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Is There Asymmetric Price Transmission in the U.S. Fluid Milk Market?

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Abstract

This study is used to examine the characteristics of the retail price adjustment within the U.S. fluid whole milk market. We employ an error correction model (ECM) to test for asymmetry in the transmission of farm milk price changes to changes in the retail price. In this analysis we use monthly data of farm and retail whole milk prices encompassing the January 2001 to December 2011 period for 16 U.S. cities. Several cities were found to display asymmetric price transmission where retail prices tended to respond more quickly with farm price increases vs. decreases.

Keywords: Retail Milk Price, Price Transmission, Price Asymmetry.

I. Introduction

Over the last 25 years, the price of raw farm milk has become extremely volatile. For example, in November 2008 the U.S. average farm *All Milk* price was $17.10/cwt.[1] By February 2009, 4 months later, this had decrease to $11.60/cwt, a 32% decrease. The volatility at the farm level has raised concern as to the existence of asymmetric transmission between retail and farm price changes and the possible relative welfare impacts on consumers, dairy farm operators, processing firms, and retailers. According to the USDA, raw milk costs represent more than half of the retail price of fluid milk (USDA, 2016). Historically asymmetry in price transmission refers to the environment where the retail price tends to respond faster to farm price increase when compared to the timing of upstream price decreases.

There is an extensive literature on dairy-related milk pricing asymmetry. Based on monthly farm and retail data over the 1971-1981, Kinnucan and Forker (1987) found asymmetric price movements in four U.S. dairy products: butter, cheese, fluid milk and ice cream. Capps and Sherwell (2007) identified price asymmetry in fluid milk prices in seven cities using an asymmetric error correction model (ECM).

Meyer and von Cramon (2004) note that asymmetry in price transmission can be classified based on whether it is the speed and/or magnitude of price transmission that is asymmetric. Lass, Adanu, and Allen (2001) found both short run and long run retail price asymmetry in the Boston and Hartford milk markets with asymmetry found in both speed and magnitude. Price asymmetry can also be differentiated depending on whether the transmission process concerns vertical or spatial markets. Most studies of dairy markets are focused on vertical market impacts, for instance, the impacts of farm price changes on retail milk prices.

The causes of price asymmetry have typically been found to be the presence of differential market power and/or significant adjustment costs. Most studies of price asymmetry have focused on the existence of non-competitive market structures. In his analysis of a variety of markets, Peltzman (2000) uses two proxies for market power: (i)

[1] The *All Milk* price is a representative farm-level milk price defined as total receipts from the sale of raw milk divided by the total cwt marketed. In the U.S. fluid milk is priced once a month. Refer to Cropp and Jesse (2008) for more detail.

the number of competitors and (ii) market concentration as measured by the Herfindahl-Hirschman index. Meyer (2004) notes that only the presence of unequal market power is capable of leading to long lasting price change asymmetries. However, some important issues are still unresolved and require careful investigation. For generic commodities are there regional differences in the degree of asymmetry? Has the farm-retail price relationship changed over time? What is the role of retail market structure on this asymmetry?

In this paper, we investigate whether there is asymmetric retail price response of changes in farm milk prices. We compare the differences in retail price changes from both a speed and magnitude perspective. We employ monthly data of raw farm milk and retail whole milk prices over the January 2001 to December 2011 for 16 U.S. metropolitan areas. First, we introduce the two main econometric methods used to test for the presence of asymmetry. Then several tests are conducted to determine the specific econometric model for each city based on city-specific time series properties. Third, we compare the estimated regression coefficients associated with retrial price increases vs. decreases to investigate whether price asymmetry exists in each region's retail milk market. Finally, the concentration ratio is used in the understanding of each regions price asymmetry characteristics.

II. Asymmetric Price Transmission Evaluation Methodologies

There have been a variety of modeling techniques used to test for the presence and degree of asymmetric price transmission (Meyer and von Cramon, 2004). In this section, we discuss the development of econometric methods used in the present analysis.

2.1 The Wolffram-Houck Model

Wolffram (1971) employed a variable splitting technique to estimate asymmetric price adjustment. Houck (1977) extended this model and developed a static system to test based on the magnitude of increases and decreases of retail and farm prices:

(1) $$R_t^* = \alpha_0 + \alpha_1 F_t^{*+} + \alpha_2 F_t^{*-} + \varepsilon_t$$

Where $R_t^* = R_t - R_0 = \sum_{i=0}^{t} \Delta R_i$ is the retail price deviations from the initial value. R_t is the retail price at time t, F_t^{*+} is the cumulative increase of farm price from initial value (e.g., t = 0) to the current time period, t. F_t^{*-} is similarly the cumulative decrease of farm price from initial value at time t=0 until the value observed in time t. α_0 is the time trend coefficient. Price asymmetry is tested by determining whether $\alpha_1 = \alpha_2$.

There are possible delay responses of retail prices to changes in farm prices. Ward (1982) and Kinnucan and Forker (1987) introduced lagged cumulative change variables in a dynamic extension to (1):

$$R_t^* = \alpha_0 + \sum_{k=0}^{m_1} \alpha_{1,k} F_{t-k}^{*+} + \sum_{k=0}^{m_2} \alpha_{2,k} F_{t-k}^{*-} + \varepsilon_t \quad (2)$$

In (2) the coefficients $\alpha_{1,k}$ and $\alpha_{2,k}$ represent the net effect of rising and falling farm prices on retail prices, respectively. m_1 and m_2 represent the number of lags for rising and falling farm prices included in the model, which can differ across scenario.

The use of (2) enables us to distinguish between short-run vs long-run price asymmetry. The test of long-run symmetry is:

$$H_0 : \sum_{k=0}^{m_1} \alpha_{1,k} = \sum_{k=0}^{m_2} \alpha_{2,k} \quad (3)$$

We can conclude price asymmetry by rejecting the null hypothesis. Meanwhile, the short-run symmetry can be determined by testing whether the coefficients of the first period are equal or not, i.e. $\alpha_{1,1} = \alpha_{2,1}$.

2.2 The Error Correction Model

As Granger and Newbold (1974) note, regressions between non-stationary or highly autocorrelated stationary time series lead to spurious regression. The empirical applications based on the Wolffram-Houck specification do not account for the time series property of retail and farm level milk prices into consideration and are not consistent with cointegration. The retail and farm price may share a long-run relationship if cointegrated via the following:

$$R_t = \beta_0 + \beta_1 F_t + \varepsilon_t \quad (4)$$

After a shock in retail price R_t or farm price F_t, they will move towards the long-run relationship. Hence, tests based on equation (1) or (2) don't take this relationship into

consideration. Von Cramon-Taubadel (1997) first pointed out this limitation and suggested the inclusion of error correction terms (ECT) (Appel, 1992; Kinnucan and Forker, 1987; Pick et al., 1990; Zhang et al., 1995). As Granger and Lee (1987) originally proposed an alternative price transmission specification can be represented as:

$$\triangle R_t = \alpha_0 + \sum_{k=0}^{m_1} \alpha_{1k}^{+} \triangle F_{t-k}^{+} + \sum_{k=0}^{m_2} \alpha_{1k}^{-} \triangle F_{t-k}^{-} + \alpha_3 ECT_t + u_t \tag{5}$$

where $ECT_t = R_t - \beta_0 - \beta_1 F_t$, the residual from the cointegration relationship of retail and farm price. The constant term can be omitted. Many empirical research contains it to test whether labor, energy or other input costs have changed (Awokuse and wang, 2009).

Von Cramon-Taubadel and Loy (1999) made further modification by segmentation of farm price changes. This specification, known as the *threshold ECM*, allows for different speeds of adjustment between the rising and falling of farm price.

$$\triangle R_t = \alpha_0 + \sum_{k=0}^{m_1} \alpha_1^{+} \triangle F_{t-k}^{+} + \sum_{k=0}^{m_2} \alpha_1^{-} \triangle F_{t-k}^{-} + \alpha_3^{+} ECT_t^{+} + \alpha_3^{-} ECT_t^{-} + u_t \tag{6}$$

where ECT_t^{+} represents the error correction term for price increases and ECT_t^{-} the error correction term for decreases. Equation (6) is based on a linear error correction in that a constant proportion of any deviation from the long-run equilibrium is corrected, regardless of the size of the deviation.

As suggest by Goodwin and Holt (1999), an alternative version of the above is one which allows for some stickiness in retail price responses. Under this threshold model specification there will be retail prices changes only if the farm level changes are above (below) a certain level. In Figure 1, no error correction takes place when ECT lies between [C_1, C_2]. Goodwin and Holt (1999) estimated a three-regime threshold ECT which allows for a neutral band, which means there is no error correction when the price deviation is small compared to price adjustment costs. We can represent their model via the following:

$$\triangle R_t = \begin{cases} \alpha_0 + \sum_{k=0}^{m_1} \alpha_1^{+} \triangle F_{t-k}^{+} + \sum_{k=0}^{m_2} \alpha_1^{-} \triangle F_{t-k}^{-} + \phi_1 ECT_t + u_t & \text{if } \mathrm{ECT_t} > \mathrm{c_2} \\ \alpha_0 + \sum_{k=0}^{m_1} \alpha_1^{+} \triangle F_{t-k}^{+} + \sum_{k=0}^{m_2} \alpha_1^{-} \triangle F_{t-k}^{-} & \text{if } \mathrm{c_1} < \mathrm{ECT_t} < \mathrm{c_2} \\ \alpha_0 + \sum_{k=0}^{m_1} \alpha_1^{+} \triangle F_{t-k}^{+} + \sum_{k=0}^{m_2} \alpha_1^{-} \triangle F_{t-k}^{-} + \phi_3 ECT_t + u_t & \text{if } \mathrm{ECT_t} < \mathrm{c_1} \end{cases} \tag{7}$$

We extend (7) substituting the linear correction with the cubic polynomial error correction specification proposed by Escribano (2004). This extension is represented in equation (8):

$$\Delta R_t = \alpha_0 + \sum_{k=0}^{M_1} \alpha_{1k}^{+} \Delta F_{t-k}^{+} + \sum_{k=0}^{M_2} \alpha_{1k}^{-} \Delta F_{t-k}^{-} + \gamma_1 ECT_t + \gamma_2 ECT_t^2 + \gamma_3 ECT_t^3 + u_t \quad (8)$$

Note that with the use of polynomial error correction allows the change in the retail pieces to be a continuous, nonlinear function of the ECT and eliminate the characteristic of having a *knife-edged regime switch* that characterizes threshold method represented in eq. (7) (Mainardi, 2001; Stewart and Blayney, 2011).

2.3 Selection between the two models

Note that eq. (8) is similar to the first-difference version of eq. (2) except for the additional presence of the ECT terms. The ECT model nests the Houck specification when the lag lengths for ΔF^{+} and ΔF^{-} in the two models are the same, i.e. $m_1=m_2=M$. When taking cointegration into consideration, we can conclude that the ECT model is superior to Houck one. If the retail and farm prices are not cointegrated, we can use Akaike Information Criterion (AIC) (Akaike, 1974) or the Schwarz Information Criterion (SIC) (Schwarz, 1978) to determine the optimal lag length and therefore decide which method is suitable.

III. Description of The Farm and Retail Milk Price Data

The data used in the analysis are weekly scanner-based retail price data for whole milk covering the period January 2001–December 2011. Retail milk prices ($/gal) were obtained from the IRI Marketing Data Set, which contains milk-based revenues and quantities sold of a majority of retail grocery and drug store outlets in 46 U.S. cities[2]. The level of detail is at the Universal Product Code (UPC) level.

To match with available farm value, we select 16 cities for which we were able to obtain both retail and farm prices. We aggregate the milk sales to represent monthly sales level by assigning the weekly data to the month in which most of that week's data applies.

[2] IRI collects data from chains/stores that has more than two million dollars in sale.

We partitioned the monthly sales data into four milk types differentiated by content: skim milk (0−0.5%), 1% milk (0.5%−1.5%), reduced fat milk (1.5%−3.25%) and whole milk (>3.25%).

For farm-level milk price data, we use the Announced Cooperative Class I prices report by USDA's *Dairy Market News* and pertain to the Federal Milk Marketing Order (FMMO) system.[3] Class I milk under the FMMO system is milk that is used for bottling purposes. The minimum Class I milk price is determined on a Wednesday by the 23rd of the month prior to the month of production to which the price pertains. This price represents the *minimum* amount that has to be paid by participating dairy processors for standardized milk being used for Class I purposes. This Class I price is composed of (i) a *Base Mover* that is determined by monthly plant prices for cheddar cheese blocks, butter, non-fat dry milk and dried whey; and (ii) a Class I price differential, which tends to increases the farther a processing plant is located from the upper Midwest. The Cooperative Class I milk price is the price that major cooperatives servicing a particular metropolitan area are charging for Class I milk. The Class I prices in these announcements generally are higher than the FMMO established minimum Class I prices as these over-order prices include charges for various services performed by the cooperative for the processor and the fat, protein and other solids concentration of marketed milk differs from that of standardized milk.[4]

Figure 2 is used to provide a comparison of retail whole milk and Coop Class I farm prices for Boston over the study period. The correlation between these two series is 0.851. The average basis of the study period was $2.15/gallon. The range of these basis values was from $1.801-$2.54 with a relatively small coefficient of variation of 0.080.

Descriptive statistics associated with these price series are shown in Table 1[5]. The average retail price of retail whole milk ranged from $2.58/gal per gallon in New Orleans to $4.05 per gallon in Phoenix. The farm prices range from $1.06 per gallon in Seattle to $1.35 per gallon in New Orleans over the study period. The variance of farm prices is

[3] For an overview of the FMMO system refer to Cropp and Jesse (2008)
[4] Under FMMO regulations, standardized milk is composed of 3.5% fat, 2.99% protein and 5.69% other solids
[5] When reviewing the range of milk prices shown it should be remembered that these prices include the Class I differential. This differential ranges from $1.70/cwt in the Minneapolis region to $3.43/cwt in the Seattle area.

smaller than that of retail prices. The price of farm price fails to indicate the low rank of retail price. For instance, Seattle has the lowest farm price, but the retail price is among one of the highest in the 16 cities. Similar situation happens for Hartford.

IV. Evaluation of Alternative Model Specifications

As noted above, we undertook a variety statistical test to identify city-specific pricing model specifications. We used the time series characteristics of the monthly milk price data to determine the specific econometric model used in our final estimation. Figure 3 provides an overview of our testing procedures. This is followed by a discussion of each test and test results.

4.1 The Augmented Dickey-Fuller (ADF) Test

The error correction model can be applied only when the retail and farm prices are co-integrated of the same order. We use the Augmented Dickey-Fuller (ADF) test to check on the stationarity of retail and farm price series in each city (Brockwell and Davis, 1991). Under the ADF test the null hypothesis is that the price series are non-stationary and need to be differenced. The final lag-length is determined via Akaike Information Criterion (AIC) values (Greene, 2000). If a variable is stationary, it is integrated of order zero, I(0). Similarly, if its first difference is stationary, then the variable is integrated of order one, I(1). If the farm price and retail price are integrated of different orders, the error correction terms are not used.

The ADF test results obtained for our 16 metro area sample is presented in Table 2. For Chicago, Phoenix and St. Louis we found the retail milk price is integrated of order 0 while the farm price was found to be integrated of order 1. For the remaining areas both series where found to be integrated of order 1. Via Enders (2015), if two variables are integrated of different orders, then one concludes these two variables are not co-integrated. For those cities whose price series are integrated of the same order, we then use the Johansen test to check for co-integration.

4.2 Johansen Cointegration Test

Error correction terms are included if retail and farm prices are cointegrated. We use Johansen's test to check for cointegration. The number of cointegrating vectors between two variables can be at most one. The null hypothesis under the Johansen test is that the rank (r) of cointegrating vectors between retail and farm prices is zero (i.e. r = 0), which implies non-integration. There are two test statistics available under the Johansen test: λ_{max} and λ_{trace}. The alternative hypothesis using λ_{trace} is one or more cointegrating vectors (r > 0). Alternatively with the use of λ_{max} the alternative hypothesis is that r = 1 (Enders, 2015).

The results of Johansen's cointegration tests are shown in Tables 3. For this analysis we use the test results associated with the more restrictive, λ_{max} statistic. Farm and retail milk prices were found to be cointegrated for the Boston, Cleveland, Oklahoma City, Omaha, Philadelphia, St. Louis and Seattle markets. Given these results, we use models which allow for asymmetric ECM specifications, i.e. models (A), (B), (E) or (F), for these cities with cointegration.

4.3 Granger Causality Test

Granger causality tests were used to test whether the farm prices cause retail price and whether retail price causes farm price. To evaluate the causal relationship between our two price series we estimate an OLS model of farm price on its own lagged and lagged retail price values as well as a similar regression of retail price on its own lagged and lagged farm prices. If farm prices Granger cause retail milk prices, then in the regression where retail milk price is the dependent variable, the F-test corresponding to all lagged farm price coefficients should be statistically significant. If retail prices fail to Granger cause farm prices, then, in the regression where farm price is the dependent variable, the F-test corresponding to all coefficients associated with lagged retail prices should not be statistically significant. Our Grainger causality regressions without cointegration can be represented via following:

(9) $$\Delta F_t = \sum_{k=1}^{K} (\alpha_{1k}^{+} \Delta F_{t-k}^{+}) + \sum_{k=1}^{K} (\alpha_{1k}^{-} \Delta F_{t-k}^{-}) + \sum_{k=1}^{K} \alpha_{2k} \Delta R_{t-k}$$

(10) $$\Delta R_t = \sum_{k=1}^{K}(\alpha_{1k}^{+}\Delta F_{t-k}^{+}) + \sum_{k=1}^{K}(\alpha_{1k}^{-}\Delta F_{t-k}^{-}) + \sum_{k=1}^{K}\alpha_{2k}\Delta R_{t-k}$$

If the farm and retail prices are cointegrated, we can represent our causality test regression as:

(11) $$\Delta F_t = \sum_{k=1}^{K}(\alpha_{1k}^{+}\Delta F_{t-k}^{+}) + \sum_{k=1}^{K}(\alpha_{1k}^{-}\Delta F_{t-k}^{-}) + \sum_{k=1}^{K}\alpha_{2k}\Delta R_{t-k} + \gamma_1 ECT_t$$

(12) $$\Delta R_t = \sum_{k=1}^{K}(\alpha_{1k}^{+}\Delta F_{t-k}^{+}) + \sum_{k=1}^{K}(\alpha_{1k}^{-}\Delta F_{t-k}^{-}) + \sum_{k=1}^{K}\alpha_{2k}\Delta R_{t-k} + \gamma_2 ECT_t$$

Given the above model specifications, we test the null hypothesis that $\alpha_{1k}^{+} = \alpha_{1k}^{-}$ and $\sum_{k=1}^{K}\alpha_{1k}^{+} = \sum_{k=1}^{K}\alpha_{1k}^{-}$. If the null hypothesis is rejected, we estimate a multiple-equation, vector autoregression (VAR) specification.

As shown in Tables 4, the Granger causality tests indicate that Chicago, Detroit, Hartford, Milwaukee, Phoenix and St. Louis, econometrically support the underlying assumption that farm prices Granger cause retail prices. For Cleveland, Dallas, Oklahoma City, Phoenix and Seattle, the farm price depends on the retail price. The null hypothesis no cointegration is supported by the data.

4.4 Quandt-Andrews Test:

We use the Quandt-Andrews Test to determine market specific parameter instability and structural change (Andrews, 1993). The test-specific regression model can be represented as the following:

(13) $$\Delta R_t = \alpha_0 + \sum_{k=1}^{m}(\alpha_{1k}^{+}\Delta F_{t-k}^{+}) + \sum_{k=1}^{n}(\alpha_{1k}^{-}\Delta F_{t-k}^{-}) + \mathrm{I}_{at}\sum_{k=1}^{m}(\alpha_{2k}^{+}\Delta F_{t-k}^{+}) + \mathrm{I}_{at}\sum_{k=1}^{n}(\alpha_{2k}^{-}\Delta F_{t-k}^{-}) + u_t$$

Let I_{at} be a dummy variable identifying whether a set of farm price changes happen after the structural change point. This model can be modified to include error correction terms or multiple-equations based on the test results from Johansen and Granger Causality Tests. To test whether there is a structural break, the null hypothesis is $\sum_{k=1}^{m}\alpha_{2k}^{+} = 0$ and $\sum_{k=1}^{n}\alpha_{2k}^{-} = 0$, which indicates the coefficients are stable in the econometric model over the entire study period. Since the potential breaking point is unknown, we need to compute and compare the likelihood ratio-like test statistics for potential points and choose the one with largest chi-squared critical value of its

asymptotic distribution. As proposed by Hansen (1997), respect to the number of observations needed, the proper interval for testing is [20%, 80%], hence we conduct the test for $t \in [20, 112]$. The null hypothesis of there is no structural break is rejected if there exists at least one test statistic larger than its corresponding critical value. If we reject the null hypothesis of no structural break we include the area-specific structural break in our final model of testing for the presence of price asymmetry.

4.5 Summary of Alternative Model Specifications

Depending on the results of the various model specification tests the final econometric model estimated will be area specific. The model used in each area is shown in Table 5 and specified equations are in Table 6.

V. Farm to Retail Price Transmission Characteristics

5.1 Testing for Price Asymmetry

The estimation results for the market areas and model types included in this analysis are shown in Tables 6, 7(a) and 7(b). When the error correction term is tested to be needed, we use polynomial distributed lag models to estimate the lag structure (Almon, 1965). The lag length is determined by Akaike information criteria (AIC). Generally, the lag length equals one for most of cities, which indicates the time for retail price to adjust to changes in farm price is one months.

For all the cities except Boston and Omaha, the coefficients associated with farm price positive changes for time t are greater than the coefficients associated with the negative changes. The effect of a negative farm price change at time t-1 is significant for most cities, while the positive changes do not have a significant influence. Generally, we can conclude from the regressions that the retail price always responds faster and greater when the farm price goes up than when it is declining. The effect of farm price increase is always immediate while it takes one more month to respond to a negative farm price change. To add some power to the above conclusions we undertake a formal hypothesis tests to identify asymmetries if any.

Table 8 is used to show the estimated F-statistics associated with testing for price asymmetry when the estimation model is a single equation [i.e., model types: (B), (D), (F) or

(H)]. Table 9 is used to show the resulting χ^2-statistic when multi-equations [i.e., model type: (A), (C), (E) or (G)]. From these test results, we can see that some cities (Houston, New Orleans, Philadelphia, Hartford and St. Louis) have significant price asymmetry in the short run, in other words, the response of retail prices in the first month differ depending on the direction of farm price changes. This result is consistent with Bailey (1989). However, the test statistics of the hypothesis that the sum of coefficients for positive and negative farm price changes are equal can only be rejected for the St. Louis. This implies that the long-run effects of positive and negative price changes have equal effect on retail market for cities without a structural break using the Type (A), (B), (C) or (D) regression specifications.

For cities with a structural break, the response of retail price may be different before and after the date. The tests for Hartford and Seattle show asymmetric price transmission before certain dates but little evidence of asymmetry afterward. Among the 16 cities, seven of them show asymmetric response of retail prices to farm price changes in the first month. In contrast, in the long run, we cannot reject the hypothesis of symmetry of price transmission for these two cities.

5.2 Alternative Explanations of Pricing Asymmetry

Short-run (SR) and long-run (LR) price transmission elasticities[6] can be evaluated via the following specifications:

(13) $$\varepsilon_{POS_SR} = \frac{\Delta P_{rp}}{\Delta P_{fv}} \cdot \frac{P_{fv}}{P_{rp}} = \alpha_{10}^{+} \cdot \frac{P_{fv}}{P_{rp}} \qquad if \;\; \Delta P_{fv} > 0$$

(14) $$\varepsilon_{NEG_SR} = \frac{\Delta P_{rp}}{\Delta P_{fv}} \cdot \frac{P_{fv}}{P_{rp}} = \alpha_{10}^{-} \cdot \frac{P_{fv}}{P_{rp}} \qquad if \;\; \Delta P_{fv} < 0$$

(15) $$\varepsilon_{POS_LR} = \frac{\Delta P_{rp}}{\Delta P_{fv}} \cdot \frac{P_{fv}}{P_{rp}} = \sum_{k=0}^{m_1} \alpha_{1k}^{+} \cdot \frac{P_{fv}}{P_{rp}} \qquad if \;\; \Delta P_{fv} > 0$$

(16) $$\varepsilon_{NEG_LR} = \frac{\Delta P_{rp}}{\Delta P_{fv}} \cdot \frac{P_{fv}}{P_{rp}} = \sum_{k=0}^{m_2} \alpha_{1k}^{-} \cdot \frac{P_{fv}}{P_{rp}} \qquad if \;\; \Delta P_{fv} < 0$$

Where the subscripts *rp* and *fv* represent the retail price and farm value, respectively.

[6] Price transmission elasticity equals the percentage of absolute farm price change transmits to percentage of retail price change.

The resulting average elasticity values are shown in Table 10. The variance of the average elasticities, accounting for parameter and data variability are estimated via Delta method (Greene, 2012, p.1083-1084). For all market areas except Cleveland, the elasticities of price transmission are inelastic. We undertook t-tests of whether the elasticities between rising and falling farm price are statistically different. Surprisingly we do not find significant differences for most cities. For Minneapolis, the elasticity of positive change in farm price is larger than of negative change both in short and long-run.

Concentration ratio is calculated and shown in table 11. We calculate the ratio of top 2 to top 4 retail chains in each city based on the quantity sold per month. The sales in retail markets are quite concentrated, the quantity ratio sold by the top two retail markets are within 0.5326 to 0.9882. This result indicates that larger retail stores have market power, which may results in the asymmetry of price transmission from farm price to retail price of whole milk. Alternatively, the results need to be improved due to the limitation of data for the scanned retail prices. As noted above, the IRI dataset only collect data from chains and stores with more than 2 million dollars in sale. For each city, the average number of retail stores studied is eight, which is a relative small sample for the calculation of concentration ratio. Hence, our estimation is the upper bound of the concentration ratio. In other words, the actual result should be smaller than reported, which indicates an upper bias of the concentration ratio evaluation.

VI. Conclusions

We analyzed tests of price asymmetry according to the ECT approach for sixteen cities in the United States. Several tests are employed to obtain the regression model for each city. Empirical results suggest that in the short-run, the retail price responds differently to the increase and decrease of farm price. However, after several months, the changes in farm price can be reflected equally by the retail price.

This price asymmetry can be explained by the measure of market power. Price transmission elasticities for rising farm prices are larger than corresponding elasticities for falling farm prices. The concentration ratio (CR) is very high, the top two retail stores sold more than 50% of the total quantities of milk in all the 16 cities. For Cleveland, Minneapolis and Oklahoma City, the CR is even more than 90%.

Further investigations need to be done about distinguishing between different empirical models for asymmetric tests and measuring the adjust cost of retailers as one of the reasons for asymmetry. First, the error correction model (ECM) is used by most of researches recently. The traditional Houck method is argued to be inappropriate when variables are cointegrated. If retail and farm prices are not cointegrated, the choice between two methods are not well-defined and explanation is needed when some papers got different results when using different methods. Second, from the tests of asymmetry we can see that the speed of respond tends to be faster when farm price increase, which could be related with the adjustment cost. Data concerning inventory management is needed as proxy of analyzing adjustment cost.

References

Andrews, D.W.K. 1993. Tests for Parameter Instability and Structural Change with Unknown Change Point. *Econometrica* 61(4): 821-856.

Awokuse, T.O., and X. Wang. 2009. Threshold Effects and Asymmetric Price Adjustments in U.S. Dairy Markets. *Canadian Journal of Agricultural Economics* 57(2): 269-286

Bailey, D., and B.W. Brorsen. 1989. Price Asymmetry in Spatial Fed Cattle Markets. *Western Journal of Agricultural Economics* 14(2): 246-252.

Borenstein, S., A.C. Cameron, and R. Gilbert. 1997. Do Gasoline Prices Respond Asymmetrically to Crude Oil Price Changes? *The Quarterly Journal of Economics* 112(1): 305-339.

Brockwell, P. J. & Davis, R. A. 1991 Time series: theory and methods. *New York: Springer.*

Capps, O. Jr., and P. Sherwell. 2007. Alternative Approaches in Detecting Asymmetry in Farm-Retail Price Transmission of Fluid Milk. *Agribusiness* 23(3): 313-331.

Carmon, H. F., and R. J. Sexton. 2005. Supermarket Fluid Milk Pricing Practices in the Western United States. *Agribusiness* 21(4): 509-530.

Damania, R., and B. Z. Yang. 1998. Price Rigidity and Asymmetric Price Adjustment in a Repeated Oligopoly. *Journal of Institutional and Theoretical Economics* 154(4): 659-679

Greene, W. C. 2000. Econometric Analysis. 4th ed. *New York: Prentice Hall.*

Goodwin, B., and M. Holt. 1999. Price Transmission and Asymmetric Adjustment in U.S. Beef Sector. *American Journal of Agricultural Economics* 81(3): 630-637.

Government Accountability Office, 2004: *Dairy Industry: Information on Milk Prices, Factors Affecting Prices, and Dairy Policy Options*, Washington, DC, December

Gregory, A.W., and B. Hansen. 1996. Residual-based Tests for Cointegration in Models with Regime Shifts. *Journal of Econometrics* 10: 99-126.

Hansen, B. E. 1997. Approximate Asymptotic P Values for Structural-Change Tests. *Journal of Business & Economic Statistics* 15: 60-67

Houck, J.P. 1977. An Approach to Specifying and Estimating Nonreversible Func-tions.

American Journal of Agricultural Economics 59(3): 570-572

Hovhannisyan, V., and B. W. Gould. 2012. A Structural Model of the Analysis of Retail Market Power: The Case of Fluid Milk. *American Journal of Agricultural Economics* 94(1): 67-79

Jackson, William E., III. ‘‘Market Structure and the Speed of Price Adjustments: Evidence of Non-monotonicity.’’ Rev. Indus. Organization 12 (February 1997): 37–57.

Karrenbrock, Jeffrey D. ‘‘The Behavior of Retail Gasoline Prices: Symmetric or Not?’’ Fed. Reserve Bank St. Louis Rev. 73 (July/August 1991): 19–29.

Kinnucan, H.W., and O.D. Forker. 1987. Asymmetry in Farm-Retail Price Transmis-sion for Major Dairy Products. *American Journal of Agricultural Economics* 69(2): 285-292.

Meyer, J., and S. von Cramon-Taubadel. 2004. Asymmetric Price Transmission: A Survey. *Journal of Agricultural Economics* 55(3): 581-611.

Miller, D. J., M. L. Hayenga. 2001. Price Cycles and Asymmetric Price Transmission in the U.S. Pork Market. *American Journal of Agricultural Economics* 83(3): 551-562.

Peltzman, S. 2000. Prices Rise Faster than They Fall. *Journal of Political Economy* 108(3): 466-502.

Quandt, R. E. 1960. Tests of Hypothesis that a Linear Regression System Obeys Two Separate Regimes. *Journal of the American Statistical Association* 55(290): 324-330.

Reagan, P. B., and M. L. Weitzman. 1982. Asymmetries in price and quantity adjust-ments by the competitive firm. *Journal of Economic Theory* 27(2): 410-420

Stewart, H., and D.P. Blayney. 2011. Retail Dairy Prices Fluctuate with the Farm Value of Milk. *Agricultural and Resource Economics Review* 40(2): 201-217.

Titus O. Awokuse, and Xiaohong Wang. 2009. Threshold Effects and Asymmetric Price Adjustments in U.S. Dairy Markets. *Canadian Journal of Agricultural Economics* 57 (2009) 269–286

v. Cramon-Taubadel, S. 1998. Estimating Asymmetric Price Transmission with the Error Correction Representation: An Application to the German Pork Market. *European*

Review of Agricultural Economics 25(1): 1-18.

Ward, R.W. 1982. Asymmetric in Retail, Wholesale, and Shipping Point Pricing for Fresh Vegetables. *American Journal of Agricultural Economics* 64(2): 205-212.

Table 1. Descriptive Statistics of Whole Milk Prices at Farm and Retail Level.

City	Mean	Median	Standard Deviation	Min	Max	Coeff. Of Determ.
Retail Whole Milk Prices						
Boston	3.411	3.315	0.323	3.011	4.347	0.095
Chicago	3.066	3.036	0.276	2.587	3.715	0.090
Cleveland	3.316	3.255	0.407	2.639	4.379	0.123
Dallas	2.990	3.021	0.578	2.057	4.169	0.193
Detroit	2.718	2.675	0.291	2.035	3.428	0.107
Hartford	3.889	3.951	0.391	3.157	4.586	0.101
Houston	3.362	3.372	0.394	2.593	4.112	0.117
Milwaukee	3.270	3.370	0.424	2.544	4.006	0.130
Minneapolis	3.630	3.575	0.293	3.106	4.315	0.081
New Orleans	4.051	4.052	0.564	3.069	5.074	0.139
Oklahoma	3.503	3.398	0.459	2.744	4.458	0.131
Omaha	3.113	3.074	0.468	2.221	4.081	0.150
Philadelphia	3.837	3.771	0.523	3.066	4.820	0.136
Phoenix	2.580	2.518	0.299	2.062	3.279	0.116
St. Louis	3.616	3.662	0.286	3.092	4.412	0.079
Seattle	3.654	3.547	0.419	2.895	4.659	0.115
Coop Class I Whole Milk Prices						
Boston	1.260	1.197	0.242	0.942	1.895	0.192
Chicago	1.234	1.177	0.263	0.879	1.824	0.213
Cleveland	1.217	1.128	0.262	0.882	1.833	0.215
Dallas	1.226	1.144	0.246	0.921	1.841	0.201
Detroit	1.185	1.097	0.268	0.824	1.801	0.226
Hartford	1.251	1.188	0.242	0.934	1.887	0.193
Houston	1.282	1.196	0.249	0.972	1.893	0.194
Milwaukee	1.236	1.174	0.259	0.887	1.820	0.210
Minneapolis	1.162	1.122	0.251	0.831	1.786	0.216
New Orleans	1.352	1.244	0.282	1.001	2.090	0.209
Oklahoma	1.203	1.109	0.239	0.904	1.787	0.199
Omaha	1.160	1.085	0.257	0.822	1.774	0.222
Philadelphia	1.315	1.236	0.262	0.955	1.947	0.199
Phoenix	1.079	1.005	0.248	0.757	1.732	0.230
St. Louis	1.186	1.109	0.261	0.822	1.803	0.220
Seattle	1.064	0.989	0.247	0.742	1.712	0.232

Table 2. Augmented Dickey-Fuller (ADF) Tests for Stationarity.

		Test Statistic	
City	Levels	Retail Price	Farm Price
Boston	I(0)	-1.773	-2.594
	I(1)	-6.780**	-11.447**
Chicago	I(0)	-2.898*	-2.546
	I(1)	-6.083**	-11.085**
Cleveland	I(0)	-2.463	-2.191
	I(1)	-5.809**	-11.769**
Dallas	I(0)	-2.033	-2.299
	I(1)	-6.565**	-10.532**
Detroit	I(0)	-2.437	-2.147
	I(1)	-9.416**	-11.630**
Hartford	I(0)	-1.776	-2.579
	I(1)	-7.273**	-11.534**
Houston	I(0)	-2.089	-2.22
	I(1)	-7.033**	-10.509**
Milwaukee	I(0)	-1.674	-2.509
	I(1)	-10.854**	-10.701**
Minneapolis	I(0)	-2.473	-2.722
	I(1)	-10.193**	-10.930**
New Orleans	I(0)	-1.664	-2.077
	I(1)	-6.115**	-11.750**
Oklahoma	I(0)	-1.501	-2.262
	I(1)	-12.030**	-11.667**
Omaha	I(0)	-1.702	-2.302
	I(1)	-10.817**	-11.595**
Philadelphia	I(0)	-0.944	-2.293
	I(1)	-6.228**	-11.668**
Phoenix	I(0)	-2.934*	-2.449
	I(1)	-10.212**	-11.484**
St. Louis	I(0)	-3.266*	-2.278
	I(1)	-6.377**	-11.441**
Seattle	I(0)	-1.751	-2.449
	I(1)	-7.053**	-11.486**

Note: * = statistically significant at 5%; ** = statistically significant at 1%. I(•) = integration order
The null hypothesis is the price series are non-stationary and need to be differenced.

Table 3. Johansen's Cointegration Tests for Farm and Retail Milk Prices.

	Test Statistic	
City	λ_{trace}	λ_{max}
Boston	26.89*	22.31*
Chicago	20.02*	12.34
Cleveland	24.50*	17.53*
Dallas	11.47	8.22
Detroit	19.25*	12.18
Hartford	13.72	11.31
Houston	15.65*	9.89
Milwaukee	15.96*	12.81
Minneapolis	16.89*	9.64
New Orleans	16.82*	13.55
Oklahoma	29.29*	27.15*
Omaha	23.57*	20.99*
Philadelphia	20.09*	19.23*
Phoenix	16.59*	9.52
St. Louis	18.44*	14.23*
Seattle	21.80*	17.61*

Note: * = statistically significant at 5%; ** = statistically significant at 1%.
The null hypothesis is the two variables are not cointegrated.

Table 4. Granger Causality Tests.

City	Effect	Cause	F-statistic	P-value
Boston	Farm	Retail	2.57	0.0804
	Retail	Farm	1.30	0.2759
Chicago	Farm	Retail	2.04	0.1340
	Retail	Farm	10.53	0.0001*
Cleveland	Farm	Retail	3.80	0.0251*
	Retail	Farm	2.36	0.0988
Dallas	Farm	Retail	3.57	0.0311*
	Retail	Farm	2.26	0.1085
Detroit	Farm	Retail	0.84	0.4353
	Retail	Farm	8.07	0.0005*
Hartford	Farm	Retail	1.09	0.3390
	Retail	Farm	7.13	0.0012*
Houston	Farm	Retail	1.99	0.1407
	Retail	Farm	0.36	0.6957
Milwaukee	Farm	Retail	1.09	0.3386
	Retail	Farm	11.22	0.0000*
Minneapolis	Farm	Retail	2.90	0.0590
	Retail	Farm	1.77	0.1749
New Orleans	Farm	Retail	1.34	0.2651
	Retail	Farm	1.20	0.3048
Oklahoma	Farm	Retail	5.37	0.0058*
	Retail	Farm	1.36	0.2595
Omaha	Farm	Retail	2.66	0.0743
	Retail	Farm	1.73	0.1810
Philadelphia	Farm	Retail	2.49	0.0867
	Retail	Farm	0.55	0.5769
Phoenix	Farm	Retail	5.01	0.0081*
	Retail	Farm	6.45	0.0022*
St. Louis	Farm	Retail	0.82	0.4434
	Retail	Farm	7.47	0.0009*
Seattle	Farm	Retail	5.79	0.0040*
	Retail	Farm	1.53	0.2217

Note: * = statistically significant at 5%; ** = statistically significant at 1%.
The null hypothesis is there is no Granger causality.

Table 5. Result of Model Selection.

City	Error Correction Term	Single-Equation Approach	Structural Break	Structural Break Date	Model Specification
Boston	Yes	Yes	No	---	(B)
Chicago	No	No	No	---	(C)
Cleveland	Yes	No	No	---	(A)
Dallas	No	No	Yes	Jul-2003	(G)
Detroit	No	No	No	---	(C)
Hartford	No	No	Yes	Oct-2006	(G)
Houston	No	Yes	No	---	(D)
Milwaukee	No	No	Yes	Jan-2002	(G)
Minneapolis	No	Yes	Yes	Jan-2002	(D)
New Orleans	No	Yes	Yes	Jul-2004	(H)
Oklahoma	Yes	No	No	---	(A)
Omaha	Yes	Yes	Yes	Dec-2001	(F)
Philadelphia	Yes	Yes	Yes	Dec-2009	(F)
Phoenix	No	No	Yes	Jul-2004	(G)
St. Louis	Yes	No	Yes	Jul-2008	(E)
Seattle	Yes	No	Yes	Oct-2008	(E)

Note: * = statistically significant at 5%; ** = statistically significant at 1%.

Table 6. Model Equation Specification.

Model Type	Equation Specification
(A)	$\begin{cases} \Delta R_t = \alpha_0 + \sum_{k=0}^{M_1} \alpha_{1k}^{+} \Delta F_{t-k}^{+} + \sum_{k=0}^{M_2} \alpha_{1k}^{-} \Delta F_{t-k}^{-} + \gamma_1 ECT_t + \gamma_2 ECT_t^2 + \gamma_3 ECT_t^3 + u_t \\ \Delta F_t = \beta_0 + \sum_{k=0}^{m_1} \beta_{1k}^{+} \Delta R_{t-k}^{+} + \sum_{k=0}^{m_2} \beta_{1k}^{-} \Delta R_{t-k}^{-} + \gamma_1 ECT_t + \gamma_2 ECT_t^2 + \gamma_3 ECT_t^3 + v_t \end{cases}$
(B)	$\Delta R_t = \alpha_0 + \sum_{k=0}^{M_1} \alpha_{1k}^{+} \Delta F_{t-k}^{+} + \sum_{k=0}^{M_2} \alpha_{1k}^{-} \Delta F_{t-k}^{-} + \gamma_1 ECT_t + \gamma_2 ECT_t^2 + \gamma_3 ECT_t^3 + u_t$
(C)	$\begin{cases} \Delta R_t = \alpha_0 + \sum_{k=0}^{M_1} \alpha_{1k}^{+} \Delta F_{t-k}^{+} + \sum_{k=0}^{M_2} \alpha_{1k}^{-} \Delta F_{t-k}^{-} + u_t \\ \Delta F_t = \beta_0 + \sum_{k=0}^{m_1} \beta_{1k}^{+} \Delta R_{t-k}^{+} + \sum_{k=0}^{m_2} \beta_{1k}^{-} \Delta R_{t-k}^{-} + v_t \end{cases}$
(D)	$\Delta R_t = \alpha_0 + \sum_{k=0}^{M_1} \alpha_{1k}^{+} \Delta F_{t-k}^{+} + \sum_{k=0}^{M_2} \alpha_{1k}^{-} \Delta F_{t-k}^{-} + u_t$
(E)	$\begin{cases} \Delta R_t = \alpha_0 + \sum_{k=0}^{M_1} \alpha_{1k}^{+} \Delta F_{t-k}^{+} + \sum_{k=0}^{M_2} \alpha_{1k}^{-} \Delta F_{t-k}^{-} + \mathrm{I}_{at} \sum_{k=1}^{m} (\alpha_{2k}^{+} \Delta F_{t-k}^{+}) + \mathrm{I}_{at} \sum_{k=1}^{n} (\alpha_{2k}^{-} \Delta F_{t-k}^{-}) \\ \qquad + \gamma_1 ECT_t + \gamma_2 ECT_t^2 + \gamma_3 ECT_t^3 + u_t \\ \Delta F_t = \beta_0 + \sum_{k=0}^{m_1} \beta_{1k}^{+} \Delta R_{t-k}^{+} + \sum_{k=0}^{m_2} \beta_{1k}^{-} \Delta R_{t-k}^{-} + \mathrm{I}_{at} \sum_{k=1}^{m} (\beta_{2k}^{+} \Delta R_{t-k}^{+}) + \mathrm{I}_{at} \sum_{k=1}^{n} (\beta_{2k}^{-} \Delta R_{t-k}^{-}) \\ \qquad + \theta_1 ECT_t + \theta_2 ECT_t^2 + \theta_3 ECT_t^3 + v_t \end{cases}$
(F)	$\Delta R_t = \alpha_0 + \sum_{k=0}^{M_1} \alpha_{1k}^{+} \Delta F_{t-k}^{+} + \sum_{k=0}^{M_2} \alpha_{1k}^{-} \Delta F_{t-k}^{-} + \mathrm{I}_{at} \sum_{k=1}^{m} (\alpha_{2k}^{+} \Delta F_{t-k}^{+}) + \mathrm{I}_{at} \sum_{k=1}^{n} (\alpha_{2k}^{-} \Delta F_{t-k}^{-}) + \gamma_1 ECT_t + \gamma_2 ECT_t^2 + \gamma_3 ECT_t^3 + u_t$
(G)	$\begin{cases} \Delta R_t = \alpha_0 + \sum_{k=0}^{M_1} \alpha_{1k}^{+} \Delta F_{t-k}^{+} + \sum_{k=0}^{M_2} \alpha_{1k}^{-} \Delta F_{t-k}^{-} + \mathrm{I}_{at} \sum_{k=1}^{m} (\alpha_{2k}^{+} \Delta F_{t-k}^{+}) + \mathrm{I}_{at} \sum_{k=1}^{n} (\alpha_{2k}^{-} \Delta F_{t-k}^{-}) + u_t \\ \Delta F_t = \beta_0 + \sum_{k=0}^{m_1} \beta_{1k}^{+} \Delta R_{t-k}^{+} + \sum_{k=0}^{m_2} \beta_{1k}^{-} \Delta R_{t-k}^{-} + \mathrm{I}_{at} \sum_{k=1}^{m} (\beta_{2k}^{+} \Delta R_{t-k}^{+}) + \mathrm{I}_{at} \sum_{k=1}^{n} (\beta_{2k}^{-} \Delta R_{t-k}^{-}) + v_t \end{cases}$
(H)	$\Delta R_t = \alpha_0 + \sum_{k=0}^{M_1} \alpha_{1k}^{+} \Delta F_{t-k}^{+} + \sum_{k=0}^{M_2} \alpha_{1k}^{-} \Delta F_{t-k}^{-} + \mathrm{I}_{at} \sum_{k=1}^{m} (\alpha_{2k}^{+} \Delta F_{t-k}^{+}) + \mathrm{I}_{at} \sum_{k=1}^{n} (\alpha_{2k}^{-} \Delta F_{t-k}^{-}) + u_t$

Table 7(a). Estimation Results of Single-Equation Specifications

	Boston	Houston	Minneapolis
Regression of Farm Price on Retail Price			
Positive change in farm price at time t (ΔF_t^+)	0.348**	0.869**	0.656**
	(0.086)	(0.147)	(0.128)
Negative change in farm price at time t (ΔF_t^-)	0.445**	0.412**	0.471**
	(0.085)	(0.143)	(0.134)
Positive change in farm price at time t-1 (ΔF_{t-1}^+)	0.102	-0.054	
	(0.092)	(0.150)	
Negative change in farm price at time t-1 (ΔF_{t-1}^-)	0.055	0.409**	
	(0.086)	(0.139)	
Positive change in farm price at time t-2 (ΔF_{t-2}^+)			
Negative change in farm price at time t-2 (ΔF_{t-2}^-)			
Positive change in farm price at time t-3 (ΔF_{t-3}^+)			
Negative change in farm price at time t-3 (ΔF_{t-3}^+)			
Level error correction term (ECT)	-0.111*		
	(0.05)		
Square error correction term (ECT^2)	-0.171		
	(0.149)		
Cubic error correction term (ECT^3)	0.165		
	(0.512)		
Intercept	0.01	0.005	-0.005
	(0.008)	(0.012)	(0.011)
Model Fit and Diagnostics			
R^2	0.395	0.374	0.293
AIC	-374.447	-254.018	-232.854

Note: * = statistically significant at 5%; ** = statistically significant at 1%.
Numbers in parentheses are standard errors.

Table 7(b). Estimation Results of Single-Equation Specifications

	New Orleans	Omaha	Philadelphia
Regression of Farm Price on Retail Price			
Positive change in farm price at time t (ΔF_t^+)	0.881**	0.948**	0.719**
	(0.125)	(0.176)	(0.053)
Negative change in farm price at time t (ΔF_t^-)	0.434**	0.970**	0.438**
	(0.119)	(0.180)	(0.052)
Positive change in farm price at time t-1 (ΔF_{t-1}^+)	0.161	-0.089	0.081
	(0.127)	(0.184)	(0.054)
Negative change in farm price at time t-1 (ΔF_{t-1}^-)	0.362**	0.267	0.322**
	(0.115)	(0.175)	(0.052)
Positive change in farm price at time t-2 (ΔF_{t-2}^+)		0.073	
		(0.185)	
Negative change in farm price at time t-2 (ΔF_{t-2}^-)		0.061	
		0.171	
Positive change in farm price at time t-3 (ΔF_{t-3}^+)		(-0.123)	
		0.192	
Negative change in farm price at time t-3 (ΔF_{t-3}^+)		0.041	
		(0.168)	
Level error correction term (ECT)		-0.134	0.000
		(0.089)	(0.024)
Square error correction term (ECT^2)		-0.587*	-0.008
		(0.281)	(0.045)
Cubic error correction term (ECT^3)		-0.898	-0.199
		(0.849)	(0.139)
Intercept	0.001	0.042*	0.009
	(0.010)	(0.021)	(0.006)
Model Fit and Diagnostics			
R^2	0.467	0.489	0.794
AIC	-288.952	-181.673	-491.542

Note: * = statistically significant at 5%; ** = statistically significant at 1%.
Numbers in parentheses are standard errors.

Table 8(a). Estimation Results of Multiple-Equation Specifications

	Chicago	Cleveland	Dallas	Detroit	Hartford
Regression of Farm Price on Retail Price					
Positive change in farm price at time t (ΔF_t^+)	0.971**	1.201**	1.092**	1.056**	0.790**
	(0.146)	(0.226)	(0.168)	(0.200)	(0.083)
Negative change in farm price at time t (ΔF_t^-)	0.607**	0.747**	0.781**	0.950**	0.537**
	(0.153)	(0.219)	(0.157)	(0.199)	(0.078)
Positive change in farm price at time t-1 (ΔF_{t-1}^+)	0.158	-0.082	-0.019	0.144	0.087
	(0.149)	(0.230)	(0.171)	(0.207)	(0.085)
Negative change in farm price at time t-1 (ΔF_{t-1}^-)	0.459**	-0.191	0.397**	0.135	0.266**
	(0.145)	(0.217)	(0.153)	(0.190)	(0.075)
Level error correction term (ECT)		-0.205*			
		(0.103)			
Square error correction term (ECT^2)		0.055			
		(0.174)			
Cubic error correction term (ECT^3)		-0.597			
		(0.500)			
Intercept	-0.005	-0.014	0.006	-0.004	0.004
	(0.014)	(0.022)	(0.014)	(0.018)	(0.008)
Regression of Retail price on Farm Price					
Positive change in retail price at time t (ΔR_t^+)	0.852**	0.379**	0.572**	0.493**	1.076**
	(0.100)	(0.080)	(0.080)	(0.084)	(0.121)
Negative change in retail price at time t (ΔR_t^-)	0.218**	0.351**	0.371**	0.378**	0.968**
	(0.084)	(0.084)	(0.110)	(0.096)	(0.176)
Positive change in retail price at time t-1 (ΔR_{t-1}^+)	-0.031	0.215**	0.042	0.048	-0.089
	(0.100)	(0.080)	(0.080)	(0.085)	(0.121)
Negative change in retail price at time t-1 (ΔR_{t-1}^-)	0.036	0.054	0.156	0.095	-0.118
	(0.085)	(0.077)	(0.110)	(0.092)	(0.175)
Level error correction term (ECT)		0.041			
		(0.061)			
Square error correction term (ECT^2)		-0.008			
		(0.104)			
Cubic error correction term (ECT^3)		0.203			
		(0.296)			
Intercept	-0.023	-0.014	-0.004	-0.002	-0.009
	(0.012)	(0.016)	(0.010)	(0.014)	(0.012)
Model Fit and Diagnostics					
R^2	0.29	0.221	0.218	0.1	0.429
AIC	-516.583	-423.901	-597.641	-487.597	-817.328

Note: * = statistically significant at 5%; ** = statistically significant at 1%.
Numbers in parentheses are standard errors.

Table 8(b). Estimation Results of Multiple-Equation Specifications--Continue

Regression of Retail price on Farm Price					
Positive change in retail price at time t (ΔR_t^+)	0.844**	0.537**	0.679**	0.479**	0.963**
	(0.108)	(0.060)	(0.111)	(0.096)	(0.112)
Negative change in retail price at time t (ΔR_t^-)	0.539**	0.309**	0.563**	-0.02	0.605**
	(0.113)	(0.077)	(0.104)	(0.092)	(0.111)
Positive change in retail price at time t-1 (ΔR_{t-1}^+)	0.076	0.146*	0.084	0.087	0.052
	(0.110)	(0.063)	(0.111)	(0.105)	(0.111)
Negative change in retail price at time t-1 (ΔR_{t-1}^-)	-0.172	-0.123	0.06	-0.006	0.025
	(0.108)	(0.088)	(0.104)	(0.104)	(0.114)
Positive change in retail price at time t-2 (ΔR_{t-2}^+)	0.067			0.156	0.087
	(0.114)			(0.085)	(0.066)
Negative change in retail price at time t-2 (ΔR_{t-2}^-)	0.157			0.02	0.044
	(0.115)			(0.081)	(0.132)
Positive change in retail price at time t-3 (ΔR_{t-3}^+)	0.067			-0.032	0.586
	(0.114)			(0.100)	(0.443)
Negative change in retail price at time t-3 (ΔR_{t-3}^-)	0.157			0.025	
	(0.115)			(0.082)	
Level error correction term (ECT)		0.136**		0.012	
		(0.051)		(0.060)	
Square error correction term (ECT^2)		-0.014		0.009	
		(0.172)		(0.123)	
Cubic error correction term (ECT^3)		0.182		-0.334	
		(0.347)		(0.241)	
Intercept	-0.004	-0.024	-0.001	-0.015	-0.02
	(0.013)	(0.012)	(0.013)	(0.018)	(0.012)
Model Fit and Diagnostics					
R^2	0.272	0.214	0.248	0.243	0.32
AIC	-620.642	-610.042	-608.968	-349.656	-716.359

Note: * = statistically significant at 5%; ** = statistically significant at 1%.

Numbers in parentheses are standard errors.

Table 9. The F-statistics of the price asymmetry test.

City	Without a structural break		With a structural break		
	Short-run	Long-run	Before	After	Date
Boston	0.48	0.09	---	---	---
Houston	3.69**	0.00	---	---	---
Minneapolis	0.78	---	0.38	0.58	2001-Dec
New Orleans	4.93**	1.24	10.01**	6.58*	2004-Jul
Omaha	0.01	1.14	0.33	0.25	2001-Dec
Philadelphia	10.48**	0.17	---	---	---

Note: * = statistically significant at 5%; ** = statistically significant at 1%.

Table 10. The χ^2 statistics of the Price Asymmetry Test.

City	Without a structural break		With a structural break		
	Short-run	Long-run	Before	After	Date
Chicago	2.26	0.05	---	---	---
Cleveland	1.55	1.70	---	---	---
Dallas	1.37	0.11	0.02	0.14	1-Jul
Detroit	0.10	0.09	---	---	---
Hartford	3.67**	0.23	4.44**	2.3	Oct-06
Milwaukee	0.10	0.51	0.33	0.15	Jan-02
Ok_City	0.63	0.27	---	---	---
Phoenix	0.22	0.17	0.67	0.2	1-Dec
St. Louis	5.42**	2.86*	1.69	0.6	1-Jul
Seattle	2.13	0.00	3.89**	0.77	1-Oct

Note: * = statistically significant at 5%; ** = statistically significant at 1%.

Table 11. Elasticities of Price Transmission for Whole Milk.

City	Short-run Rising	Short-run Falling	T-statistic (pos_sr = neg_sr)	Long-run Rising	Long-run Falling	T-statistic (pos_lr = neg_lr)
Boston	0.2332	0.1984	-0.1421	0.4352	0.7131	-0.7952
	(1.6167)	(1.0118)		(1.9935)	(1.9330)	
Chicago	0.9739	1.475	0.3757	1.7148	3.2643	-0.8593
	(2.1702)	(0.7258)		(7.5778)	(12.8097)	
Cleveland	0.7112	2.609	1.0497	2.964	3.5449	-0.225
	(1.8420)	(1.2105)		(16.4386)	(11.4413)	
Dallas	0.3115	-0.5203	-0.797	-0.1354	-0.1586	0.0155
	(0.9450)	(1.6688)		(10.0698)	(5.9518)	
Detroit	3.3186	0.1353	-1.0619	4.1822	3.4823	0.1619
	(1.7968)	(0.7774)		(24.1275)	(24.7641)	
Hartford	-0.0606	0.1284	0.4331	-0.1637	0.279	-0.7764
	(2.2952)	(1.2740)		(3.8238)	(2.1689)	
Houston	1.0023	0.6112	-0.2898	2.8209	0.3617	1.2728
	(1.7254)	(1.2769)		(13.5687)	(6.3207)	
Milwaukee	-0.2911	-0.2048	0.1168	-0.2389	-0.6932	0.4417
	(1.6394)	(1.3194)		(5.8818)	(5.7294)	
Minneapolis	0.2598	-0.0688	-1.7257	0.5580	-0.1194	2.3964
	(2.8507)	(0.6705)		(1.7261)	(1.4328)	
New Orleans	-0.8236	-0.546	0.2658	-1.8364	-0.8922	-0.6162
	(1.6585)	(1.7701)		(10.5181)	(5.6174)	
Oklahoma	1.6699	-0.6615	0.2658	3.3951	-1.1102	-0.6162
	(1.4382)	(1.7171)		(24.1353)	(9.3461)	
Omaha	0.1602	-1.277	-1.0052	0.0154	-2.0467	1.0432
	(1.2016)	(1.6478)		(10.6065)	(11.7768)	
Philadelphia	0.6398	0.2116	-0.8489	0.9489	0.8446	0.1477
	(1.6164)	(1.6886)		(4.0663)	(3.8617)	
Phoenix	1.3917	-0.3813	-0.8825	2.8356	-0.6519	1.1689
	(1.6824)	(0.8318)		(21.4239)	(6.3738)	
St. Louis	-2.1948	-0.0928	0.9012	-4.6766	-0.0715	-1.2832
	(3.3472)	(0.2270)		(26.5731)	(4.2716)	
Seattle	0.5554	-0.2152	-1.4832	1.1568	-0.4337	1.9346
	(2.0704)	(1.4888)		(5.6133)	(2.8842)	

Note: * = statistically significant at 5%; ** = statistically significant at 1%.

Numbers in parentheses are standard errors.

Table 12. Concentration Ratio of Whole Milk.

City	CR2	CR3	CR4
Boston	0.6666	0.8712	0.9405
Chicago	0.6401	0.7948	0.8952
Cleveland	0.9675	0.9894	0.9971
Dallas	0.5953	0.7253	0.8129
Detroit	0.8347	0.9647	0.9840
Hartford	0.6763	0.7964	0.8908
Houston	0.6809	0.8432	0.9216
Milwaukee	0.7681	0.8797	0.9475
Minneapolis	0.9266	0.9719	0.9926
New Orleans	0.7464	0.8774	0.9512
Oklahoma	0.9882	0.9995	0.9998
Omaha	0.8995	0.9797	0.9975
Philadelphia	0.5326	0.7222	0.8228
Phoenix	0.7273	0.8942	0.9876
St. Louis	0.7673	0.9455	0.9839
Seattle	0.6271	0.7677	0.8679

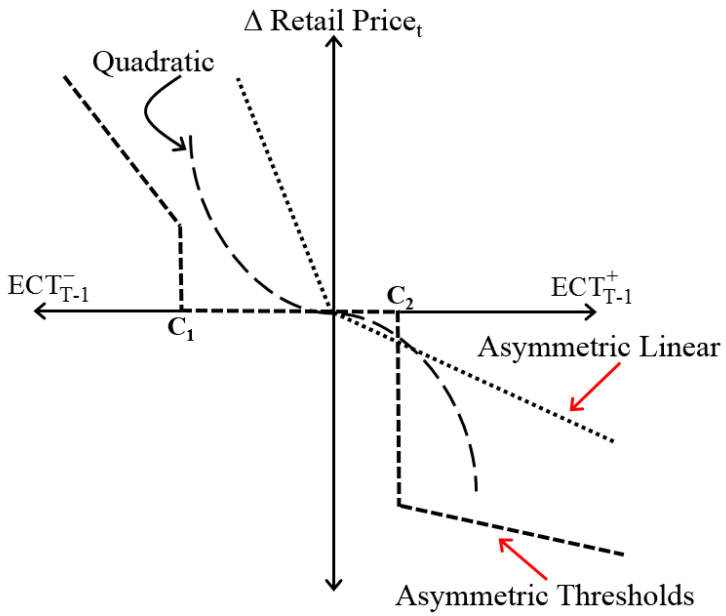


Figure 1. Types of Error Correction

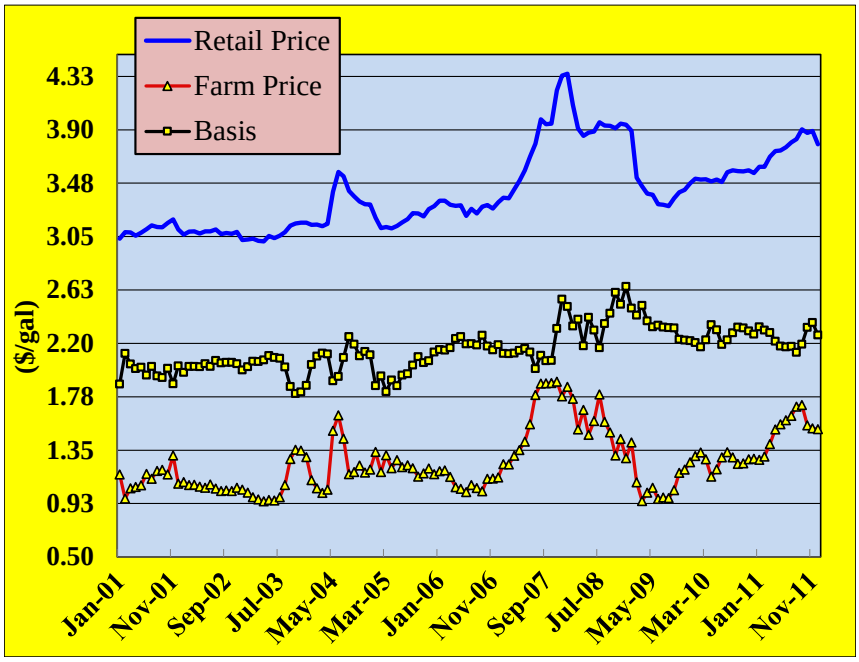
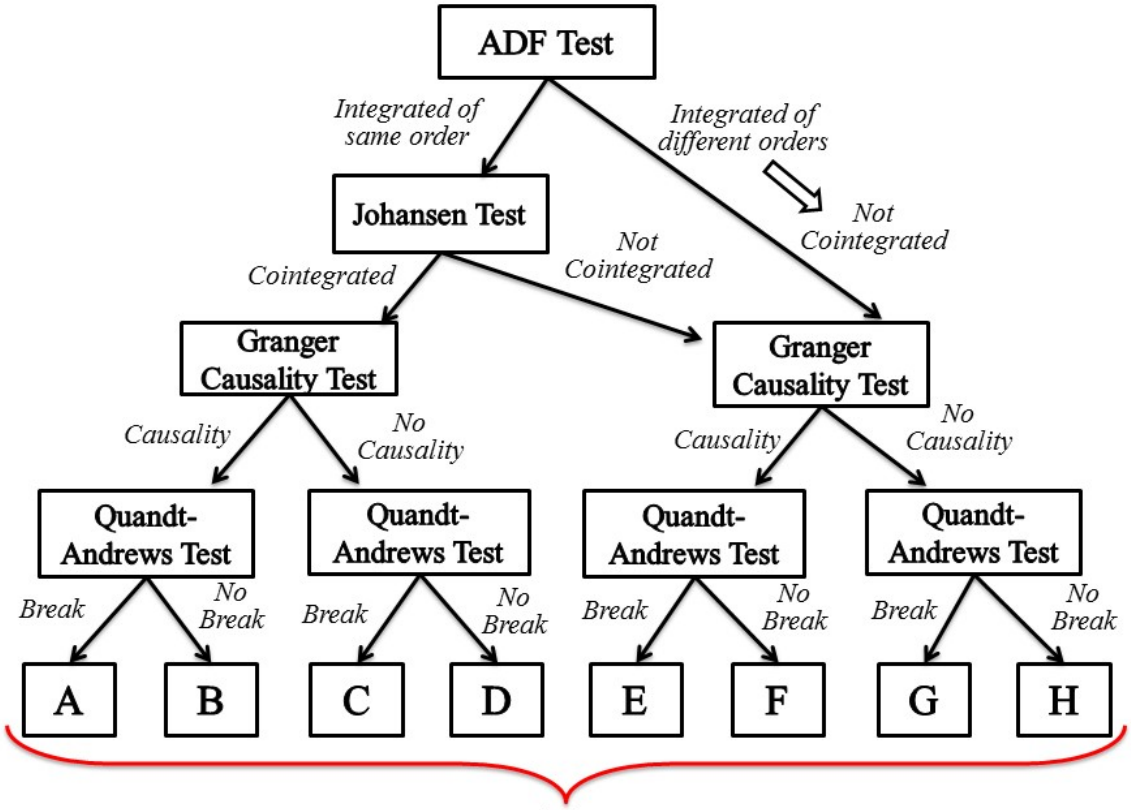


Figure 2. Comparison of Whole Milk Retail Price, Farm Price and Basis



Model Type	(A)	(B)	(C)	(D)	(E)	(F)	(G)	(H)
Error Correction Term	Yes	Yes	No	No	Yes	Yes	No	No
Single-Equation Approach	No	Yes	No	Yes	No	Yes	No	Yes
Structural Break	No	No	No	No	Yes	Yes	Yes	Yes

Figure 3. A Summary of Test Procedures Used for Model Selection