Modeling the Farm-Retail Price Spread for Beef

Michael K. Wohlgenant and John D. Mullen

A new model for the farm-retail price spread, which accounts for both farm supply and retail demand changes, is introduced. This model is applied to beef, and its empirical performance relative to the markup pricing formulation is evaluated using nonnested testing procedures. The results are consistent with theory and indicate the markup pricing model is misspecified.

Key words: beef, marketing margins, markup pricing, nonnested testing.

In recent years the real farm price of beef has declined despite a secular decline in beef production. This suggests demand as well as supply changes are important in explaining price changes. Farm-level demand for beef is influenced by changes in both consumer demand and the farm-retail price spread for beef. This paper focuses on factors affecting the price spread by estimating and testing alternative empirical specifications of the farm-retail price spread for beef.

A common approach to modeling price spread behavior is to assume the price spread is a combination of both percentage and constant absolute amounts (Waugh; George and King). This suggests an empirical specification in which the price spread is related to retail price and marketing input prices. This modeling approach has been applied to beef by Freebairn and Rausser, Arzac and Wilkinson, and Brester and Marsh. As emphasized by Gardner (p. 404) the problem with this approach is that the relationship between farm and retail prices can be depicted accurately if changes occur solely in supply or demand, not both. Because demand as well as supply changes appear to be important for beef, an alternative approach to modeling price spread behavior seems desirable.

As is demonstrated below, relating the price spread to industry output and marketing input prices where both prices are deflated by retail beef price allows simultaneous changes in demand and supply conditions. Hence, this model, referred to as the relative price model, is more theoretically appealing. While still similar in many respects, neither the relative price model nor the George and King formulation is a special case of the other, so nonnested econometric testing procedures are used. Out-of-sample forecast tests also are employed to test the adequacy of the new specification. Overall, the results indicate superiority of the relative price model over the markup pricing specification.

Theoretical Considerations

The relative price spread model can be derived from an industry-wide specification of derived demand by processors for quantity of the farm output. Assuming the farm product is predetermined with respect to price from year to year because of biological lags in the production process, derived demand for the farm output is written in price dependent form as

$$(1) \qquad P_f = f(Q, P_r, C),$$

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where P_f is the price of the farm output, Q is the quantity of the agricultural commodity processed, P_r is the price of the retail product, and C is a vector of marketing input prices (wage rates, transport costs, etc.). Neoclassical theory of the firm implies demand for a factor of production is invariant to proportionate changes in all input and output prices (Varian, chap. 1). This means equation (1) can be written in terms of relative prices as

$$(2) \qquad P_f/P_r = f(Q, 1, C/P_r) = g(Q, C/P_r).$$

Heien (p. 128) calls equation (2) the "farm-retail margin." This equation shows the theoretical determinants of the farm-retail price ratio. Increases in farm-level output and increases in relative marketing costs would be expected to lower the farm-retail price ratio. To obtain a specification for the farm-retail price spread note that when farm price is measured in the same units as the retail product that the relative price spread is by definition equal to one minus the relative farm price. Thus, using equation (2), the specification for the relative price spread is

$$(3) \qquad (M/P_r) = 1 - g(Q, C/P_r) = h(Q, C/P_r)$$

or, in terms of the absolute spread, as

$$(4) \qquad M = P_r h(Q, C/P_r),$$

where $M = P_r - P_f$ is the farm-retail price spread.[1]

In constrast to the markup pricing model, this model indicates that there is no fixed relationship between the price spread and retail price. In general, the relationship between the prices will change as output and relative marketing input prices change. This formulation is consistent with the theory of food price determination put forth by Gardner. It suggests that shifts in retail demand and farm supply have two possible avenues of influence on the farm-retail price spread: quantity of output and retail price. Increases in output and increases in relative marketing costs lead to a higher relative price spread. Because shifts in both demand and supply can cause output and retail price to change, a complete analysis of the price spread is only possible through analyzing the complete set of market behavior equations. The present paper is primarily concerned with specification of the structural equation defining the linkage between farm and retail prices.

An alternative way to obtain equation (4) is to view the spread as the price of a bundle of marketing services. On this interpretation, firms would be expected to provide marketing services to the point where the marginal value of these services (M) equals marginal cost. That is,

$$(5) \qquad M = k(Q, C),$$

where $k(\cdot)$ is the marginal cost function of marketing services. The marginal cost function is homogenous of degree one in input prices (Varian), implying $k(Q, C) = (1/t)k(Q, tC)$ for all $t > 0$. With $t = (1/P_r)$, this yields an equation of exactly the same form as (4).

The foregoing analysis suggests an alternative specification for the price spread relation of the same form as equation (5) but with both M and C deflated by some general price index such as the consumer price index. Such a specification of marketing margin behavior is prevalent in the literature (e.g., Buse and Brandow). The choice between (4) and (5), therefore, can be thought of as a choice between relative and real price specifications for price spread behavior.

Empirical Specifications and Nonnested Testing Procedures

Based on the previous theoretical considerations, three empirical specifications are hypothesized for the farm-retail price spread for beef. These are:

$$(6) \qquad M_t = a_0 + a_1 P_{rt} + a_2 IC_t + \epsilon_{1t},$$

$$(7) \qquad M_t = b_1 P_{rt} + b_2 P_{rt} Q_t + b_3 IC_t + \epsilon_{2t}, \text{ and}$$

$$(8) \qquad M_t = c_0 + c_1 Q_t + c_2 IC_t + \epsilon_{3t},$$

where M_t is the farm-retail price spread for beef, cents per pound (retail price of choice beef minus retail equivalent of farm price net of by-product value), P_{rt} is the retail price of choice beef (¢/lb.), IC_t is an index of marketing costs for beef, 1967 = 100 (simple average of index of earnings of employees in packing plants, and producer price index of fuels and related products and power), and Q_t is per capita quantity of beef produced (million pounds of beef, carcass weight, divided by civilian population in millions to remove trend growth in output). Price data (M_t, P_{rt}, and IC_t) are

[1] The farm price is assumed to be net of by-product values.

deflated by the consumer price index.[2] Equation (6) is the markup pricing hypothesis augmented by the index of marketing costs. Equations (7) and (8) are linear specifications for the relative price spread formulation (4) and the real price spread formulation (5), respectively.

Note that equation (7) does not contain an intercept. The reason for this can be seen by comparing equation (3) with equation (4). Specifically, the theory underlying this specification suggests that the price spread relation is homogenous of degree 1 in input and output prices. Clearly, equations (3) and (4) do not produce identical empirical specifications since, if the error term in one of these equations is assumed to be homoscedastic, it must be heteroscedastic in the other. Specifying the relative price spread hypothesis as equation (7) has the advantage that the comparison with equations (6) and (8) leads to easily identifiable nonnested hypotheses (see, e.g., Quandt).

Because no one specification for price spread behavior is a special case of the other, nonnested testing procedures need to be employed. A number of tests have been proposed (see Godfrey and Pesaran). The simplest of these tests is the J-test proposed by Davidson and MacKinnon. This test can be implemented as follows. Suppose the null nypothesis is the markup pricing model (6) and the alternative hypothesis is the relative price spread model (7). Consider the compound regression model

$$(9)\quad M_t = (1 - \lambda)(a_0 + a_1P_{rt} + a_2IC_t) + \lambda\hat{M}_{2t} + \epsilon_t = a_0' + a_1'P_{rt} + a_2'IC_t + \lambda\hat{M}_{2t} + \epsilon_t,$$

where $\hat{M}_{2t}$ is the predicted value of M_t from the regression model (7). Under the null hypothesis (6), the value of λ is zero; that is, the relative price model can explain none of the variation in price spread not already accounted for by the markup model. Davidson and MacKinnon (also see Pesaran) show that one may validly test whether $\lambda = 0$ by estimating equation (9) and employing a conventional t-test. The J-test can be used also to test the truth of a hypothesis against several alternatives at once. For example, to test (6) against both (7) and (8), one would estimate the compound model consisting of the right-hand-side variables in (6) and the predicted values $\hat{M}_{2t}$ and $\hat{M}_{3t}$ from (7) and (8) and then test the significance of these predicted values using a conventional F-test.

Godfrey and Pesaran present monte-carlo results which indicate the J-test has low power for small samples. They propose two additional tests which seem to have good small-sample properties.[3] These tests are an adjusted Cox-type test ($\tilde{N}_o$-test) and Wald-type test (W-test). As with the J-test, computed values for the $\tilde{N}_o$ and W-tests can be compared with the tabled t-value with the appropriate degrees of freedom. Formulas for these test statistics are not presented here in order to save space; they can be found in Godfrey and Pesaran (sec. 2).

Econometric Results and Hypothesis Testing

Equations (6)–(8) were estimated with U.S. annual time-series data covering the period 1959–83, a total of twenty-five observations. These results, together with equation (7) with an intercept included, are reported in table 1. All parameter estimates have the expected signs. The fact that the intercept in equation (7) is not significantly different from zero provides some support for the relative price specification. However, by the usual statistical criteria and consistency with a priori expectations, it is difficult to choose between these models. The only model that appears somewhat suspect is equation (8), which has a substantially lower R^2 and a low Durbin-Watson statistic.[4]

Table 2 presents pair-wise nonnested tests of each of the three models. Here, H_1 through H_3 correspond to models (6)–(8), respectively. Each group of three rows relates to a particular hypothesis being tested. The first element of each column is the value of the J-test, the second element is the value of the $\tilde{N}_o$-test, and the third element is the value of the W-test. Although the main interest is in testing the truth of H_1 against H_2 and H_3, all pair-wise tests are presented because the test results can

[2] Sources for beef data are USDA *Livestock and Meat Statistics* and *Livestock and Poultry Outlook and Situation.* Other data were obtained from the *Economic Report of the President* and USDL *Employment and Earnings of the United States.*

[3] Davidson and MacKinnon (pp. 783–84) discuss another test based on an F-test from estimating a compound model which includes the regressors from both the null and alternative hypothesis. This test is not employed here because it yields the same results as the J-test for the three specifications considered.

[4] The models were also estimated with a linear trend, to account for technical change in meat processing, and lagged retail price, to account for price leveling (Parish). However, in all cases these variables were statistically insignificant.

Table 1. Econometric Estimates of Alternative Specifications of the Farm-Retail Price Spread for Beef, 1959–83

Model	Intercept	Explanatory Variables				Statistics	
		P_{rt}	$P_{rt}Q_t$	Q_t	IC_t	R^2	D-W[a]
M_t, eq. (6)	5.524 (4.861)[b]	.199 (.051)			.084 (.013)	.72	2.03
M_t, eq. (7)		.183 (.032)	$.783 \times 10^{-3}$ $(.316 \times 10^{-3})$		.083 (.011)	NA[c]	NA
M_t, eq. (7) with intercept	4.699 (4.424)	.189 (.052)	$.757 \times 10^{-3}$ $(.316 \times 10^{-3})$		.079 (.012)	.78	2.33
M_t, eq. (8)	19.229 (4.134)			.049 (.040)	.079 (.016)	.56	1.30

[a] Durbin-Watson statistics.
[b] Standard error of the coefficient.
[c] NA—not applicable.

be sensitive to the ordering of the null and alternative hypothesis (see Davidson and MacKinnon, p. 783). All entries in the table can be compared with the tabled value for the two-sided t-test with twenty-two degrees of freedom. At the 5% significance level, this critical value is 2.074.

The pair-wise tests in table 2 indicate rejection of both H_1 and H_3 but nonrejection of H_2 relative to the other hypotheses. In only one case (H_1 vs. H_3) do the test results yield ambiguous conclusions. In this case, the J-test indicates rejection but the $\tilde{N}_o$- and W-tests indicate nonrejection. This is consistent with the findings of Godfrey and Pesaran, who find a tendency for the J-test to reject when it should not.

As noted earlier, the J-test can be used also to test each hypothesis against the other two jointly by estimating a compound model consisting of the regressors of the null hypothesis and the predicted values of the dependent variables for the two alternative hypotheses. These test statistics, which are computed using the conventional formulas for F-tests, yield values for H_1 of 2.73, for H_2 of .54, and for H_3 of 10.33. In each case, the F-value has two numerator and twenty denominator degrees of freedom. With a 5% critical value of 3.49, this suggests rejection of only H_3. While this result may appear favorable for the markup pricing hypothesis, the F-test gives disproportionate weight to H_3, which based on the results in table 2 appears to be an inferior alternative to either H_1 or H_2. In other words, the relevant comparison seems to be between H_1 and H_2. The pair-wise results in table 2 clearly indicate a preference for the relative price spread model.

Out-of-Sample Forecast Tests

Equations (6) and (7) also were subjected to out-of-sample forecast tests. Recursive residual analysis, described by Galpin and Hawkins, was used to assess the extent of parameter instability over the sample period. Recursive residuals are derived by sequentially deleting observations from the model and by using the estimated parameters from the reduced sample to generate year-ahead forecast errors. Under the null hypothesis that the model specification is correct, these (standardized) recursive residuals will be normally distributed. A

Table 2. Pair-wise Nonnested Tests for H_1 through H_3

Null Hypothesis	Alternative Hypothesis		
	H_1	H_2	H_3
H_1		2.39 −3.23 −2.92	2.39 −.60 −.59
H_2	1.06 −.37 −.37		1.06 −1.19 −1.06
H_3	4.66 −7.72 −6.10	4.66 −5.93 −3.47	

Note: H_1–H_3 corresponds to equations (6)–(8), respectively. The first element in each column is the J-test, the second is the $\tilde{N}_0$-test, and the third is the W-test. See text for explanation of these tests.

total of twenty-two recursive residuals were generated for each model, and normality was tested using the Shapiro-Wilk statistic. For each model the null hypothesis of normally distributed errors was not rejected at a 10% significance level.

The CUSUM test suggested by Brown, Durbin, and Evans was also applied. Under the null hypothesis, the sum of the recursive residuals (standardized by the standard deviation of the sample) is expected to follow a random walk around zero. While none of the CUSUM plots for the three models closely resembled a random walk, neither did they cross the "critical boundaries"; hence, the plots do not indicate a process of gradual structural change.

On these criteria the relative price and George and King models seem to be correctly specified and the residuals have the desirable properties. However, other aspects of the behavior of the recursive parameters and residuals gave cause for concern about the stability of the models. The normal probability plots did not pass through the origin, and the CUSUM plots and residuals suggested that both models were systematically overpredicting the price spread from around 1973. Dufour argues that "structural changes will be indicated by tendencies to either overpredict or underpredict" (p. 34). A plot of the recursively estimated parameters also suggested structural change around 1966. The structural shift seemed most pronounced for the George and King model and involved the parameter on retail price changing from negative to positive and that on marketing input prices becoming much smaller. This type of behavior results either from the presence of outliers in the base period for the recursive estimation or from some form of model misspecification. Because the parameters changed most in the late 1960s it would seem that misspecification was more likely the problem.[5]

The out-of-sample forecasting performance of both models was compared using a mean-squared-error (MSE) test developed by Ashley, Granger, and Schmalansee. In this test, the difference in the out-of-sample forecast errors between the two models was regressed on the sum of the forecast errors from the two models. When both intercept and slope are positive, a conventional F-test can be employed to test that both models have equal forecasting performance (see Ashley, Granger, and Schmalansee, p. 1155). This test was applied to ten out-of-sample forecast errors (1974–83) for equations (6) and (7) using parameters estimated with data over the period 1959–73. The computed F-value was 67.76, indicating strong rejection of the null hypothesis that the two MSE's are the same.[6] The relative price spread model has a much smaller MSE and hence is preferred on this criterion.

Discussion

Overall, the test results indicate rejection of the markup pricing specification compared to the relative price spread specification. The reason for the difference in empirical performance appears to be significant shifts in retail demand as well as farm supply, which are reflected both in retail price and quantity and, therefore, the relationship between retail and farm price. The inferior performance of the markup pricing model is consistent with the conclusion by Gardner (p. 406) that, with both supply and demand shifts, no markup pricing relationship can depict accurately the relationship between retail and farm price.

It is interesting to compare the results of this study with those of Buse and Brandow, who also included retail price and quantity in their margin specifications. In contrast to the findings reported here, their results indicated volume had a small and insignificant effect on the farm-retail price spread for beef. This was true for both quarterly and annual data. While Buse and Brandow's finding of an insignificant relationship could be a function of the time period used (1921–41, 1947–57), it also could be an artifact of the specific functional form they used. In particular, their model related the price spread linearly to retail price and quantity

[5] The two price spread models were also subjected to within-sample structural tests. Plots of the price spread suggested that structural change may have occurred in the early 1970s coincident with rising oil prices. Moschini and Meilke found some evidence of structural change in demand for meat around this time. Hence, the significance of slope and intercept dummy variables for the period 1973 to 1983 for the three models was tested using an F-test. The relative price model was structurally stable. The intercept dummy variable for George and King model was significant at the 5% level but not the 1% level after failing to reject the hypothesis that the slope parameters were stable.

[6] Even if we take one-half the computed F-value for the test value, as Ashley, Granger, and Schmalansee recommend, we still strongly reject the null hypothesis that the MSEs are the same.

without an interaction term between price and quantity. The results in table 2 clearly indicate this interaction term is preferred to a linear quantity term. The implication of this finding is that quantity affects margin behavior mainly through its effect on the percentage markup—a larger volume leads to a higher percentage markup and vice versa.

A related implication of the relative price spread model concerns estimation of derived demand elasticities. George and King show that with their specification of price spread behavior (and assuming fixed input proportions), derived demand elasticities can be obtained as the product of retail price elasticities of demand and elasticities of price transmission between retail and farm prices. As emphasized by Hildreth and Jarrett (p. 110), this relationship obtains only when quantity does not appear in the processor behavioral equation. Otherwise, the formula must be modified to account for the influence of retail price on output quantity. In that case, the correct derived demand formula to use is

$$\frac{\eta \cdot e}{1 - (\eta/\epsilon)},$$

where η is the price elasticity of retail demand, e is the elasticity of price transmission, and ϵ is the price elasticity of the retail supply function.[7] Using this formula, derived demand elasticities from the relative price spread model can be compared with those derived from the George and King model. At the sample mean prices and quantities of 90.49¢ and 100.36 pounds, the elasticity of price transmissions for the George and King and relative price spread models are .75 and .81, respectively. The retail supply elasticity for the relative price spread model at the sample means is estimated to be 9.4. Assuming a retail demand elasticity of −.6, we obtain elasticities of derived demand for the George and King and relative price spread models of −.45 and −.46, respectively. Thus, despite the inferior statistical performance of the George and King model, this model and the relative price spread model yield almost the same derived demand elasticities at the sample means. The reason for the similarity in estimates is the large retail supply elasticity at this data point.

Despite the similarity in derived demand elasticities with these two models, the choice of an econometric model for price spread behavior will depend upon its ultimate use. If the model is to be used to obtain derived demand elasticities, then the George and King model might suffice. However, if the model is intended to be used in policy applications relating to shifts in both retail demand and farm supply, then preference would be for the relative price model. The reason for this is that the relative price model can account for shifts in supply and demand which have different consequences for the relationship between farm and retail prices. For example, a policy aimed at reducing farm output supply (which would cause quantity to fall and retail price to rise) would have different consequences for the relationship between farm and retail prices than a policy aimed at increasing retail demand (which would cause both quantity and retail price to rise). The relative price spread model proposed here is an improvement on previous specifications of margin behavior precisely because it can account for simultaneous changes in farm output supply and retail demand.

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[7] On the assumption the retail product is produced in fixed proportions with the farm product, the retail supply elasticity can be derived by differentiating equation (1). See Hildreth and Jarrett for details.

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