Retailers' response to wholesale price changes: New evidence from scanner-based quantity-weighted beef prices

Retailers' Response to Wholesale Price Changes: New Evidence from Scanner-Based Quantity-Weighted Beef Prices

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ABSTRACT

Retailers are often criticized for a slow response in retail prices to changes in wholesale beef prices. The unresponsiveness, especially when wholesale and farm prices are declining, is seen as a lack of competitiveness and as a reason for congressional action to regulate behavior of processors or retailers. The validity of the historical analyses of retailer price responsiveness is questionable, however. Traditional Bureau of Labor Statistics (BLS) retail beef prices biased upward and do not account for large volumes sold at discounted prices. This article uses newly available scanner-based quantity-weighted retail prices to suggest that retailers' response to changes in wholesale beef prices is significantly larger and possibly quicker than is shown by traditional BLS measures of retail prices. Recent efforts to prompt legislation to regulate how firms behave along the beef supply chain, which are based only on arguments that retailers are not responding to price changes at the wholesale level, may be inappropriate. [L110, L660, D400]

1. INTRODUCTION

Upward biases in the Bureau of Labor Statistics (BLS) retail price and index series have been well documented (e.g., on the Consumer Price Index [CPI], Hausman, 2003 and Moulton, 1996; on food prices, Nakamura, 1998). One source of the bias results from the failure of BLS price series to account for consumers' increased consumption of reduced price goods and reduced consumption of more expensive goods. Hausman and Hausman and Leibtag (2005) suggest an approach that employs both price and quantity data to address biases in price indices. Broadly speaking, this approach amounts to constructing price series or price indices by using quantities as weights. The increasing availability of weekly data collected by scanner devices in retail stores makes the construction of quantity-weighted price series feasible.

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Previous research has noted the implications of price measurement on productivity, Gross Domestic Product (GDP) measures, and on cost-of-living measures, such as the CPI. We use newly available quantity-weighted retail prices to extend this body of knowledge and to suggest that the often cited failure of retail prices to respond to changes in wholesale- and farm-level prices, which has been based on traditional BLS retail prices, may be inaccurate.

Because of the important policy issues involved, the interest in meat prices and price behavior is long standing. Prices at the farm, wholesale, and retail levels do not always move in parallel. When farm-to-retail price spreads increase, because farm-level prices decline and retail prices do not, livestock producer groups cite the increases in price spreads and the related declines in farmer share of the consumer's dollar as evidence of a lack of competitiveness among middlemen along the beef supply chain and of the need for Congressional intervention on behalf of livestock producers. This concern is evident in testimony by USDA Chief Economist Keith Collins during a hearing by the Committee on Agriculture, U.S. House of Representatives, February 10, 1999:

> We know that every time farm prices have dropped sharply, retail prices have not followed the drop down. (There is) a built in time lag … we have called the lack of time lags competitive in other industries, and we are concerned about it in this (livestock) industry. And we need more work on that.

Producers are particularly interested in retailers' behavior when beef production has surged and lower prices at retail are needed to pull the increased volume through the supply pipeline and into consumption. They are critical when they do not see prompt price declines at retail. But the historical monthly retail price series for beef collected by the BLS, on which such criticism is based, consists of a simple average of beef prices sampled during the month in retail outlets across the country. BLS retail prices do not account for the large portion of retail sales that usually occurs during periods when retail prices are at discounted levels. BLS series are not, therefore, accurate measures of the prices to which consumers are responding in their buying decisions.

The result of using an inaccurate measure of retail prices and price changes could be more criticism from producers than is merited and an exaggerated tendency to push for congressional policies that place regulations on how business can be done along the supply chain. An example of producer-level action is Senate Amendment No. 2534 to the Agricultural, Conservation, and Rural Enhancement Act of 2001 (Farm Bill), which made it illegal for nonproducers along the supply chain to own, feed, or control livestock more than 14 days prior to slaughter. The legislation, which passed the Senate but failed in the House, would have blocked many of the modern contractual arrangements being used to bring coordination of effort and quality control to the beef supply chain. Regulations always have both predictable and unpredictable consequences and can therefore have both positive and negative unanticipated impacts on the producer and consumer. It can be argued, for example, that contract arrangements and even vertical integration were necessary conditions for quality control that brought processor-level expenditures to improve product offerings to the benefit of both producer and consumer. When such important considerations are involved, any tendency toward increased regulation becomes an important policy issue.

The objective of our article is to investigate the hypothesis that the *actual* reaction of retailers to changes in wholesale and farm prices is different than that implied by Keith Collins' congressional hearing testimony. In particular, we make use of new quantity-weighted prices to argue that retailers are more responsive to changes in wholesale prices than what BLS retail prices suggest. The new series consist of weekly scanner prices that are weighted by the sales volume registered at the price during that week and are then aggregated into a quantity-weighted monthly price series. These retail series have been developed by the Economic Research Service (ERS) in the USDA in response to the longstanding criticisms that the BLS prices are incorrect and biased and the related congressional directive for a better and more transparent price system consistent via the Livestock Mandatory Reporting Act of 1999.[1]

We employ a cointegration approach that takes into account the relationship among farm, wholesale, and retail beef prices. The cointegration results are used to investigate whether the ERS scanner-based monthly retail price series reacts differently to changes in wholesale prices (or boxed beef) prices[2] than does the retail price series based on the historical monthly BLS price data. In general terms, we find that retailers' responses to changes in wholesale prices are larger when using ERS scanner price data. Our results suggest that using a "lack of responsiveness" argument based on analysis of BLS data to justify concerns about lack of competitiveness or to support calls for regulations on how retailers and processors can interact and do business may, therefore, be inappropriate.

We do not suggest that the perceived failures of retail prices to respond quickly to changes in wholesale- and farm-level prices are the only reasons for proposed legislative solutions via regulation of highly concentrated markets, nor do we suggest that retailers should always respond immediately to changes at the wholesale levels. Indeed, in some circumstances, there are legitimate reasons for the retailer to be cautious in changing prices. Changing price labels and price bar codes is costly, and the retailer will never want to decrease prices significantly if recent wholesale- and farm-level price declines are likely to be temporary.

Furthermore, theory suggests that it is a priori unclear whether a more competitive retail market (as measured by concentration) should transmit, or pass through, a larger magnitude of input cost increases (in this case wholesale prices) than a less competitive market. Cotterill, Egan, and Buckhold (2001) show that the relationship between pass-through rates and competition can be either positive or negative depending on the assumed demand functional form: Under a linear demand functional form, pass-through rates increase with competition, whereas under a log-linear demand functional form, pass-through rates decrease with competition.[3]

Any government intervention in this industry should thus be based on a more exhaustive analysis that determines: (a) the market conditions (demand and supply) of the retail sector and the expected price transmission from wholesale to retail prices (how the market works), and (b) a better understanding of how the price

[1]Bill number H.R 1906 available at http://thomas.loc.gov/cgi-bin/thomas.

[2]We explain in more detail later why we focus on wholesale and retail prices and not on farmer prices.

[3]Assumptions of this model are homogeneity of products and constant marginal cost.

transmission behaves (what the market looks like). Our contribution to the base of information on which policy and market regulation decisions will be made is to take advantage of the new retail price information to clarify the behavior of price transmissions at the wholesale and retail interface.

We have chosen the beef market for several reasons. The beef processing sector consolidated rapidly during the 1985–1995 period, with the CR-4 concentration index increasing from near 40 in the mid-1980s to above 80 during the early 1990s and into 2006. The retail food sector was consolidating at the same time, so scrutiny of this sector by public agencies was to be expected. The recent availability of scanner-based data and the longstanding criticism by farmers regarding retailers' lack of responsiveness added other reasons to focus on beef. But our results have implications in other industries because the inaccuracy in the measurement of retail prices is likely to occur in many markets (MacDonald, 1995; Nakamura, 1998).

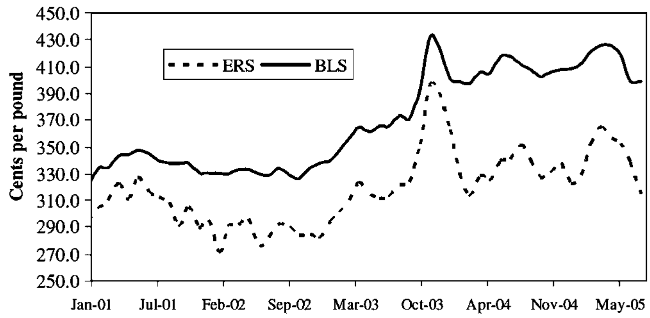
2. DATA

Monthly price data are used for this study. The data include prices for beef at three market levels: farm, wholesale, and the retail level. There are two retail series: the historical BLS and the new ERS retail monthly series. Farm (live cattle) and wholesale (boxed beef) prices are available weekly, but the historical BLS retail series is only available as a monthly series. *Farm price* is the monthly average Western Kansas fed steer price, 11–1300 lb, 35–65% Choice grade, measured in cents per pound. *Wholesale price* is the monthly Choice boxed beef cutout value in cents per pound for 750–900 lb boxes. Both wholesale and farm prices are adjusted to their retail dollar equivalents using adjustment factors employed by ERS in calculating published price spreads. The *BLS retail price* is the traditional simple average monthly retail price for Choice beef collected by the BLS in cents per pound.

ERS retail price is the monthly retail composite price in cents per pound, a quantity-weighted monthly price series constructed from weekly prices weighted by the weekly sales volumes. Prices and quantities are read from scanner tapes in supermarkets and acquired by commercial data firms. Then, a third party (outside USDA) processes the data and delivers aggregated information to ERS so that the identity of the stores or chains is not revealed. Although not a representative sample, supermarkets included in the sample are those with sales over $2 million and together account for approximately 20% of all supermarket sales in the U.S. More details on ERS scanner retail prices are available at: http://www.retail-lmic.info/CD/questions.htm.

The ERS retail scanner-based prices start on 2002, but ERS provided a special calculation of a quantity-weighted monthly price series that was extended back to January, 2001 and is available through August, 2005. Unfortunately, the Federal Mandatory Price Reporting Act that enforced and ensured funds for the calculation of these data expired on October, 1, 2005, and ERS retail prices are not available after August of 2005.

There are apparent differences between the two retail series as can be seen in Figure 1. The mean of BLS Choice prices is 369.93 cents per pound and the mean of ERS prices for Choice beef is 318.58 cents per pound. Thus, on average, BLS



Source: Economic Research Service, USDA.

Figure 1 Monthly BLS and ERS beef retail prices in nominal dollars.

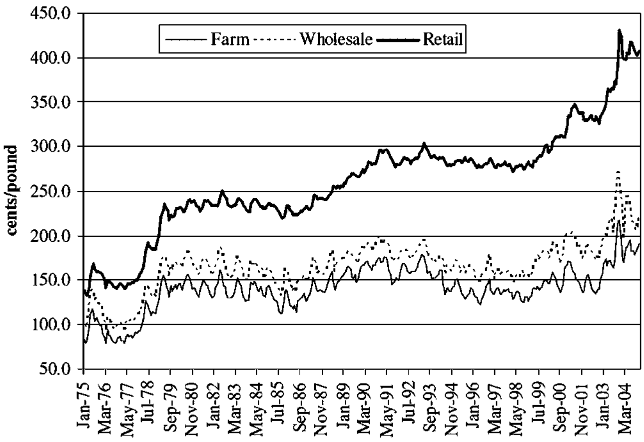
prices are 16% higher than the ERS prices. A parametric test of equality of means and a nonparametric test of equality of medians confirm that BLS prices are higher than ERS prices (p-value <0.01 in both tests). BLS prices range from 321.40 to 431.70 cents per pound while ERS prices range from 270.50 to 395.90 cents per pound, suggesting that ERS prices are more variable than BLS prices. A test of equality of variances confirms this (p-value <0.05). Furthermore, nonparametric tests reject the null hypothesis of equality of distributions (p-value <0.01 for Kruskal-Wallis and Wilcoxon rank-sum tests). This preliminary analysis of the two retail series indicates that there are significant differences between the two.

The (somewhat weak) evidence of a larger volatility of the ERS series is arguably a necessary condition if we are to be able to confirm through empirical analysis our hypothesis that the ERS series reacts differently to changes in wholesale prices. A more sophisticated econometric analysis is needed to provide a rigorous test of the hypothesis. The retail-wholesale relationship receives most of the attention in the analysis because it is at retail where most of the "lack of responsiveness" criticism has been focused. Because boxed beef and farm-level prices move in tandem and are highly correlated (correlation = 0.95, p-value <0.01), we interpret any lack of response to wholesale price changes as measured by boxed beef values also as a lack of response to live cattle prices.

3. ECONOMETRIC ANALYSIS

Visual inspection of time series (Figure 2) plots shows that movements in farm, wholesale and retail prices (at the retail weight equivalent) are related and tend to move generally together over time. Therefore, cointegration emerges as a plausible econometric framework for the analysis of the relationship among variables.

If a group of variables is cointegrated, there exists a long-run relationship or equilibrium across the variables. In econometric terms, this means that a linear combination of the variables is stationary. Thus, if all variables are in equilibrium at



Source: Bureau of Labor Statistics (BLS) series obtained from Economic Research Service at the USDA.

Figure 2 Monthly farm, wholesale, and retail prices in nominal dollars for beef at retail weight equivalent (January 1975 to November 2004).

time 0 and there is a change in one of them at time 1, then all variables will move so as to restore the long-run equilibrium.[4]

In practice, there are several factors affecting the change of the variables back to the long-run equilibrium. Specifically, there are different rates of adjustment of each of the variables and possibly more than one cointegrating vector. Also, there may be other explanatory variables that enter the analysis and, frequently, there are more than two series in the cointegrating vector. Finally, short-run dynamics play a role in this adjustment (Johansen, 2005). A tool that can capture all these factors is "impulse-response analysis" (Luktepohl & Reimers, 1992). Essentially, this technique incorporates all relevant information from the econometric analysis of the cointegrated series and allows plotting the reaction of one variable when another disturbs the long-run relationship.

To test the hypothesis of whether ERS scanner prices are more responsive to changes in wholesale prices than the traditional BLS prices, two cointegration models are estimated: one containing ERS quantity-weighted monthly average prices and the other containing the traditional simple average monthly BLS prices. Then, a comparison of the responses of the two measures of retail prices to changes in the same wholesale price series is conducted using impulse-response analysis.

3.1. Cointegration

Cointegration is related to two other concepts in time series analysis: (1) vector autoregressive (VAR) models and (2) vector error correction models (VECM).

[4]Suppose that the farm price is cointegrated with the wholesale price with a (1,–1) cointegrating vector. The cointegrating vector indicates the coefficients of the linear combination, that is PriceFarm – PriceWholesale = 0. If the farm price increases by one unit and "destabilizes" the linear combination by making PriceFarm-PriceWholesale = 1, cointegration implies that the wholesale price must increase by a unit to restore the relationship. This example considers a one-to-one long-run relationship, which need not be the case.

In this study, a VAR model of lag p, VAR(p), would be written as

$$y_t = A_0 + A_1 y_{t-1} + \ldots + A_p y_{t-p} + e_t \quad (1)$$

where

- $y_t =$ a 3×1 vector (y_{1t}, y_{2t}, y_{3t}) and $y_{1t} =$ farm price, $y_{2t} =$ wholesale price, $y_{3t} =$ retail price,
- $A_i =$ a 3×3 matrix of coefficients, for $i = 1, \ldots, p$, and
- $e_t =$ a 3×1 vector (e_{1t}, e_{2t}, e_{3t}) of white-noise error terms.

To allow for nonstationary series, as is the case here, the VAR model is reparameterized into a VECM, as follows:

$$\Delta y_t = A_0 + \Pi y_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta y_{t-i} + e_t \quad (2)$$

With an appropriate transformation of the coefficients Π and Γ_i, Equation 2 is equivalent to Equation 1. The term Πy_{t-1} in Equation 2 determines whether the three variables are cointegrated. If all variables are individually integrated of order one and there are no cointegrating relations, then the matrix of coefficients Π is equal to zero. If there are r cointegrating relations and stationary linear combinations of the three variables, then[5]

$$\begin{aligned} z_{1t} &= b_{11} y_{1t} + b_{12} y_{2t} + b_{13} y_{3t} \\ &\vdots \\ z_{rt} &= b_{r1} y_{1t} + b_{r2} y_{2t} + b_{r3} y_{3t} \end{aligned} \quad (3)$$

Equation 2 is then expressed as

$$\Delta y_t = A_0 + \alpha z_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta y_{t-i} + e_t \quad (4)$$

where α is a $3 \times r$ matrix of adjustment or "loading" coefficients and Z_{t-1} is a $r \times 1$ vector of cointegrating relations. With one cointegrating relation ($r = 1$), α is a 3×1 matrix and its three elements $\alpha_1, \alpha_2, \alpha_3$, measure how each of the three variables adjust to the long-run equilibrium ($z_t = 0$) when there is a shock or disturbance in the system $z_{t-1} \neq 0$.[6] Any shock to the linear combination is only temporary because the variables are cointegrated and should return to the long-run equilibrium. The α coefficients characterize the long-run dynamics of the system, while the Γ coefficients establish the short-run dynamics. A large (small) value of α implies that the variable will respond rapidly (slowly) to the deviations from the long-run equilibrium.

Using Equations 3 and 4, the matrix Π is written as $\Pi = \alpha b$, where b is the $r \times 3$ matrix of coefficients given in Equation 3. Johansen (1988) provides a way to test the number of cointegrating relations and to recover the corresponding α and b

[5]The maximum number of r is n-1, where n is the number of variables in the system, in this case $n = 3$.

[6]In this example, z_{rt} is equal to zero; more generally, it can be any constant.

coefficients. This procedure amounts to testing the rank of the matrix Π, which determines the number of cointegrating relations. This is done by calculating the eigenvalues of Π, ordering them from largest to smallest, and testing which ones are significantly different from zero (i.e., p-value smaller than 5%). With only the largest eigenvalue significantly different from zero, then only one cointegrating relation exists; if two largest eigenvalues are significantly greater than zero, then two cointegrating relations exist, and so on.

Technically, the analysis above assumes that the retail price responses to increases and decreases in wholesale prices are symmetric. Therefore, we test for the existence of asymmetric responses where retail prices react differently to wholesale price increases than to wholesale price decreases. We employ the two-step procedure introduced by Granger and Lee (1989) and estimate the following nonsymmetric error correction equation:

$$\Delta y_{rt} = B_0 + \sum_{i=1}^{p-1} \alpha_i \Delta y_{wt-i} + \sum_{i=1}^{p-1} \beta_i \Delta y_{ft-i} + \delta_1 \xi_{t-1} + \delta_2 \xi_{t-1}^{+} + \varepsilon_t \tag{5}$$

where the subscripts r, w, f denote retail, wholesale and farm prices, respectively; $\alpha_i' s$, $\beta_i' s$, and $\delta_i' s$ are coefficients to be estimated, ξ_t is the error correction term from a first stage OLS regression (in levels) of the retail price on wholesale and farm prices, and $\xi_{t-1}^{+} = \max(\xi_{t-1}, 0)$. The speed of adjustment to the long-run equilibrium is captured by $\delta_1 < 0$, and the null hypothesis of symmetry is tested through δ_2 where a statistically insignificant coefficient is an indication of a symmetric response to price increases and decreases. If, on the other hand, as noted in some research literature (see next section), retail prices respond faster to wholesale price increases than to wholesale price decreases, the δ_2 coefficient should be significantly positive.

4. RESULTS[7]

The augmented Dickey-Fuller test is used to test the null hypothesis of nonstationarity in levels and in differences. When the variables are in levels, the null hypothesis of a unit root for the four variables cannot be rejected at the 5% significance level. For the differenced variables, the null hypothesis of a unit root is rejected for all series at a (stronger) 1% significance level.[8] Therefore, all variables are I(1) as required for cointegration.[9]

The number of lags to be included in Equation 4 is determined with the use of lag selection criteria in a VAR. These criteria make pair-wise comparisons between two VAR models, each with a different number of lags.[10] The literature includes a variety

[7]The econometric analysis was performed using the JMulti software developed by Luktepohl and Kratzig (2004).

[8]Results of these tests are robust to different specifications. Unit root tests in levels were conducted with and without a trend. For the differenced variables, the tests were done with and without a constant.

[9]Higher orders of cointegration were tested and rejected.

[10]The model with less lags (restricted model) is rejected if additional lags of the unrestricted model bring in additional explanatory power. This is done up to the point where the restricted model can no longer be rejected.

of criteria. However, according to Liew (2004), Aikaike's information criterion (AIC) and the final prediction error (FPE) are the best criteria when using small samples (60 or less observations). These two criteria are used as this study employs 56 observations. Table 1 reports the results of the lag selection procedure for each of the two models considered: Farm-Wholesale-BLS Retail (the BLS model) and Farm-Wholesale-ERS Retail (the ERS model). Both models have the same farm and wholesales prices but differ in the retail price. The AIC and FPE results suggest that the number of lags is 4 when using BLS prices and 2 when using ERS prices.

The trace test of the Johansen procedure is reported in Table 2. If the p-value of the likelihood ratio (LR) statistic for the *n*th largest eigenvalue is smaller than the critical level (here we adopt the usual 0.05 level), then the null hypothesis that the rank of the Π matrix is n is rejected and the test proceeds to the next largest eigenvalue. The results indicate that there is one cointegrating relation in each of the two models considered.

Table 3 reports the adjustment coefficients (the α's in Equation 4) and the cointegrating vectors (the *b's* in Equation 3) for both models with their respective t-statistics reported in parentheses. The farm price is normalized in the cointegrating equation.

The corresponding cointegrating equations are

$$P_{\text{farm}} - 0.395P_{\text{whole}} - 0.361P_{\text{BLS}} = 0 \tag{6}$$

$$P_{\text{farm}} - 4.293P_{\text{whole}} + 2.737P_{\text{ERS}} = 0 \tag{7}$$

Beta coefficients in the cointegrating Equation 7 are an order of magnitude larger than those in Equation 6, which would suggest that long-run adjustments to

TABLE 1. Results of Lag Selection Criteria

	Number of lags	
Test criterion	BLS	ERS
Akaike info criterion	4	2
Final prediction error	4	2
Hannan-Quinn criterion	4	1
Schwarz criterion	1	1

TABLE 2. Results of Trace Tests on Rank of Π Matrix in the Vector Error Correction Model

		BLS model		ERS model	
Null hypothesis	Eigenvalue (largest to smallest)	LR	p-value	LR	p-value
Rank 0	1	48.37	0.0008	54.99	0.000
Rank 1	2	19.67	0.0588	19.16	0.069
Rank 2	3	2.55	0.6716	3.30	0.5359

TABLE 3. Estimated Adjustment Coefficients (α's) and Cointegrating Coefficients (β's)

	BLS*			ERS*		
	Farm	Wholesale	Retail	Farm	Wholesale	Retail
Alphas	−0.231	0.442	−0.492	0.015	0.107	−0.194
	(−1.079)	(1.670)	(−2.817)	(0.523)	(2.926)	(−5.276)
Betas	1.00	−0.395	−0.361	1.00	−4.293	−2.737
	(—)	(−3.406)	(−5.499)	(—)	(−8.392)	(−6.012)

*t-statistics in parenthesis.

disequilibrium errors through retail prices might be larger in the ERS model than in the BLS model. However, Lutkepohl and Reimers (1992) and Johansen (2005) identify and stress the importance of the short-run dynamics of the system (the $\Gamma_i' s$ in Equation 4) and the adjustment coefficients (alphas) in determining the adjustment of series to disequilibrium errors. Because the short-term dynamics and the adjustment coefficients are different between the two models, an analysis that combines both elements is needed to determine which of the two series adjusts more rapidly to disequilibrium errors.

Impulse-response analysis is utilized to analyze the response of one variable to the change of another while capturing all factors governing the dynamics of the system (Lutkepohl & Reimers, 1992). To allow and correct for potential contemporaneous correlation between each of the error terms in Equation 4, an "orthogonalized" variant of the impulse-response analysis is used (Hamilton, 1994, chap. 11). In this case, the instantaneous or same period response of one variable to a change in another depends on the ordering of the series. We adopt a "natural" ordering of price transmission (from upstream to downstream) in the impulse-response analysis. Indeed, previous research (e.g., Goodwin & Holt, 1999) has found that price transmission occurs from farm prices to wholesale prices and from wholesale price to retail prices. With this ordering, a change in the farm price is allowed to have contemporaneous effect on wholesale and retail prices, while a change in wholesale price is allowed to have only a contemporaneous effect on retail prices. Changes in retail prices are allowed to have only a lagged effect on the other two variables.

Figures 3 and 4 display the 20-month responses of BLS prices and ERS prices to a shock equal to one standard deviation in the wholesale price of the "orthogonalized" model.[11] The response of BLS prices is not statistically different from zero in any period (the lower band in the confidence interval is always below zero and the upper band always above zero). The responses of ERS prices are, on the other hand, almost always statistically different from zero at the 5% significance level (the lower confidence band is typically above zero and the upper band always above zero) and never approach zero. A complete return to zero is not expected: With cointegrated series, a shock in one of the variables (in this case the wholesale price) causes a

[11]The shock is equal to one standard deviation in the "orthogonalized" wholesale error term.

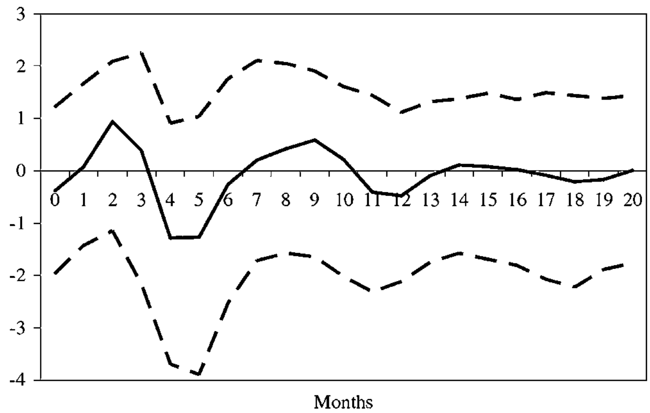


Figure 3 Response of BLS prices to an impulse of one standard deviation of the orthogonalized error of wholesale prices, cents per pound (- - - 95% confidence intervals).

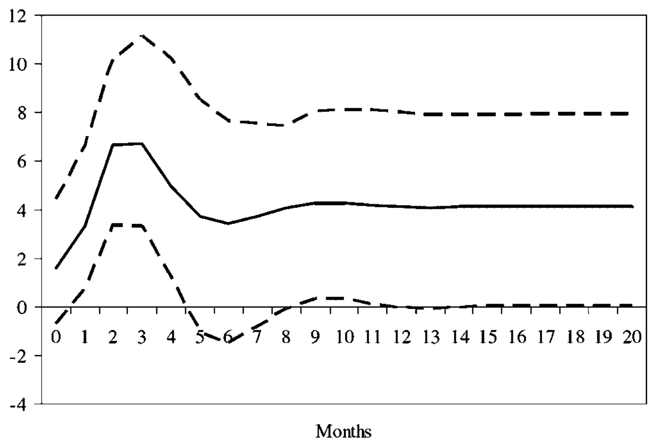


Figure 4 Response of ERS (scanner) prices to an impulse of one standard deviation of the orthogonalized error in wholesale prices, cents per pound (- - - 95% confidence intervals).

movement in all series that eventually becomes permanent because of their nonstationary nature.

In addition to being almost always statistically significant, ERS prices appear to respond quicker and by a larger amount: BLS and ERS responses are, respectively, −0.39 and 1.57 at time 0, 0.07 and 3.31 at time 1, 0.94 and 6.68 at time 2, and 0.39 and 6.72 at time 3. The difference in the magnitude of response between the two series is illustrated in Figure 5, which plots both ERS and BLS responses (confidence intervals omitted for clarity). The impulse-response functions in Figures 3 through 5 suggest that the ERS series adjusts to changes in wholesale prices, while the BLS series does not.[12] Although weakly significant (15% level), the ERS response is

[12]The statistically significant response of ERS prices to wholesale price changes is consistent with the betas of the ERS model being an order of magnitude larger than those of the BLS model.

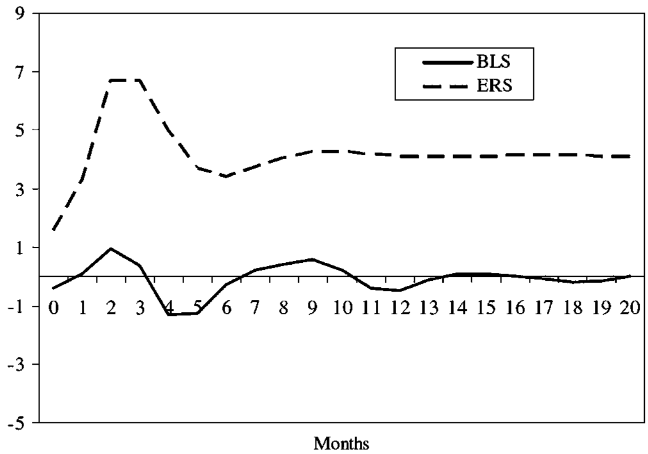


Figure 5 Response of ERS and BLS prices to an impulse of one standard deviation of the orthogonalized error in wholesale prices, cents per pound.

positive in the contemporaneous period (month 0), and remains significantly above zero for later periods. ERS prices adjust quickly to changes in wholesale prices and the magnitude of the cumulative response would suggest a more effective price discovery.[13]

As noted earlier, the econometric analysis assumes symmetric responses of retail beef prices to changes in wholesale prices. There is evidence that suggests the existence of asymmetric responses of retail prices to changes in wholesale prices when BLS retail prices are used. This asymmetry, when identified by researchers, typically takes the form of retail prices responding faster and in greater magnitude to wholesale price increases than to wholesale price decreases (Goodwin & Holt, 1999; Hahn, 2004). We conducted asymmetry tests in both models using Equation 5 and failed to reject the null hypothesis of symmetric responses in either of the two retail series: for ERS series $\delta_2 = 0.11$ (p-value = 0.64) and for BLS series $\delta_2 = -0.23$ (p-value = 0.57). These results ensure that the symmetry assumption is not violated and that our results are not biased.

An explanation for the lack of asymmetry in the BLS series, which is at odds with previous findings in the cited references, may be due to the short nature of the analysis period that was limited to January, 2001–August, 2005. However, reflection on what has been learned about the differences between the BLS and ERS price series suggests that historical analyses of asymmetry may themselves be biased. The findings based on the ERS data pose the interesting possibility that analysis of historical BLS data have significantly underestimated the magnitude of retailers' decreases in retail prices in response to decreases in wholesale prices. The failure of BLS data to account for the impact of price specials would suggest lack of proper measurement of retailers' responses to decreases in wholesale prices. The new

[13]In addition, the parameters that measure the speed adjustment in Equation 5—for ERS series (p-value = 0.16) and for BLS series (p-value = 0.57)—suggest that ERS series adjust faster because its p-value is statistically significant at a (weak) 20% level, whereas BLS series are far from being significantly different from zero.

scanner data show that for many cuts of beef there exist essentially two prices during a year, a *regular* and a *discount* weekly price such as \$7.98 and \$5.98, respectively, for a cut like sirloin steak. ERS prices will differ sharply from the simple average BLS prices during periods (often several consecutive weeks based on observation of scanner data) that the special \$5.98 price is present. Quantities sold may surge by levels up to tenfold at the \$5.98 price and pull the quantity-weighted prices sharply lower. Retailers are likely to try to coordinate price specials with times of higher expected production and lower boxed beef and cattle prices. If BLS prices do, in fact, significantly understate the responsiveness of retailers to wholesale price declines, the resulting lack of accurate measurement is a logical reason for the asymmetry finding in analyses using historical BLS data. We leave this question for future research if current efforts to extend the mandatory price reporting legislation are successful and the ERS retail series are extended.

5. SUMMARY AND CONCLUSIONS

Biases in price series as calculated by the BLS have been widely documented at the aggregated level. We provide evidence that this bias is not only present at a more disaggregated level, but that measurement inaccuracy can have implications in other widely studied issues, such as retailers' pricing behavior and its relationship to congressional policymaking and considerations of market regulation.

When the farm-retail or wholesale-retail price spread shows an increase because boxed beef and cattle prices have declined and retail prices are stable or higher, producers get concerned. In such an environment, Congress is often being asked to step in on behalf of producers by regulating how business can be done along the supply chain, especially at the processing and retailing levels. But market regulation is a drastic step and should be considered only if it enhances welfare and if measures of responsiveness are accurate and demand and supply conditions are fully understood.

Research by Lensing and Purcell (2006) shows that historic retail prices as collected by BLS are in fact too high. Our efforts show, as producer groups often assert, that the BLS prices do not always respond to price declines at the wholesale level. But recent research by Lensing and Purcell indicates that the BLS measures of retail prices used in current and historical analyses are incorrect and do not represent what is actually occurring at the retail level. Lensing and Purcell hypothesized that when the new ERS retail price data are analyzed, retailers' responses to wholesale- and farm-level price declines will be larger and quicker than we have historically believed.

We considered two cointegrating models, one with the traditional BLS monthly retail prices and the other with the new quantity-weighted monthly ERS scanner prices. Because the same wholesale values enter both cointegration models, our findings about price responsiveness by retailers have direct implications for the wholesale-retail price spread and indirect yet very predictable implications to the highly correlated farm-retail price spreads.

When wholesale prices are increasing, ERS prices increase in greater magnitude (and perhaps quicker) than BLS retail prices. But when wholesale prices are decreasing, retail ERS scanner-based prices also decrease in greater magnitude than the BLS prices. As a result, when wholesale prices decline, the wholesale-retail price spread based on ERS prices decreases in greater magnitude than the wholesale-retail spread based on BLS prices. Our results suggest that Keith Collins' assertion that retailers are not "following the drop down" in wholesale prices might, therefore, be incorrect because the assertion is based on flawed retail price data.[14]

It appears that the base for the long-standing criticisms of retailers' failure to respond to wholesale- and farm-level price declines disappears when ERS quantity-weighted monthly average retail prices are used to measure retailer response and calculate wholesale-retail price spreads. Our results also indicate that asymmetric responses to wholesale price changes are not present in either the ERS or the BLS retail prices. While this validates our assumption of symmetric responses in the impulse-response analysis, we interpret it with some caution given the short time period available for analysis and the importance of sample size to tests for asymmetry.

There are two additional caveats that should be mentioned. First, ERS retail prices are not obtained from a representative sample of supermarkets. Second, as pointed out in the introduction, a higher rate of input price transmission, or pass-through, does not necessarily mean that the market is more competitive. Before suggesting intervention in this industry, authorities will not only have to take into account the "right" measure of retailers' responsiveness but also whether this measure of price transmission is in accordance with prevailing supply and demand conditions.

As data collection through supermarket scanners becomes more widespread, the construction of quantity-weighted price series becomes more nearly feasible and available to researchers. The focus of attention here was on the beef sector. But there is likely to be growing interest in research in other food products and industries in which retailers' price response behavior is attracting attention and prompting policy concerns.

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[14]The original statement states unresponsiveness to farm prices rather than wholesale prices. However, it is wholesale prices that retailers face as an input cost. Moreover, as argued earlier, wholesale prices move in tandem with farm prices, extending our argument to movements in farm prices. Also, the original statement focused on the speed of adjustment, but our findings indicate that the difference between how the two series respond is mainly one of magnitude.

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