Asymmetric Adjustments in Vertical Price Transmission in the US Beef Sector: Testing for Differences among Product Cuts and Quality Grade

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Abstract

This study examines the dynamic relationship between wholesale and retail prices of beef products, accounting for product differentiation in cuts and quality grades. We test for long-run association in price transmission relationship in presence of asymmetry caused by threshold-type adjustments. The results indicate that there are significant asymmetric effects such that decrease or increase in wholesale beef prices tend to have different effects on the retail beef prices, and this effect varies across quality grades. Superior quality beef tend to show longer persistence to increase in prices and are adjusted at a slower rate than relatively inferior quality beef. This shows that adjustment of beef prices at retail market is influenced by the level of quality, thus supporting our original hypothesis.

Introduction

In a competitive market, a rise in input costs at the wholesale level, for example, is expected to be passed through to consumers via retailers. Similarly, a reduction in input costs is also expected to be reflected in the price paid by consumers to retailers. However, popular press news articles frequently lament the fact that retail prices rise rapidly, but fall slowly in response to price changes in upstream markets. Many empirical studies in economics have examined this asymmetric price response. Most studies model the price transmission (PT) between two or more markets using price margin behavior. Causality, time lag in response to price change, asymmetric response, and market structure are commonly considered in these studies. Peltzman (2000) analyzed prices of 77 consumer goods and 165 producer goods, including 120 agricultural products, and tested for the asymmetric price transmission (APT). He reports APT in about two-thirds of the goods studied. Fackler and Goodwin (2001), Meyer and von Cramon-Taubadel (2004), Frey and Manera (2005), and Goodwin (2006) review such studies. Goodwin (2006) notes that most studies in the literature prove existence of APT and provide a variety of explanations. Conforti (2004) summarizes that there are six factors affecting PT in spatial/vertical product markets: transaction/transportation costs; market power; increasing returns to scale in production; product homogeneity and differentiation; exchange rates; and, border and domestic policies.

While these studies have greatly improved the understanding of PT, most studies have considered aggregate commodity prices. Most studies on PT in the U.S. beef supply chain have been carried out considering beef as an aggregate product. For example, Boyd and Brorsen (1988) did not find evidence against symmetry in US pork marketing channel. In contrast, Hahn (1990) rejected symmetry in the adjustments of both pork and beef prices between retail and wholesale levels. Goodwin and Holt (2001) examine the U.S. farm, wholesale, and retail beef commodity prices and conclude that though there is APT it is modest in magnitude economically. Pozo et al (2013) performed TAECM estimation found no evidence against symmetry in beef vertical PT.

However, the U.S. meat industry has attempted to differentiate products based on intrinsic and extrinsic attributes. One such extrinsic attribute is quality certification implemented by the U.S. Department of Agriculture (USDA). Such certification could be considered as vertical product differentiation of beef. Grades are decided based on measurement of beef tenderness, juiciness, and flavor by USDA meat graders' subjective assessment, and by electronic instruments. Prime, Choice and Select are the three quality grades awarded; most of the graded beef sold in supermarkets is either USDA Choice or USDA Select (USDA 2012). Wholesale cutout beef data available from USDA- Agricultural Marketing Service (AMS) (2013) show that the shares of Choice, Select, and Prime in total graded and branded beef products are about 47%, 36%, and 1%, respectively. Prime is considered the highest quality grade based on marbling, and is generally sold in restaurants and hotels. Choice grade is considered to be superior in quality to the Select grade.

Our argument is that existence of such differentiated products could be a source of APT. Two potential reasons for APT here are market power and transaction costs such as adjustment cost and menu cost. Market power could arise since consumers' decision to purchase beef might be influenced by beef quality. Preference for specific quality attributes implies that sellers can discriminate prices based on quality, which in turn could lead to potential for manipulating prices away from long-run equilibrium prices. Transaction costs are incurred while adjusting prices for advertising and labeling costs. It may be assumed that sellers' decision on advertising and pricing strategy is influenced by exogenous factors such as seasonality in demand and supply, and unobservable geographical and demographic attributes. Seller may be reluctant to respond to changes in those costs unless the deviation of costs exceeds certain threshold. In such cases, larger changes in costs lead to a different price schedule than smaller changes. To model the effects of unobservable transaction costs specific models are available such as the threshold autoregressive (TAR) and momentum-TAR (MTAR) models.

In case of beef quality grades, in general, the higher quality (Choice) grade receives a higher price than the lower quality (Select) grade both at retail and wholesale levels. Moreover, for beef products of given quality grade, different cuts of meat show differences in prices. Meat cuts from the middle portion of carcass tend to be priced higher than those from the ends, and cuts like ribs, loin and sirloin are typically priced higher than round or briskets (please cite). Keeping in mind that quality differences could determine consumer purchase decision, the response in price of higher quality grade beef (say Choice) may be different from that of lower quality (say Select) to retain market share. Estimating price response of beef as an aggregate product would ignore such differential responses. Therefore, our results may not reflect the actual pricing behavior.

The objective of this study is to model the vertical price adjustment in the beef supply chain accounting for the differences in dynamic price adjustment behavior among beef products differentiated by quality. The intuition for this study is that given identifiable differences in quality, adjustments in prices of similar products towards cost changes may not be identical. Thus, this study contributes to the literature by viewing beef as a differentiated product unlike most previous analyses. The contribution is by looking for potential differences in PT for different product cuts and quality grades which might otherwise have been ignored/averaged out previously. Another contribution is in using appropriate methods to model the time series properties of data, namely the TAR and MTAR models.

In the ensuing sections, we provide the theoretical explanations of incentives for a monopolist to discriminate prices based on quality, followed by section on econometric models of PT asymmetry, data and estimation, and results from the analysis.

Price Discrimination Based on Product Quality

In this scenario, a monopolist is faced with the problem of choosing both the qualities of its products as well as their prices. Then, the preference of a consumer-surplus maximizing consumer can be represented as (Mussa and Rosen 1978; Maskin and Riley 1984):

$$(1)\quad U_i = \theta_i\, s - P_i = \theta_i\, V(q) - P_i\ (V(q)) = \theta_i\, V(q) - p(q)$$

where i represents individual consumer, U is the utility derived by the ith consumer by purchasing the product, θ_i is the consumer-specific taste/quality parameter (or the marginal utility of quality), s is an index of taste observable to the monopolist, P is the price paid, q is the unit cost of production and is given by $q = c(s)$ where c is the cost of producing quality s while c is an increasing function of quality, and $V(q)$ is the inverse function for $q = c(s)$. The price paid P is a function of quality, i.e. $\mathrm{P} = \mathrm{P(s)}$.

Let there be only one quality product being offered, while there are two types of consumers represented by their taste parameters θ_i where $\theta_2 > \theta_1$ and $\theta_2 > 1$. It is not optimum for the firm to price considering only one type of consumers: if only θ_1 consumers are considered, θ_2 consumers would still have a higher willingness to pay; if only θ_2 consumers are considered, then θ_1 consumers are priced out. As a solution, the firm can choose to provide two quality grades for which the firm charges price P_1 on θ_1 and price P_2 on $\theta_{2.}$. The firm can achieve sales to both groups of consumers and the profit π from sales to N consumers is given by:

$$(2)\quad \pi = N[\lambda(P_1 - c(s_1) + (1-\lambda)\ (P_2 - c(s_2))]$$

where λ represents the proportion of θ_1 consumers out of N consumers. Price P_1 must be set such that the θ_1 consumers would be willing and able to purchase. Price P_2 must be set such that θ_2 consumers prefer to purchase the higher quality product while ignoring the lower quality product. The prices are given by: $P_1 = \theta_1 s_1$, and $P_2 = \theta_2 s_2 - (\theta_2 - \theta_1)s_1$. The firm uses its ability to design product quality as per the requirements of consumers so that they self-select to purchase a given quality product. Pepall et al (2011) show that for a cost function given by $c(s) = \dfrac{s^2}{2}$, from the profit-maximizing condition, the difference in two prices is given by $P_1 - P_2 = \theta_2(\theta_2 - \theta_1)/\lambda$ though the marginal cost is only $s_2 - s_1 = \theta_2 - \theta_1/\lambda$, therefore is a case of price discrimination except when $\theta_2 = 1$. This shows that profit-maximizing monopolist has an incentive not only to discriminate price but also to keep the price of higher quality higher. Under menu pricing, the seller designs a menu of product bundles so as to make the consumer self-select the appropriate bundle based on the unobservable taste preferences which are otherwise hard to measure. However, bundling and bundle maintenance add to the seller's variable cost.

Considering the above seller pricing behavior, we hypothesize that products of higher quality would show stronger evidence for asymmetric PT and display longer lags. Increase in upstream prices would show more persistence and slower response than decreasing prices. This fits the profit-maximizing assumption for the seller. For higher quality beef we expect to see stronger evidence for positive asymmetry between retail price as dependent price and wholesale price as explanatory price, and test for its occurrence. Similarly, for higher quality beef we expect larger threshold effect in adjustment, or the threshold range for these products is expected to be bigger. This stickiness is to account for the expected higher transaction costs like greater promotion, display, and labeling costs.

Data and Econometric Models of APT

We have pooled two sets of time-series data on beef prices from publicly available sources. Wholesale data on boxed beef cutout values were collected from USDA Weekly Beef Archive (USDA, 2013). Retail data on individual grades and cuts were collected from the Bureau of Labor Statistics (BLS) consumer price indices and producer price indices (BLS, 2013). In some cases, the grade-cut data did not exactly match at the retail and wholesale levels because Select-grade data is not directly available at the retail level from BLS. We have assumed the BLS description "graded and ungraded, excluding USDA Prime and Choice," to be representative of Select grade, since retail sales mostly comprise Choice and Select graded products. Final data has 134 monthly observations on three cuts each for Choice and Select at wholesale and retail levels. The three cuts are chuck, round and sirloin, where sirloin can be considered to be of better quality than the other two. Since the Prime grade forms a very negligible share, it has not been included in further analysis.

The standard procedure for measuring APT involves estimation of the long-run equilibrium relationship among the concerned prices using tests for stationarity and cointegration. The Engle and Granger (EG) (1987) procedure for testing cointegration between two nonstationary I(1) price variables P_{1t} and P_{2t} consists of two stages; the first stage is estimation of an assumed long-run equilibrium relationship between the two variables using least squares as follows:

$$(3) \quad P_{1t}^{ik} = \beta_1 + \beta_2 P_{2t}^{ik} + \mu_t$$

where P_{1t}^{ik} and P_{2t}^{ik} are prices of beef products of ith grade and kth cut, 1 and 2 are the two market levels (retail and wholesale), $\beta_{1,}$ the intercept and β_2 the slope are the parameters to be estimated, and μ_t is an error term that may be autocorrelated. The intercept term is assumed to capture the effects of factors such as transportation costs and demographic effects. The error term represents the deviation from the long-run equilibrium. This step implies that the regular tests of stationarity/unit roots on levels and first-differences (represented by the Δ operator) have been conducted to ascertain that the variables are I(1).

The second step involves testing for the stationarity of the first-differenced error term using an Augmented Dickey-Fuller (ADF) test as follows:

$$(4) \quad \Delta\mu_t = \alpha \ \mu_{t-1} + \sum_{m=1}^{M} \gamma \ \Delta\mu_{t-1} + \varpi_t$$

where m is the lags of autoregressive terms selected by an appropriate Information Criterion, and ϖ_t is a white noise error term. If we reject the hypothesis that $\alpha = 0$ then the residual μ_t is stationary and the two variables P_{1t} and P_{2t} are cointegrated. Such linear representation of deviations (μ_t) has been criticized for the fact that they do not account for the effects of transaction costs on adjustments to deviation. Econometrically this procedure has been criticized by Enders and Granger (1998) who show that if μ_t is asymmetric, then the traditional unit root tests like the ADF test are misspecified. They proposed an alternative test which has more power than the traditional tests. Their model involves an asymmetric representation of adjustment as a TAR as follows:

$$(5) \quad \Delta\mu_t = I_t\left[\alpha_0^{OUT} + \alpha_1^{OUT}\mu_{t-1}\right] + (1 - I_t)\left[\alpha_0^{IN} + \alpha_1^{IN}\mu_{t-1}\right] + \sum_{k=1}^{K}\phi_k\Delta\mu_{t-1} + \upsilon_t$$

where I_t is the Heaviside indicator function defined as follows:

$$(6) \quad I_t = \begin{cases} 1 & if \quad \mu_{t-d} \geq \delta \\ 0 & if \quad \mu_{t-d} < \delta \end{cases}$$

Therefore, in this representation, the autoregressive (AR) process of μ_t is separated into two regimes, namely the outer and the inner regimes, based on whether the threshold variable μ_{t-d} is greater than the threshold value δ or not. The outer regime represents the deviations outside the range $[-\delta, +\delta]$ and the inner regime is defined by the observations falling inside this range. The parameter d represents the threshold delay parameter. It measures the time taken for switching from one regime to another.

A critical step is to test for threshold (nonlinearity) effects, and obtain estimates of the threshold parameter. Tsay (1986) suggested an F-test for the null hypothesis of an AR process against quadratic nonlinearity. This test is an improvement over Keenan test (1985) which adopted the idea of Tukey's (1949) one degree of freedom test for nonadditivity. Rejection of the null hypothesis indicates the presence of threshold δ. Hansen (1999) uses a residual bootstrap procedure to obtain p-values for the supremum Lagrange Multiplier (sup LM) test developed by Hansen (1996) for the existence of one threshold, and extends it for the case of two thresholds. After confirming presence of threshold effects, actual identification of the threshold δ is done by using the grid search procedure of Chan (1993). The procedure is based on the fact that if the value of the threshold is known then slope estimates of the regimes are least squares estimates. The threshold value is selected such that the lowest sum of squared residuals is obtained by estimating (5) for the observations sorted in ascending order by threshold variable μ_{t-d}. Hansen and Seo (2002) developed a test for a null hypothesis of linear cointegration versus an alternative of threshold cointegration. The test uses a maximum likelihood estimation of a bivariate threshold vector error correction model, unlike the univariate Keenan and Tsay tests.

The TAR model represents the idea that price variables might exhibit asymmetry where one price remains above the other, but for short intervals the lower price peaks above the higher prices (Ghoshray 2002). However, sometimes the underlying price series might behave such that falling prices lead to sharper decrease in gap between the two prices than when prices are increasing. For such conditions MTAR model is more appropriate since it adjustment represents greater momentum in one direction than the other (Ghoshray 2002). The MTAR model represents steep variations in residuals while the TAR model represents deep movements (Enders and Granger 1998). The representation of MTAR is similar to the TAR in (5) except for the changed definition of the Heaviside indicator function in (5):

$$(7) \quad I_t = \begin{cases} 1 & if \quad \Delta\mu_{t-d} \geq \delta \\ 0 & if \quad \Delta\mu_{t-d} < \delta \end{cases}$$

In the TAR or MTAR forms of (5), if the hypothesis that $\alpha_1 = \alpha_2$ is rejected, then asymmetry exists. If the hypothesis that $\alpha_1 = \alpha_2 = 0$ is rejected, it indicates significant threshold effects. Both tests can be conducted using an F-statistic.

We first investigate time series properties of the variables using unit root tests, and two cointegration tests (EG, and Johansen procedures). After this we test for threshold effects using the Tsay (1987) test and Keenan (1985) tests. However, both these tests detect quadratic nonlinearity and cannot be adequate for TAR models (Chan and Cryer, 2008). Hence, we then estimate the thresholds using Chan (1993) grid search process and test the statistical significance of these thresholds using Hansen (1999) test. Both TAR and MTAR models are estimated while searching for thresholds.

Results and Discussion

For the purpose of observing the basic characteristics of the beef price series used in this analysis, we arbitrarily divided each series in half, one from January 2002 to June 2007 and the other from July 2007 to February 2013. Descriptive statistics for the two periods are provided in Table 1. As expected there is a clear difference in average price between the two grades both at the wholesale and retail levels. Choice has a higher average price than Select for all cuts, but the difference in prices is the highest for sirloin. Though the standard deviation is larger for sirloin cuts in all the cases, the coefficient of variation does not show such behavior. Between the two time periods, there is a clear increase in average levels of all prices. The increase in prices is the least for sirloin cuts (about 2-10%), the highest for chuck cuts (about 17-24%), while for round cuts it is about 10-18%. Compared to the first period, the coefficient of variation is either stable or lower for five of the six retail price series in the second period, whereas it has increased for four of the six wholesale prices.

Output from Unit Root Tests and Linear Cointegration Tests

To understand the relationship between prices of quality varieties of beef, it is important to consider the underlying data generating mechanism for such price series. First the prices were converted into logarithms since it helps in controlling possible non-normality. Order of integration was obtained by conducting unit root tests, and results are in Table 2. Trend and constant terms were included in the tests based on visual inspection of series, while lag length was selected based on the Bayesian Information Criterion (BIC). The ADF test did not reject the presence of unit root in all 12 of the prices at level, while PP test rejected the hypothesis of unit root in five of the twelve beef price series. The DF-GLS test (results not presented here) showed all the level series to contain unit root. Therefore, we took the first difference and tested again for the presence of unit root. None of the first-differenced data followed unit root process, thus showing the prices to be integrated of order one.

Table 3 shows the output from cointegration tests from the Johansen procedure and the EG procedure where the corresponding lag lengths were selected based on BIC. The Johansen test does not reject the existence of at least one (r=1) cointegrating relationship in each of the six price pairs. The EG procedure too supports the existence of cointegrating relationship in all six price pairs. Therefore, the wholesale and retail markets for these products are shown to be integrated by these two tests. However, if there are threshold effects, then the results from tests such as EG would have lesser power. For example, Ghoshray (2009) found that though the null hypothesis of no cointegration could not be rejected in the EG test, the results differed when threshold effect was introduced in the adjustment process. Therefore, next we test for threshold nonlinearities.

Nonlinearity Tests, Threshold Model Estimation and Hypothesis Tests

We have used the packages "tsDyn" (Stigler, 2010) and "TSA" (Chan and Ripley, 2012) available in the R software for detecting nonlinearities and thresholds. Table 4 shows the nonlinearity and threshold test results when retail prices are the dependent variables. Two nonlinearity tests used here are the Keenan test and Tsay test, while Hansen test is used for testing the statistical significance of the identified thresholds. The delay parameter is selected as the lag value showing the largest TAR-F statistic in the Tsay test, though many of the empirical applications assume a delay value of one (Goodwin and Holt, 1999; Goodwin and Piggott, 2001; Lee and Gomez 2012). Three versions of Hansen test are used to test for null hypotheses of linearity against alternative of one threshold, linearity against two-thresholds, and one-threshold against two-thresholds. These are denoted in the table as Hansen-1, Hansen-2, and Hansen-3, respectively The first two can be seen as linearity tests, whereas the third can be seen as a specification test; once the 1vs2 and/or 1vs3 rejected the linearity and indicate the presence of threshold effect, the 2vs3 test indicates if a model with one or two thresholds is preferable (Stigler, 2010). Hansen and Seo (2002) test is further used to check if the null hypothesis of linear cointegration versus the alternative of threshold cointegration holds.

Results show that the Keenan test and Tsay test show significant quadratic nonlinear effects in all three of the Choice cuts but only in the chuck cut of Select grade. However, as noted previously the two tests cannot be adequate for TAR models. We use the delay values yielding the largest test statistic in the Tsay test for threshold detection in the Hansen tests. Delay parameter shows the time delay in switching from the outer and inner regimes (OUT and IN in equation 5) in the error correction mechanism. From the results, it is seen that the Choice cuts have larger and significant delays, while the Select cuts do not show statistical significance in two cases. This might indicate that time required for regime switching is more delayed in the case of the Choice cuts. While Hansen-1 doesn't reject the hypothesis of linearity in any equation, Hansen-2 strongly rejects it in all the cases. Hanse-3 results do not show evidence against one-regime. Therefore, we conclude that all equations relationships contain two regimes (one threshold). Hansen and Seo (2002) test rejects the null of linear cointegration in all but Select-round equation. The percentage of observations falling in the upper regime are those deviations outside the interval $[-\delta, \delta]$. Choice chuck has relatively higher percentage in this region which might indicate that for this product price adjustment takes place less frequently as more deviations are in the outer region. Such observations for the Select cuts ranges between 33-45%. Finally, the threshold value can be interpreted as representing the region $[-\delta, \delta]$ where no price adjustment towards the long-run equilibrium occurs. This region is considerably larger in the case of all three Choice cuts, with the highest being sirloin; among Select cuts only sirloin cut shows larger threshold value. For these products, prices are more sensitive than for the other two products. This might be taken as supportive of our hypothesis that better quality products show larger threshold effects.

We have also used Akaike Information Criterion (AIC) and BIC to guide in model selection, i.e., zero vs one vs two regime models. Table 5 shows these values. The AIC values of the one-regime model are the lowest in all cases, while BIC values support this argument in the case of Select grade only. Overall, these values could be thought of as supporting the findings from Table 4.

Table 6 shows the estimates from TAR and MTAR models. We used the model AIC values to select the optimum number of lags in regimes. Comparing the model AIC and BIC values, we see that the TAR model shows better fit than MTAR models. Therefore, the long-run adjustment process for these relationships can be assumed to possess deep movements in one direction than the other. Coefficient of the inner regime is represented by $\alpha_1 - IN$ while that of the outer regime by $\alpha_1 - OUT$. The finding that threshold cointegration tests are significant indicate that the linear cointegration tests such as the Johansen and EG test are less powerful. First four out of six relationships have both the coefficients significant at 10% level of significance or lower. For Choice-round, Choice-sirloin and Select-sirloin, $|\alpha_1 - IN| <$ $|\alpha_1 - OUT|$, which means that the model exhibits more adjustment towards negative deviations than for positive deviations. This supports our hypothesis that for higher quality products decrease in upstream prices are likely to elicit greater response than decrease, and increase in prices tend to persist longer. These rate of adjustments are quite large, with the largest adjustment occurring for the two Choice cuts. Therefore, for higher quality retail beef the adjustments show faster response for decreasing prices and show greater extent of adjustment, than to increasing wholesale prices. For Select-round and Select-sirloin, the adjustment process is significant in only one direction than the other. For them adjustment takes place in only one direction. The table shows that the TAR model rejects the hypothesis of symmetry in all six cases, while the MTAR rejects in only one case. Since TAR model is found to be better fit than MTAR model, we can conclude that there is indeed asymmetric reaction in the adjustment behavior in all six cases. The threshold F test again shows that there are significant threshold effects, and thus estimating a linear cointegration and linear autoregressive adjustment process is not the correct procedure.

Conclusions

This study analyzed the differences in responses prices of retail beef products of different qualities. Choice and Select were the two quality grades each with three different qualities of cuts. We found that there are significant threshold effects in the adjustment process, i.e., all price changes at the wholesale level do not elicit the same type and extent of response at the wholesale level. Threshold delays showed that significant larger delays were mostly observed in the case of Choice grade products. Since such delay is indicative of lags in response of retail prices, it showed that prices of Choice products, which are superior in quality to Select, show delayed response to changes in wholesale prices. Choice products had considerably larger threshold values, indicating that it takes larger changes in wholesale prices of these products to elicit responses at the retail level, than for Select products. From the threshold tests, it was confirmed that the estimated thresholds were statistically significant. The TAR model fitted the models better than the MTAR model. Choice products showed larger extent of adjustment to negative price deviations and a faster rate of adjustment. In other words, positive deviations would show more persistence. Overall, we can say that these observations have supported our hypothesis that better quality products would show stronger threshold effects and more response towards falling prices than to increasing prices. This study could further be improved by developing error correction models and impulse response functions to check for extent of asymmetry.

Table 1. Descriptive Statistics of Retail and Wholesale Beef Products

	January 2002 to June 2007			July 2007 to February 2013		
Price Series	Average	Standard Deviation	Coefficient of Variation	Average	Standard Deviation	Coefficient of Variation
Retail Prices ($/lb)						
Choice-chuck	3.13	0.24	0.08	3.91	0.39	0.10
Choice-round	3.93	0.27	0.07	4.40	0.29	0.07
Choice-sirloin	5.99	0.53	0.09	6.12	0.39	0.06
Select-chuck	3.00	0.28	0.09	3.63	0.27	0.07
Select-round	3.78	0.28	0.08	4.34	0.33	0.08
Select-sirloin	5.06	0.44	0.09	5.38	0.35	0.07
Wholesale Prices ($/cwt)						
Choice-chuck	100.16	11.01	0.11	131.64	20.05	0.15
Choice-round	118.51	13.75	0.12	142.44	18.47	0.13
Choice-sirloin	215.09	29.35	0.14	224.68	25.31	0.11
Select-chuck	100.03	10.86	0.11	130.52	19.44	0.15
Select-round	116.28	13.38	0.12	142.11	18.14	0.13
Select-sirloin	185.93	20.62	0.11	205.60	22.45	0.11

Table 2. Results from Unit Root Tests

Price Series	Prices at levels		First-differenced prices	
	ADF	PP	ADF	PP
Retail beef prices				
Choice-chuck	-2.663	-3.321*	-13.855***	-14.205***
Choice-round	-2.618	-3.557**	-15.99***	-16.514***
Choice-sirloin	-2.637	-3.068	-11.401***	-11.593***
Select-chuck	-3.036	-4.860***	-15.687***	-17.164***
Select-round	-1.888	-2.7	-13.49***	-14.047***
Select-sirloin	-2.945	-2.923	-12.56***	-12.668***
Wholesale beef prices				
Choice-chuck	-1.979	-2.871	-9.725***	-9.613***
Choice-round	-2.199	-3.129*	-8.859***	-8.6***
Choice-sirloin	-1.885	-3.297*	-8.29***	-7.898***
Select-chuck	-2.028	-2.924	-9.94***	-9.854***
Select-round	-2.299	-3.250*	-9.337***	-9.166***
Select-sirloin	-2.176	-3.109	-8.185***	-7.904***
Fuel price	-1.945	-1.908	-8.557***	-8.688***
Hourly retail worker earnings	-1.156	-1.143	-15.903***	-16.495***
5-area slaughter cattle (heifer) price	-0.582	-0.882	-7.701***	-7.345***
Hourly meat processing worker earnings	-1.73	-1.852	-12.618***	-12.574***

*significant at 10%, ** significance at 5%, *** significance at 1% levels of significance

Table 3. Cointegration Test Output from Johansen Procedure and Engle-Grange Procedure

	Johansen Procedure			E-G Procedure
Retail -Wholesale Price Pair	lags	Trace Statistic# for r=0§	Trace Statistic# for r=1§§	t-statistic
Choice-chuck	2	39.751	8.259*	-0.393***
Choice-round	2	53.776	8.602*	-0.521***
Choice-sirloin	2	37.581	8.405*	-0.233***
Select-chuck	1	64.323	11.190*	-0.469***
Select-round	1	56.363	11.270*	-0.312***
Select-sirloin	2	45.593	7.724*	-0.208***

r =number of maximum cointegrating vectors;
§ critical value is 25.32; **§§** critical value is 12.25;
* indicates non-rejection of existence of at least one cointegrating vector;
*** indicates significance at 1% level of significance or higher.

Table 4. Tests for nonlinearities and threshold effects; dependent variable-retail prices

	Choice-chuck	Choice-round	Choice-sirloin	Select-chuck	Select-round	Select-sirloin
Keenan Test	3.178*	1.990	3.107*	2.859*	0.252	1.977
Tsay Test	4.616**	2.012**	4.711***	4.051**	1.027	0.860
Threshold Delay	1	4	3	1	10	8
Hansen-1	12.759	15.335	43.854***	17.904	31.492	16.432
Hansen-2	26.283***	34.296***	66.275***	30.562***	66.982***	43.026***
Hansen-3	12.289	16.932	16.699	11.094	28.208	23.483
Hansen and Seo Test	22.019***	66.805*	24.309***	61.092*	33.88611	55.244***
% Observations in Outer Regime	72.65	30.17	18.97	45.38	33.33	40.52
Threshold Values	-0.0198	0.0141	0.0261	0.0026	0.0097	0.0141

*significant at 10%, ** significance at 5%, *** significance at 1% levels of significance

Table 5. Information Criteria for Regime Selection; dependent variable-retail prices

	AIC			BIC		
	Linear	1 Regime	2 Regimes	Linear	1 Regime	2 Regimes
Choice-chuck	-913.082	-930.485	-919.530	-863.946	-849.555	-827.039
Choice-round	-959.116	-979.639	-971.077	-907.089	-884.257	-855.463
Choice-sirloin	-870.642	-909.916	-881.443	-818.615	-843.438	-817.856
Select-chuck	-902.742	-924.540	-923.290	-850.715	-904.307	-842.360
Select-round	-970.335	-981.917	-964.788	-921.199	-935.671	-869.407
Select-sirloin	-865.020	-879.086	-868.349	-812.994	-850.182	-816.323

Table 6. Regime Estimates and Hypothesis Tests from Threshold Models; Dependent Variable- Retail Prices

Model		Choice-chuck	Choice-round	Choice-sirloin	Select-chuck	Select-round	Select-sirloin
TAR							
	AIC	-932.485	-981.639	-898.936	-922.063	-978.022	-878.691
	BIC	-854.446	-889.148	-844.019	-901.831	-931.776	-852.678
	$\alpha_1 - IN$	-0.951***	-0.354***	-0.153*	-0.712***	-0.374***	-0.127
	$\alpha_1 - OUT$	-0.368***	-1.080***	-0.804**	-0.374***	-0.035	-0.526**
	Symmetry Test	8.410***	9.581***	3.563*	4.014**	3.445*	3.150*
	Threshold F-Test	43.633***	40.921***	8.601**	51.705***	9.629***	8.459**
MTAR							
	AIC	-912.399	-954.923	-878.123	-907.058	-975.858	-866.378
	BIC	-834.359	-862.432	-823.207	-886.825	-929.612	-840.365
	$\alpha_1 - IN$	-0.595***	-0.324***	-0.141	-0.546***	-0.311***	-0.168
	$\alpha_1 - OUT$	-0.401***	-0.701***	-0.237	-0.431***	0.049	-0.155
	Symmetry Test	0.978	2.43	0.181	0.46	3.741*	0.005
	Threshold F-Test	26.181***	19.918***	3.34	35.012***	8.034**	3.553

*significant at 10%, ** significance at 5%, *** significance at 1% levels of significance

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