from farm to wholesale to retail. Third we estimate the long run price linkage equation using Phillips and Loretan's method. In general, we find that the lagged residual terms are important to address the problem of autocorrelation, and in some cases the estimated long run price transmission coefficient is sensitive to the lag length specification of the lagged residual terms. This set of estimation is currently under refinement.

Work to be completed

To entertain structure change in the long run price linkage equation, we need to codify the procedure of Bai and Perron and determine the Monte Carlo simulation procedure to obtain critical values for inferences on the estimated parameters. After the long run equations are estimated, we need to estimate the short term dynamics, within the autoregressive distributed lag model of Hendry, and conduct policy simulations to assess the dynamic interrelationships among the prices.

Appendix A: Perron's three models of testing for unit root (DRAFT)

Model 1: Entertain a gradual change in the intercept at the break date τ under both the unit root null and the breaking trend alternative (Innovational Outlier Model):

$$
\begin{aligned}
&(\textit{null}) \quad y_t \quad = m + a \ y_{t-1} + d \ D\_TEMP_t + e_t, \ \text{ with } a = 1 \\
&(\textit{alternative}) \quad y_t \quad = m + b \ t + (m_2 - m) \ D\_PERM_t + e_t \\
&(\textit{composite}) \quad y_t \quad = m + a \ y_{t-1} + b \ t + d \ D\_TEMP_t + q \ D\_PERM_t + \sum_{i=1}^{L} c_i \, \Delta y_{t-i} + e_t
\end{aligned}
$$

where $\theta \equiv \mu_2 - \mu$, $D\_TEMP_t = 1$ for $t = \tau + 1$ and zero otherwise, $D\_PERM_t = 1$ for $t > \tau$ and zero otherwise. 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 incorporate the assumption that the structural change is of a gradual type (Perron 1989).

Model 2: Entertain a gradual change in both the intercept and the slope at the break date τ under both the unit root null and the breaking trend alternative (Innovational Outlier Model):

$$
\begin{aligned}
&(\textit{null}) \quad y_t \quad = m + a \ y_{t-1} + d \ D\_TEMP_t + (m_2 - m) \ D\_PERM_t + e_t, \ \text{ with } a = 1 \\
&(\textit{alternative}) \quad y_t \quad = m + b \ t + (m_2 - m) \ D\_PERM_t + (b_2 - b) \ D\_TREND_t + e_t \\
&(\textit{composite}) \quad y_t \quad = m + a \ y_{t-1} + b \ t + d \ D\_TEMP_t + q \ D\_PERM_t \\
&\qquad\qquad\qquad + g \ D\_TREND_t + \sum_{i=1}^{L} c_i \, \Delta y_{t-i} + e_t
\end{aligned}
$$

where $\gamma \equiv \beta_2 - \beta$, and D_TREND = t for $t > \tau$ and zero otherwise. 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 incorporate the assumption that the structural change is of a gradual type (Perron 1989).

Model 3: Entertain a sudden change in the slope at the break date τ under both the unit root null and the breaking trend alternative (Additive Outlier Model):

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

where D_R_TREND = t - τ for t > τ and zero otherwise. Under this model, a change in the slope is allowed but both segments of the trend function are joined at the time of break. Thus, the change is presumed to occur rapidly and corresponds to the "additive outlier model" in the literature. 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.

$$(\text{Step 1}) \quad y_t = m + b\, t + g\, D\_R\_TREND_t + y_t^{\sim}$$
$$(\text{Step 2}) \quad y_t^{\sim} = a\, h_{t-1} + \sum_{i=1}^{L} c_i\, \Delta n_{t-i} + e_t$$

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