Impacts of the Northeast Dairy Compact on New England Retail Prices

Daniel A. Lass, Mawunyo Adanu, and P. Geoffrey Allen

Northeast Dairy Compact impacts were estimated for Boston and Hartford retail prices using an econometric model. Asymmetric speeds of adjustment to farm price increases and decreases were found; however, tests indicated that retail prices do return to the same level following equal farm price increases and decreases. Model forecasts suggested no structural changes occurred during the out-of-sample period, July 1996 through June 1998. Simulations with and without the Compact predicted lower retail fluid milk price impacts than actual July 1997 changes. These predicted impacts separate the effects of farm price changes on retail prices from possibly confounding effects.

The announcement of the Northeast Interstate Dairy Compact (Compact) over-order premium attracted a significant amount of media attention and encouraged a great deal of speculation about impacts of the Compact on retail fluid milk prices. When the over-order premium was instituted in July of 1997, retail prices increased as anticipated. While the furor has died, questions remain about impacts the Compact had on retail fluid milk prices and whether observed increases were warranted given the effects the Compact had on the farm milk price. In the analyses below, an assessment is provided of impacts the Compact had on retail milk prices in New England.

Analyses of impacts the Compact had on retail fluid milk prices were conducted as follows. First, we specified then econometrically estimated the relationship between the farm price of fluid milk (the Class I price) and retail milk prices in the Boston and Hartford submarkets of the New England milk market. The estimation step used the first part of the available data (the "within-sample" set and the bulk of the data). Once the relationship between farm and retail fluid milk prices was estimated, we used the second part of the data to forecast retail fluid milk prices for the period immediately prior to Compact implementation and the period immediately following. By making forecasts for retail fluid milk prices with and without the known impacts of the Compact on the farm fluid milk, we could assess the impacts the Compact should have had on retail fluid milk prices. As a check on these predicted impacts, we also compared actual and forecast prices in New England.

In the next section we describe the retail fluid milk price model we used and the hypotheses we tested. Then we discuss data collection problems. Third, estimation results are presented and implications of the results discussed. Finally, we present forecasts for the period immediately prior to and following the Compact and discuss simulations for retail fluid milk prices with and without the Compact: In the final section, we summarize and discuss some general conclusions about the relationships between the farm Class I milk price and retail prices of fluid milk.

Theory of Retail Prices

Various marketing margin models have appeared in the literature. Lyon and Thompson compared four common specifications, including the basic price mark-up model, and were unable to distinguish among them based on non-nested hypothesis tests. The mark-up model presumes that the retail fluid milk price represents the farm fluid milk price, the Class I price, plus a mark-up to account for additional processing, handling and marketing

The authors are professor, former graduate student, and professor, respectively, with the Department of Resource Economics at the University of Massachusetts, Amherst.

Analyses were conducted under contract with the Northeast Dairy Compact Commission; support from the Compact Commission is gratefully acknowledged. The authors are solely responsible for the content of this report.

costs (Heien et al. 1980). It can be represented by the following simple mathematical expression

$$(1) \qquad R_t = \alpha F_t + \sum_{i=1}^{k} \beta_i P_{it} + e_t$$

where R_t is the retail price of fluid milk in period t, F_t is the Class I price of milk in period t, and P_{it} represents processing costs. Processing costs include expenditures on labor, packaging, transportation and other processing inputs. The model suggests that the farm milk price is marked-up according to the value of processing and marketing inputs required to bring the product to the retail shelf. Included are any costs that retailers incur in marketing the product. The parameters, α and β, determine the degree to which the farm price and processing costs are included in the retail price.

Kinnucan and Forker (1987) developed an important modification of the basic mark-up model. They found that retail price responses to farm price changes were asymmetric with respect to rising versus falling farm prices. Rising farm prices were incorporated rapidly in retail prices, while retail prices were slow to adjust to decreases in farm prices. It was also found that falling farm prices were not incorporated in retail prices to the same degree as rising farm prices. Holding all other factors constant, retail prices would rise rapidly with rising farm prices, but would fall more slowly with falling farm prices. And when the farm price first increased and then fell back to its previous level, the retail price would fail to return to its previous level.

The Kinnucan and Forker results have been generally accepted and represent important considerations for anyone analyzing farm-to-retail price transmissions. In order to incorporate their modifications into the basic price mark-up model, farm prices must be separated into rising and falling categories. Rising and falling farm prices are then included in the model separately to allow for the possibility of different effects on retail prices. To account for potential differences in the speed of retail price adjustment to rising versus falling farm prices, lagged values are used in addition to current period values. The basic price mark-up model in the equation above is expanded to:

$$(2) \qquad R_t = \delta_0 t + \sum_{i=0}^{L1} \pi_i^R FR_{t-i} + \sum_{i=0}^{L2} \pi_i^F FF_{t-i} + \beta P_t + e_t;$$

where t is a time trend variable,

$$FR_t = F_1 + \sum_{i=0}^{t-2} \max(\Delta F_{t-i}, 0)$$

measures the accumulated increases in farm price up to period t, $\Delta F_t = F_t - F_{t-1}$,

$$FF_t = F_1 + \sum_{i=0}^{t-2} \min(\Delta F_{t-i}, 0)$$

measures the accumulated decreases in farm price up to period t, and P is a marketing cost index. Theory provides no guidance for the number of lags to include. The model presented in the equation above is completely general, allowing different lag lengths for rising ($L1$) and falling ($L2$) farm prices. We evaluated a number of different lag structures during our analysis and found that current price (period t) and two lagged prices (for periods $t - 1$ and $t - 2$) best fit our data.

To determine whether price asymmetry exists in retail milk pricing in the New England market, we conducted hypothesis tests on the estimated parameters. In particular, we are interested in the following hypotheses:

$$(3) \qquad H_o\colon \pi_i^R = \pi_i^F, \text{ for lags } i = 0, 1, 2;$$

$$\text{and:} \quad H_o\colon \sum_{i=0}^{2} \pi_i^R = \sum_{i=0}^{2} \pi_i^F.$$

The alternative hypotheses in each case are that the parameters, or sums of parameters, are not equal. The first hypothesis is a test of the speeds of adjustment for rising versus falling farm prices, sometimes referred to as a test for short-run price transmission asymmetry. For example, suppose the estimated parameter for current rising farm price is statistically greater than the estimated parameter for current falling farm price. Then processors will have a greater response to an increase in the current farm price than they will to a decrease in the current farm price. This result will have provided evidence that upward adjustments in retail prices due to rising farm prices occur more rapidly than do downward adjustments due to falling farm prices. The second hypothesis test will provide evidence about whether retail prices return to the same level after equal farm price increases and decreases over the period of three months. This is sometimes referred to as a test for long-run price transmission asymmetry. Rejection of either of these hypotheses, or both, will constitute evidence of asymmetric farm-to-retail price transmission in the New England market.

In addition to these asymmetry tests, we computed mean lags for rising and falling farm price effects. For rising farm prices, the formula is (Rao and Miller 1971):

$$(4) \qquad \bar{L}^R = \frac{\sum_{i=0}^{2} |\pi_i^R| \cdot i}{\sum_{i=0}^{2} |\pi_i^R|};$$

where $\bar{L}^R$ is the mean lag for rising farm prices and the maximum lag length is two months. The mean lag for falling farm prices is computed similarly. A mean lag for rising farm prices that is smaller than the mean lag for falling farm prices will provide additional evidence that the upward speed of adjustment is more rapid than the downward speed of adjustment.

The final set of measures that indicate relative speeds of adjustment are the *current* and *short-run* elasticities of retail price with respect to rising and falling farm prices. The *current* and *short-run* elasticities are computed as follows:

$$(5) \quad \varepsilon_C^R = \pi_0^R \cdot \frac{\overline{FR}_t}{\overline{R}_t}; \text{ and } \varepsilon_{SR}^R = \sum_{l=0}^{2} \left(\pi_l^R \cdot \frac{\overline{FR}_{t-l}}{\overline{R}_{t-l}} \right)$$

where $\overline{FR}_t$, $\overline{R}_t$, and $\overline{FR}_{t-l}$ are means for the respective series. The *current* and *short-run* elasticities differ in that the latter incorporate all lagged farm price effects on the retail price. *Current* and *short-run* elasticities for falling farm prices are calculated using similar formulas. In this study, the current period and one-month and two-month lags were used to calculate *short-run* elasticities.

Data

Monthly time-series data for the period January 1982–June 1998 were collected for New England. The data were divided into two samples, one for estimation and a second for forecasting. The estimation sample included data for the period January 1982 through June 1996. The sample retained for forecasting included data for July 1996 through June 1998. This allowed a two-year period for forecasting, one year prior to implementation of the Compact and one year during which it has been in effect. U.S. level data (U.S. averages) were also collected for the period January 1982 through June 1996 and were used in model specification.

A consistent series of retail price data for New England proved somewhat difficult to obtain. After discussions with economists responsible for regional Consumer Price Indexes and economists at the U.S. General Accounting Office, we chose to use the USDA Agricultural Marketing Service (AMS) retail fluid milk price series for Boston, Massachusetts, and Hartford, Connecticut, to represent retail fluid milk prices in the New England market. These data were available monthly and were provided for our estimation and forecasting periods by the AMS staff.[1] Changes in retail prices were determined from this monthly series and a variable measuring the accumulated changes was created for use as the dependent variable in estimation.

The Class I price for the New England market measured the farm price of fluid milk. The Class I price of milk was used because we are concerned only with changes in retail fluid milk prices and the Class I price represents the cost of farm milk as an input for the fluid milk processing industry.

The USDA's Economic Research Service provided marketing cost indexes.[2] These indexes represent producer price indexes for inputs used in processing and marketing. Indexes were available for labor, packaging, and transportation inputs. In addition, a composite index of these three components was provided. It is likely that these three variables will be highly collinear making it difficult to identify individual effects when all three are included in the model. By using the composite index, estimation problems may be avoided without encountering the problems associated with omitting important variables.

Data for two additional variables were also collected to test the validity of the Kinnucan and Forker specification. The consumer price index for nonalcoholic beverages (Bureau of Labor Statistics) was included to capture effects of price changes for substitutes. Quarterly expenditures for total generic fluid milk advertising were included to capture two possible effects: an additional cost of marketing fluid milk and shifts in the fluid milk demand. The generic dairy advertising data, deflated by the media cost index, were Leading National Advertisers data provided by Dr. Harry Kaiser of the Cornell Commodity Promotion Research Program. Our list of variables is almost a subset of the variables used by Vande Kamp and Kaiser (1999) in their study of the effects of advertising on the New York City fluid milk market. Since their dependent variable is quantity sold, their model is a type of marketing cost model (Lyon and Thompson 1993).

[1] Retail fluid milk prices were prices paid by consumers for the most common brand and packaging of whole milk. The prices were in $ per gallon and were provided by John Wetterau of the USDA, AMS, Dairy Programs.

[2] Howard Elitzak of the USDA, Economic Research Service, provided the market cost indexes.

Econometric Estimation and Results

Because none of the price and cost variables are non-stationary (that is, each possesses a unit root) we adopted the initial specification proposed by Houck (1977) in his note clarifying the specification and estimation of non-reversible functions. The variables for retail fluid milk prices, current and lagged Class I milk prices, consumer price index for nonalcoholic beverages, marketing cost indexes and advertising costs were first transformed into first-differences. Variables were then created that separately measured accumulated increases and accumulated decreases in farm price, using the formulas in the definitions that follow equation (2). For example, in period five of the data set, the current cumulative rising farm price variable represents the sum of all price increases for periods one through five. These accumulated increase and decrease variables allow for different effects of rising versus falling farm prices.

Ideally, sufficient data would be available to use separate samples for model specification, estimation and forecasting. We used U.S. level data for the same estimation period, January 1982 through June 1996, and the Kinnucan and Forker model to specify the lag structure for retail fluid milk prices. We used a polynomial distributed lag model and maximum likelihood methods to correct for autocorrelation.[3] After testing several different lag lengths and polynomial orders we found that a polynomial lag of length two and order two for both rising and falling farm prices best fit the data. This duplicates the results given in Kinnucan and Forker's article and requires no restrictions on the parameters of the polynomial (so that an Almon lag specification, for example, is no different from an unrestricted model of the same lag order). The one substantive difference between our results and those reported by Kinnucan and Forker is that we failed to reject the hypothesis of long-run price transmission symmetry.

We assumed that the lag structure for the retail fluid milk price models for Boston and Hartford were the same as the U.S. for estimation. Based on results from misspecification tests, the regional models appear to be well specified. Calculated values of the Ljung-Box Q-statistic of 41.95 and 20.85 for Boston and Hartford, respectively, fell well short of the critical χ^2 [36, 0.05] value of 50.96, indicating that the residuals are white noise. Jarque-Bera tests lent empirical support for normality of the residuals for the Boston model, though not for the Hartford model.

The measures of farm-to-retail price asymmetry were calculated from the estimated econometric model results. Tables 1 and 2 present summary results for the Boston and Hartford models. The estimated coefficients for the Boston model show that retail fluid milk prices rise rapidly in response to a current period rise in the farm price (π_0^R = 3.513). There then appears to be an adjustment that occurs in subsequent periods after the initial rapid increase. However, the estimated coefficients for one-period and two-period lags were not statistically significant. The response to a current period decline in farm price is of a much smaller magnitude (π_0^F = 1.075). In fact, the greatest impact of falling farm prices occurs with a one-period lag (π_1^F = 1.584). Only the current period effect is statistically different from zero for rising farm prices, while the current period and one-period lag effects are statistically different from zero for falling farm prices.

Hypothesis tests were conducted to determine whether single period rising effects were different from single period falling effects. If the single period effects are not statistically different, then there is no statistical difference between the retail price time-paths of adjustment to farm price changes. As indicated in table 1, the direction of the price change matters for both the current period effect and the one-period lag effect. The positive calculated t-statistic indicates the current period rising coefficient is statistically greater than the current period falling coefficient, while the negative t-statistic for the one-period lag indicates the opposite. The two-period lag coefficients were not statistically different. Thus, there is empirical evidence in the Boston retail fluid milk price series that adjustments to rising farm fluid milk prices are much more rapid than adjustments to falling farm fluid milk prices.

The second aspect of farm-to-retail price asymmetry to consider is the net effect after adjustments to both rising and falling farm prices have had time to be incorporated into retail milk prices. In other words, if farm prices rose, but subsequently fell to their previous level, would retail prices also rise and then return to their previous level? In the model, retail milk price is affected by current values for accumulated rising and falling farm prices, rising and falling farm prices lagged one period, and rising and falling farm prices lagged two periods. Thus, this hypothesis test requires a comparison of the sum of the three rising coefficients to the sum of the three falling coefficients. The

[3] The regression model and autoregressive parameters were estimated simultaneously using the maximum likelihood methods of *SAS*'s autoregression procedure.

Table 1. Results for the Boston Econometric Retail Price Model, 1982–1996

	Estimates	Hypotheses	Test Statistic
Rising Farm Price Coefficients:	$\pi_0^R = 3.513\ (0.74)^a$	H_o: $\pi_0^R = \pi_0^F$; H_a: $\pi_0^R \neq \pi_0^F$	2.654*
	$\pi_1^R = -0.491\ (0.94)$		
	$\pi_2^R = -0.326\ (0.77)$	H_o: $\pi_1^R = \pi_1^F$; H_a: $\pi_1^R \neq \pi_1^F$	−1.927*
Falling Farm Price Coefficients:	$\pi_0^F = 1.075\ (0.48)^a$		
	$\pi_1^F = 1.584\ (0.49)$	H_o: $\pi_2^R = \pi_2^F$; H_a: $\pi_2^R \neq \pi_2^F$	−0.390
	$\pi_2^F = 0.033\ (0.46)$		
Sum of Rising Coefficients:	2.697 (0.615)	H_o: $\sum_{l=0}^{2} \pi_l^R = \sum_{l=0}^{2} \pi_l^F$	0.006
Sum of Falling Coefficients:	2.691 (0.585)	H_a: $\sum_{l=0}^{2} \pi_l^R \neq \sum_{l=0}^{2} \pi_l^F$	

[a]Numbers in parentheses are standard errors.
*Statistically different from zero at the 5% level of significance.

sums of the estimated rising coefficients and the estimated falling coefficients for Boston retail prices are reported in table 1. As can be observed, there is little difference between these sums. A test of the hypothesis confirms what is apparent; there is no statistical difference between the sums of rising and falling price coefficients. Thus, while the time-path of adjustment to rising farm prices is more rapid than the time-path for falling farm prices, the net retail price effect is not statistically different from zero for equal increases and decreases in farm prices.

The same analyses (table 2) were conducted using Hartford retail fluid milk price data. Individual estimated coefficients indicated only current period effects were statistically different for rising versus falling farm prices. The two current period coefficients indicated that retail prices respond more rapidly to increases in the farm price than to decreases in the farm price. The current period effect of a rising farm price was statistically greater than the current period effect of a falling farm price. Estimated coefficients for one-month and two-month lags were not statistically different. We again compared the sum of rising coefficients to the sum of falling coefficients to determine whether retail prices return to the same level after equivalent farm price increases and decreases. While the rising sum of farm price coefficients appears greater than the falling sum, we cannot conclude that the sums of the coefficients were statistically different. Thus, we also find that aspect of farm-retail price asymmetry to be absent in the Hartford area as well.

We also examined the two other measures of asymmetry: mean lags, and current and short-run elasticities, defined in equations (4) and (5). The mean lags were calculated as a weighted lag-length for rising and falling farm milk prices. Recall that for rising farm milk prices, most of the impact occurs in the current period, which places the heaviest weight on the lag of zero. For falling farm

Table 2. Results for the Hartford Econometric Retail Price Model, 1982–1996

	Estimates	Hypotheses	Test Statistic
Rising Farm Price Coefficients:	$\pi_0^R = 2.246\ (0.58)^a$	H_o: $\pi_0^R = \pi_0^F$; H_a: $\pi_0^R \neq \pi_0^F$	2.399*
	$\pi_1^R = -0.742\ (0.73)$		
	$\pi_2^R = 1.004\ (0.60)$	H_o: $\pi_1^R = \pi_1^F$; H_a: $\pi_1^R \neq \pi_1^F$	−1.317
Falling Farm Price Coefficients:	$\pi_0^F = 1.132\ (0.37)^a$		
	$\pi_1^F = 0.361\ (0.38)$	H_o: $\pi_2^R = \pi_2^F$; H_a: $\pi_2^R \neq \pi_2^F$	0.818
	$\pi_2^F = 0.415\ (0.36)$		
Sum of Rising Coefficients:	2.509 (0.501)	H_o: $\sum_{l=0}^{2} \pi_l^R = \sum_{l=0}^{2} \pi_l^F$	0.859
Sum of Falling Coefficients:	1.908 (0.472)	H_a: $\sum_{l=0}^{2} \pi_l^R \neq \sum_{l=0}^{2} \pi_l^F$	

[a]Numbers in parentheses are standard errors.
*Statistically different from zero at the 5% level of significance.

Table 3. Comparison of Mean Lags for Rising and Falling Farm Price Effects

	Mean Lags		Test	
	Rising	Falling	Statistic	Conclusion*
Boston	0.424	0.613	−0.353	Fail to Reject
Hartford	0.505	0.624	−0.488	Fail to Reject

*Conclusions for a t-test of the null hypothesis: rising mean lag equals falling mean lag. Tests conducted at the 5% level of significance.

milk prices, stronger effects occur for lags of one-month and two-months. The mean lags (table 3) for rising farm prices are closer to zero showing that the strongest effects occur during the current period for rising farm fluid milk prices in contrast to falling farm prices. Statistical tests revealed no differences in the values of the mean lag; however, it is important to remember that the current period coefficients provide weights for the lag of zero.

Current elasticities represent the immediate (current period) effects of a farm milk price increase or decrease on the retail milk price. The *short-run* elasticities represent the sum of the current and lagged effects. These elasticities, measures of the percentage change in retail price given a percentage increase or decrease in farm price, provide another way of considering current and aggregate effects of farm price increases and decreases. The elasticities presented in table 4 show that the immediate effect of a farm price increase is greater than the immediate effect of a farm price decrease, while over time the aggregate effects of increases and decreases are approximately the same.

Analyses of the farm-retail price relationship in both the Boston and Hartford markets lead us to conclude that asymmetry does exist. The asymmetry appears as a rapid upward response to an increase in the farm price, while retail prices fall more gradually when farm prices decrease. However, after three periods (three months time), retail prices would return to the same level given equivalent increases and decreases in the farm price of fluid milk.

Table 4. Current and Short-Run Retail Price Elasticities of Farm Price Changes

	Rising Elasticities		Falling Elasticities	
	Current	Short-Run	Current	Short-Run
Boston	0.457	0.351	0.140	0.350
Hartford	0.300	0.333	0.150	0.253

Forecasting and Simulation Results

Two final aspects of this farm-retail fluid milk price analysis to be considered are out-of-sample forecasting using the econometric model and simulating retail prices with and without the Compact. By forecasting out-of-sample, we can determine whether there appears to have been a change in the farm-retail fluid milk price relationship. Through simulations with and without the Compact, we can predict effects of the Compact on retail fluid milk prices. The data retained for out-of-sample forecasting included the months July 1996, through June 1998. The Compact was instituted in July 1997. Thus, there are 12 months prior to and 12 months following Compact implementation for forecasting.

Retail milk price forecasts are presented in figure 1 for Boston and in figure 2 for Hartford. The forecast graphs show that the model did well in predicting retail milk prices. Forecasting performance was measured by the mean absolute percentage error (MAPE). The within sample MAPE for Boston was 0.88% and 0.70% for Hartford. The out-of-sample MAPE values were 1.16% and 0.80% for Boston and Hartford, respectively.[4] The models forecasted well, both before and after Compact implementation.

There appear to be immediate responses to the Compact that the models did not predict completely. Institution of the Compact represents a unique policy change, identified in the model by the change in farm price. Participants in the fluid milk markets knew perfectly the magnitude of the farm price change, and knew that the change would be permanent, at least for the duration of the Compact. The model was estimated using historic data where farm price changes were not necessarily known with certainty and were of uncertain duration. Participants in the New England fluid milk markets may have initially reacted differently had the permanence of the change been viewed as uncertain. Thus, differences between actual prices and forecasts in July 1997 likely reflect perfect anticipation of a well-known policy change by processors and retailers. There was no need for processors and retailers to form expectations of changes or to make any additional adjustments during subsequent months. While these individuals had such information, the econometric model was built with data reflecting uncertainty, and required

[4] The forecast graphs belie the small forecast errors because of the restricted range of the vertical axes. The choice of the vertical axis exaggerates the magnitudes of the errors.

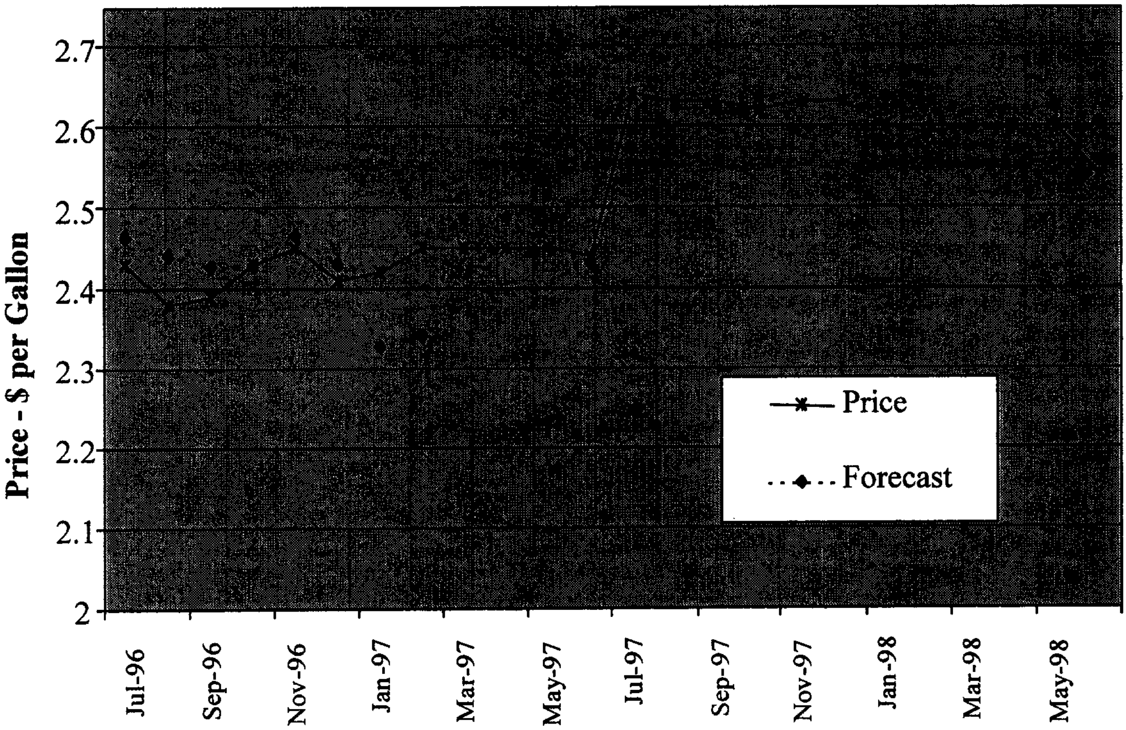

Figure 1. Boston Retail Prices and Forecasts.

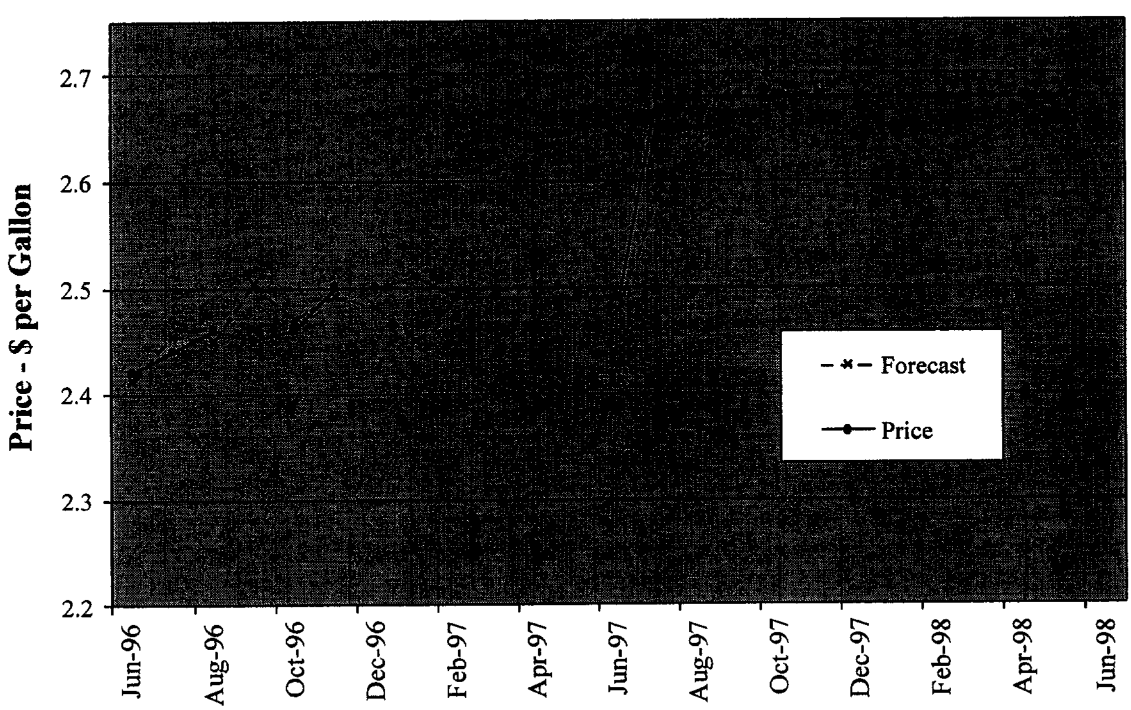

Figure 2. Hartford Retail Prices and Forecasts.

some time to fully reflect adjustments to changes that occurred. The Compact stabilized farm prices; current and lagged changes diminished to zero as the farm fluid milk price became an administered price. The autoregressive error correction component of the model also subsides and the model forecasts converge to the stable values of the actual retail prices. The fact that the econometric model converges on the actual prices suggests that the retail behavior did not depart permanently from the process that was predicted using historical data.

One method of estimating impacts the Compact had on retail milk prices is to simulate retail prices "with" and "without" the administered farm fluid milk price in place. The differences in simulations then represent estimates of the impacts of setting a farm-level fluid milk price floor. The forecasts presented in figures 1 and 2 cannot be used to measure effects of the Compact on retail prices because they include effects of changes in processing costs and the error correction component. The error correction component of the model incorporates any differences not explained by the model into the following (next month's) forecast. Retail prices did increase in July 1997 in response to the farm level price increase caused by the Compact. In forecasting with the Class I price of milk rather than the price established by the compact, there is a large initial error for July 1997, the month the Compact was instituted. The error correction component of the model then incorporated that large error into subsequent forecasts. Thus, forecasts "with the Compact" and "without the Compact" converge in just one month due to the error correction component of the model.

Our solution is to develop predictions or simulations "with the Compact" and "without the Compact" based on only the structural components of the model. These simulations are based on values for the independent variables of the model and estimated parameters listed above. Values for the marketing cost variable are the same in both simulations. Differences between the "with Compact" and "without Compact" simulations are thus due only to different farm prices of milk for fluid uses. Thus, differences in simulations throughout the forecast period are due to differences between the Compact administered fluid milk price (i.e., the Zone 1 fluid price was set at $16.94 beginning July 1997) and the Class I price announced by the Market Administrator for the New England Market Order.

The differences between simulations "with the Compact" and "without the Compact" are detailed in table 5. The greatest Compact effects were predicted for the first month of the Compact, July 1997, for Boston and the third month of the Compact, September 1997, for Hartford. The effects declined after that and are closely associated with the over-order premium. The average predicted monthly Compact effect was 6.9 cents per gallon for Boston and 5.7 cents per gallon for Hartford. Figures 3 and 4 provide a graphic depiction of the association between the over-order premiums and simulated Compact effects. Numeric measures of these associations are provided by correlation coefficients. The estimated correlation coefficients were 0.98 and 0.91 for Boston and Hartford, respectively. A simple regression model relating the Compact effect to the over-order premium implies that a one-dollar increase in the over-order premium would lead to a 5.9 cent per gallon increase in Boston retail prices during the 12 months the compact was in effect. Simple regression estimates for Hartford implied that a one-dollar increase in the over-order premium would lead to a 3.7 cent per gallon increase in retail price.

Table 5. Over-Order Premiums and Predicted Compact Effects on Retail Prices

Year and Month	Over-Order Premium ($/hundredweight)	Predicted Compact Effects (¢/gallon) Boston	Hartford
1997: July	3.00	17.5	11.8
August	2.96	17.3	8.8
September	2.84	15.1	13.5
October	1.63	6.7	8.2
November	0.91	2.9	6.5
December	0.87	4.2	4.9
1998: January	0.74	3.7	3.0
February	0.41	1.6	1.6
March	0.45	2.1	1.9
April	0.38	1.9	1.0
May	0.89	3.0	2.3
June	1.69	6.5	4.3

Summary and Conclusions

The Kinnucan and Forker model fit the monthly retail fluid milk price data well for both Boston and Hartford and specification tests indicated both models were well specified. Forecasts and forecast evaluations suggest that the model predicted retail fluid milk prices well. A key feature of the analysis was investigation of asymmetric responses of retail fluid milk prices to farm fluid milk price changes. Farm-to-retail price asymmetry was investigated in two ways. First, the effects (estimated coefficients) of rising versus falling farm prices for the current period, a one-month lag and a two-month lag were

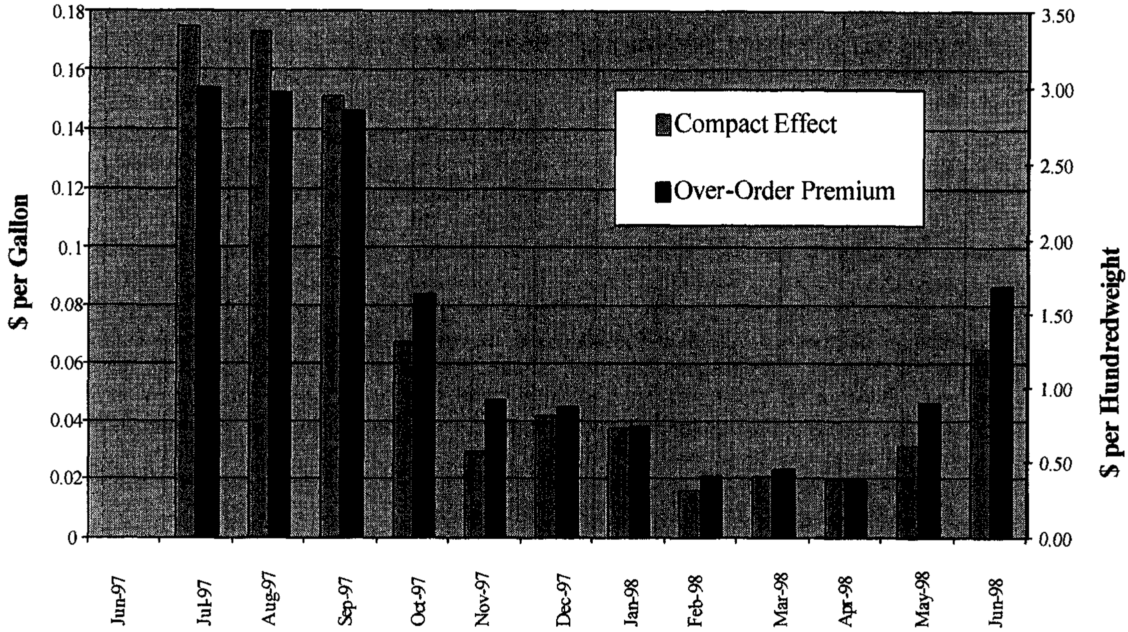

Figure 3. Compact Effects on Boston Retail Prices.

compared using t-tests. These individual parameter tests supported the existence of this form of asymmetry in both Boston and Hartford. In particular, the current period effect of rising farm fluid milk price was greater than the current period effect for falling farm fluid milk price for both Boston and Hartford.

A second aspect of asymmetry is the net effect of equal farm price increases and decreases. To determine whether retail fluid milk prices will eventually return to the same level after equal increases and decreases in farm prices, the sums of rising and falling farm price coefficients were compared. These tests indicated that retail prices

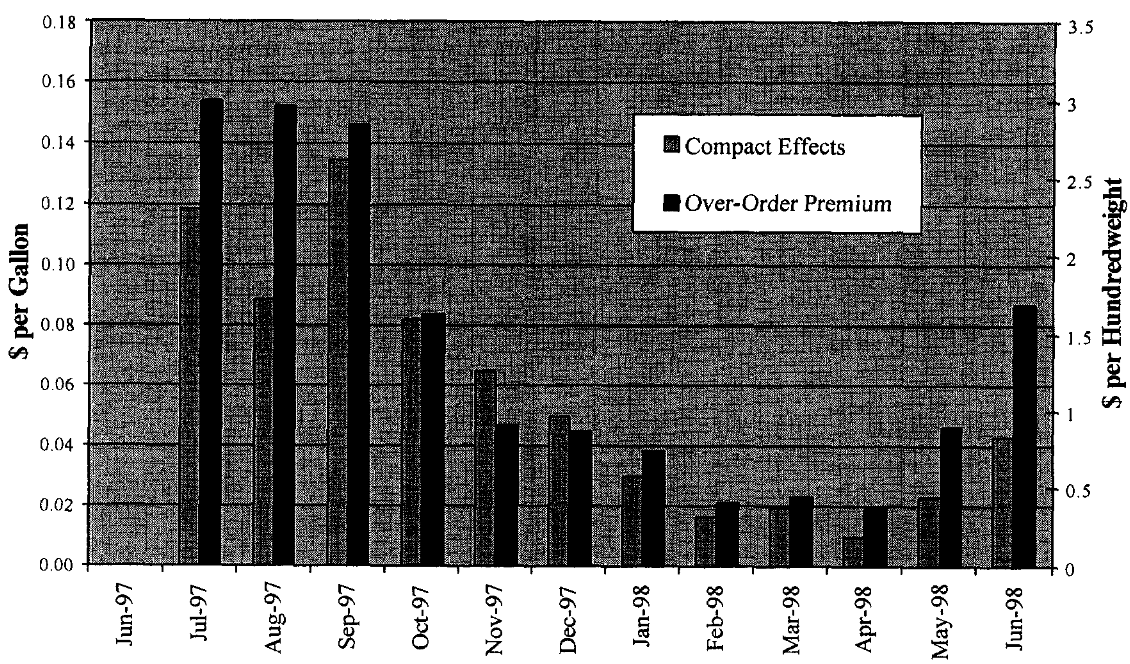

Figure 4. Compact Effects on Hartford Retail Prices.

do return to the same level upon allowing adequate time for the upward and downward adjustments. Thus, empirical evidence suggests that this second aspect of price asymmetry does not exist in Boston and Hartford.

Forecasting was a second key element of the analysis. A number of observations were saved for forecasting retail fluid milk prices "out-of-sample" including 12 months prior to and 12 months following institution of the Compact. The model forecasted actual prices well during this period; mean forecast errors were 1.16% and 0.80% for Boston and Hartford, respectively. The ability of the model to forecast actual retail fluid milk prices suggests no structural change in the farm-to-retail price relationship during the out-of-sample period, July 1996 through June 1998.

The model was also used to simulate retail prices "with" and "without" Compact. The simulations suggest that retail prices would not rise as much as had been anticipated. However, it is important to remember that these simulations are based upon a model developed to measure and forecast farm-to-retail fluid milk price relationships that was estimated using historic data. The value of the simulations depends upon evaluation of the model as a predictive tool and continuation of the historic farm-to-retail fluid milk price relationships that have existed over a long period. Measures of model fit and performance suggest the Boston and Hartford models were well specified and that they predicted farm-retail price relationships accurately.

We find that the Compact does affect retail fluid milk prices by establishing a floor under the farm fluid milk price. Increases in farm fluid milk prices that occurred after the Compact was instituted in July 1997 were transmitted to the retail fluid milk prices. Actual prices increased by $0.20 in Boston and $0.19 in Hartford in July 1997 and then there were no subsequent increases indicated by the AMS data used in this study. In fact, by June 1998 retail prices had declined in both Boston and Hartford. Compact retail fluid milk price impacts predicted by the model were somewhat less than the actual in July 1997 and diminish thereafter due to increases that occurred in the Class I price of milk in New England. These reported predicted impacts are important estimates of Compact effects because they isolate the farm fluid milk price effects on retail fluid milk prices from the possibly confounding effects of changes in processing costs.

References

Heien, D.M. 1980. "Markup Pricing in a Dynamic Model of the Food Industry." *American Journal of Agricultural Economics* 62:10–18.

Houck, J.P. 1977. "An Approach to Specifying and Estimating Nonreversible Functions." *American Journal of Agricultural Economics* 59:570–572.

Kinnucan, H.W., and O.D. Forker. 1987. "Asymmetry in Farm-Retail Price Transmission for Major Dairy Products." *American Journal of Agricultural Economics* 69:285–292.

Lyon, C.C., and G.D. Thompson. 1993. "Temporal and Spatial Aggregation: Alternative Marketing." *American Journal of Agricultural Economics* 75:523–536.

Rao, P.P., and R.L. Miller. 1977. *Applied Econometrics.* Belmont, California: Wadsworth Publishing Co.

U.S. Department of Labor, Bureau of Labor Statistics. "Consumer Price Indexes." *http://stats.bls.gov/cpihome.htm.*

Vande Kamp, P.R., and H.M. Kaiser. 1999. "Irreversibility in Advertising-Demand Response Functions: An Application to Milk." *American Journal of Agricultural Economics* 81:385–396.