Journal of Agricultural and Applied Economics, 31,1(April 1999):1–13

Estimating Market Power and Pricing Conduct in a Product-Differentiated Oligopoly: The Case of the Domestic Spaghetti Sauce Industry

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ABSTRACT

This paper develops a simultaneous-equations panel data econometric model to obtain point estimates of market power and pricing conduct in a representative product-differentiated, oligopolistic food market. The importance of this class of markets is recognized given its prevalence in the food and fiber system, especially for final consumer food products. The $1.3 billion domestic spaghetti sauce industry is featured. Although the results indicate firms exert limited market power, a portion of this power is derived from tacit price collusion. A higher degree of price collusion was found among brands within a market segment than between segments.

Key Words: market power, oligopoly, pricing conduct, product differentiation.

In the empirical marketing literature, researchers have relied heavily on the assumption of product homogeneity. The assumption has been incorporated in both temporal and spatial analyses of market power and firm conduct in output and input markets. Representative papers for various agricultural industries include meat processing (Schroeter; Koontz, Garcia, and Hudson; Azzam and Schroeter), fruit and vegetable processing (Wann and Sexton), dairy (Liu, Sun, and Kaiser), general food processing (Holloway), and tobacco (Appelbaum). There are several explanations for the continued emphasis on homogeneous product models in this realm of research. First, the assumption appears to be consistent with product characteristics in the markets investigated. A second reason stems from theoretical elegance, as these models are more parsimonious and allow for the direct estimation of a market power parameter (Varian). Finally, aggregate commodity data are typically available in secondary sources.

A broad class of imperfectly competitive markets, product-differentiated oligopolies, has been under-represented in the literature despite its prevalence in the food and fiber system especially for final consumer food products. The Connor et al. study is commonly cited as evidence that diversified food manufacturers depart from marginal cost pricing. However, many of the SIC codes in that study represent the aggregation of structural oligopolies in which firms sell differentiated products to consumers. For example, SIC code

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20860 represents carbonated soft drinks and contains nine brands and numerous associated products in the regular (e.g., non-diet) segment alone (Cotterill). The aggregation inherent in five-digit SIC codes falsely leaves the impression that the product homogeneity assumption is appropriate for every food market.

Our understanding of empirical market power and firm conduct in differentiated oligopolies is clearly in its infancy. Current analyses lag theoretical developments in this area (Tirole), and the empirical work is neither as extensive nor refined as homogeneous product industry research. Business strategists and antitrust policymakers would benefit from a broader base of empirical evidence. Thus, the empirical objective of this paper is to measure market power and pricing conduct at the brand level in a representative oligopolistic output market.

The domestic spaghetti sauce industry was chosen as an intriguing and representative processed agricultural product to study for various reasons. The industry is worth $1.3 billion annually and so is an important component of the food and fiber system. It is a structural oligopoly in which products, highly differentiated by brand, flavor, and size, are manufactured and sold.[1] Also, several features of this market may lead to heightened consumer price sensitivity. Spaghetti sauce is a durable good since its shelf life exceeds the time period between price changes (Tirole). Thus, a product becomes an intertemporal substitute for itself because consumers can stockpile when it is on sale. The products are sold in a common store location because they are substitutes in use, so consumers can make comparisons across products and brands. Another dimension of this industry is the role of transportation costs in pricing decisions. These products are relatively heavy and need to be transported from remote manufacturing locations to spatially dispersed selling markets.

This paper builds on several previous studies addressing this class of markets to enlarge the pool of empirical knowledge. Contributions are made to the literature in terms of model specification, where we control for merchandising, use a weekly time series, and empirically disentangle pricing conduct associated with rivalrous behavior from pricing conduct related to a firm's shipping costs. More generally, this New Empirical Industrial Organization (NEIO) study departs from the traditional NEIO paradigm as it identifies, in part, the source of market power and tracks a single industry through both time and space (Bresnahan). Finally, in the process of obtaining feasible error-components three-stage least-squares (EC3SLS) estimators of market power and pricing conduct with methods developed by Kinal and Lahiri, several errors in the econometrics literature were identified and corrected (Vickner and Davies).

Related Literature

Examples of empirical market power research relaxing the assumption of product homogeneity are relatively sparse. Baker and Bresnahan pioneered the use of residual demand analysis, applying it to the brewing industry. Their study was one of the seminal NEIO papers to model the attributes of a product-differentiated market, where each firm faces its own demand curve rather than an industry demand curve. Moreover, prices in these demand systems are not only endogenous but are also interdependent (Shapiro). Their practical approach to market delineation had several undesirable features. First, the estimated conduct parameter, the residual demand elasticity, was a composite of the price elasticity of demand and the conjectural variation elasticity, so individual effects could not be identified. Second, as discussed by Froeb and Werden, the ability to quantify the dynamic behavior of both consumers and purveyors is limited by the use of temporally aggregated time series.

Liang investigated the ready-to-eat breakfast cereal industry using a different model formulation to separately estimate both components of the residual demand elasticity. She used a linear demand system and corresponding set of linear price reaction functions to endogenize prices and quantify pricing conduct

[1] Due to data and model size limitations, the analysis is restricted to brand differentiation.

at the firm level. A restrictive pairwise comparison of firms and aggregate data limited the results of the study.

More recently, Cotterill analyzed the carbonated soft drink industry using an alternative NEIO model. A theoretically consistent, linear approximate Almost Ideal Demand System (LA/AIDS) model was used to represent the demand-side of the market (Deaton and Muellbauer). Consistent with the approach taken by Liang, the first-order conditions of a general price conjectural variations model were to endogenize price and quantify strategic interaction among the firms. Cotterill's study, while a generalization and extension of previous approaches, is not without problems. First, merchandising (e.g., in-store or point-of-purchase advertising) was not controlled for properly, possibly due to data availability issues. Three mutually exclusive merchandising measures were used to characterize the selling conditions: (a) the product was highlighted in a feature ad or newsprint flier, (b) the product was put on some form of display, and (c) the product was simultaneously placed in a feature ad and on a display. Scanner data suppliers such as Information Resources, Inc. (IRI), A.C. Nielsen, and Efficient Market Services measure both the percent of brand's sales and the percent of all commodity volume (ACV) made on a given merchandising condition. Cotterill used the former. Because the ACV measure quantifies the percent of stores in a geographic market maintaining one of the three merchandising conditions, we consider it a more appropriate demand shift variable.[2]

Cotterill also treated real expenditure as an exogenous variable in the model when it is clearly a function of the endogenous variables. Finally, data aggregation was an issue in the study. Quarterly time series observations were used because a less aggregate series was not available. However, both consumer and strategic behavior is affected on a weekly basis, as pricing and merchandising activities are managed in that short time frame (Kinsey and Senauer). Because twelve weeks of data were averaged into one quarterly observation, the ability to accurately measure consumer behavior, firm conduct, and, hence, market power, was diminished.

Model Development

Departures from the product homogeneity assumption result in several knotty modeling issues. First, industry output Q (*e.g.*, $\Sigma_{i=1}^{n} q_i = Q$ *for firms* $i = 1, \ldots, n$) and an overall industry demand curve do not exist as the output of each firm is measured in incommensurate units. The convenient closed form expressions for market power based on Q are immediately rendered invalid and, instead, each firm faces an individual demand curve for its own product. Demand is then a function of its own price, prices of imperfect substitutes, expenditure, and demand shift variables. Firm level demands must necessarily be obtained in a systems context to be used subsequently in the calculation of market power.

Consistent with the approach taken by Cotterill, the demand-side of the econometric model is based on the LA/AIDS model. Using the AIDS model is appropriate as it utilizes dollar market shares, thus measuring demands in commensurate units across brands. The linear approximation of the AIDS model appears to be reasonable as well in that Stone's linear price index performs well relative to competing price indexes, especially under conditions of price multicollinearity (Alston, Foster, and Green). The latter issue is important in markets where price collusion or followship is present. Since the panel data model in this study is used in part to explain pricing spatially and temporally, Moschini's transformation of the price series is inappropriate.[3] The demand system is given by

$$(1) \quad S_{kit} = \alpha_k + \sum_{l=1}^{5} \gamma_{kl} \log P_{lit} + \beta_k \log\left(\frac{x}{P*}\right)_{it} + \sum_{d \in D_k} \delta_{kd} m_{dit} + \eta_{ki}^{(s)} + u_{kit}^{(s)} \qquad \forall\, k$$

where

[2] Because stores vary in size, the stores used in the measure are weighted by sales or ACV.

[3] Moschini's transformation requires each price series to be a price index, thus eliminating the price level across the selling markets.

$$(2)\qquad \log P_{it}^{*} = \sum_{k=1}^{5} S_{kit}\log P_{kit}$$

$$(3)\qquad \sum_{k=1}^{5}\alpha_k = 1 \qquad \sum_{k=1}^{5}\beta_k = 0 \qquad \sum_{k=1}^{5}\gamma_{kl} = 0 \quad \forall\, l$$

$$(4)\qquad \sum_{l=1}^{5}\gamma_{kl} = 0 \quad \forall\, k$$

$$(5)\qquad \gamma_{kl} = \gamma_{lk} \quad \forall\, k \neq 1.$$

Equation (1) represents the share equations for each of the five brands included in the spaghetti sauce market. For example, brand l's market share (s_{kit}) equation is a function of own price ($\log p_{1it}$), competitive prices ($\log p_{2it}, \ldots, \log p_{5it}$), real expenditure ($\log(x/P_{it}^{*})$), and miscellaneous shift variables relevant to brand 1 ($\Sigma_{d\in D1}\, \delta_{1d} m_{dit}$), such as own merchandising, competitive merchandising, time trends, seasonality, and holiday effects (e.g., the set D_1). Each of the k share equations is stochastic and has two error components. The first is the random unobservable individual effect ($\eta_{ki}^{(s)}$) that only varies spatially across markets in the study. The second is the usual stochastic error ($u_{kit}^{(s)}$) that varies over both time and space. The subscript k and superscript s indicate the error components term for brand k's share equation. Brands $k = 1, \ldots, 5$ represent, respectively, Ragu, Prego, Hunt's, Classico, and Healthy Choice. The cross-sectional unit subscript i runs from selling markets 1 to 10, while the time-series subscript t runs from weeks 1 to 104. Equation (2) is Stone's linear price index. Equations (3) to (5) represent the usual adding-up,[4] homogeneity, and symmetry restrictions, respectively. Since the term x in Stone's price index represents expenditure in the spaghetti sauce market, not income, it is treated as endogenous. It is specified in the system as an identity. Thus, real expenditure in the spaghetti sauce market is endogenous and is replaced by an instrument in the estimation.

The second problem associated with relaxing the assumption of product homogeneity is that prices in the demand system are not only endogenous to the system, but are also interdependent given the strategic interaction among purveyors in the industry. Following Liang and Cotterill, consistent parameter estimates are recovered through the construction of a price reaction function for each price present in the demand system. These are given by

$$(6)\qquad \log p_{kit} = \mu_k + \sum_{l\neq k}\phi_{kl}\log p_{lit} + \lambda_k \log c_{ki} + \sum_{r\in R_k}\theta_{kr}\nu_{rit} + \eta_{ki}^{(p)} + u_{kit}^{(p)} \qquad \forall\, k.$$

For example, brand 1's price ($\log p_{1it}$) is a function of competitive prices ($\log p_{2it}, \ldots, \log p_{5it}$), observable transportation costs ($\log c_{1i}$), and miscellaneous shift variables relevant to brand 1 ($\Sigma_{r\in R_1}\, \theta_{1r}\, \nu_{rit}$), such as time trends, seasonality, and holiday effects (e.g., the set R_1). Because data for firm-specific manufacturing costs are not available, the intercept μ_k is used to capture its effect. Each price reaction function has two error components similar to the demand system equations, where the subscript k and superscript p indicate that these are error components for brand k's price reaction equation.

Unlike the Liang and Cotterill models, this specification of the price reaction function disentangles pricing conduct associated with rivalrous behavior of competing firms from pricing conduct related to each firm's shipping costs. Factors influencing the former are captured by the term $\Sigma_{l\neq k}\, \phi_{kl} \log p_{lit}$ where the sum captures all rivals' pricing conduct. The price reaction elasticities, ϕ_{kl}, quantify the percentage change in brand k's price given a 1% increase in brand l's price. A positive price reaction elasticity implies tacit price collusion and an upward sloping price reaction function. The latter is a necessary condition for a Nash equilibria in a static game of differentiated Bertrand competition (Shapiro).

The transportation cost term $\log c_{ki}$ is the product of the time invariant shipping distance

[4] In our model, the expenditure shares do not sum exactly to one as the δ_{kd} parameters are not restricted to sum to zero, hence preventing the covariance matrix from becoming singular. The usual remedy of dropping a share equation in the estimation to ensure the adding-up property holds is not possible here as it would not preserve the entire specification of endogenizing the shares and prices. We are unaware of a fully consistent procedure to achieve this.

(e.g., distance from the brand's manufacturing location to a given selling market) and a monthly spot market fuel price. Cotterill did not encounter this issue as local bottlers serviced each selling market in his model. The time subscript t is omitted in log c_{k_t} because the cost series does not vary weekly; however, the series is not quite time-invariant as it changes with monthly fuel costs. Since it could be argued that firms do not purchase fuel in a spot market, the spatial component of pricing is also measured as shipping distance only in a second specification of the model. The results of both models are compared in the Results section. The λ_k parameters, or transportation cost elasticities, quantify the percent change in brand k's price given a 1% increase in the cost of transportation. For example, a positive transportation cost elasticity may imply basing-point pricing conduct (Greenhut, Norman, and Hung).

Data Description

The market level panel data set, assembled by IRI, was collected for 104 weeks from May 15, 1994 to May 5, 1996 for 10 selling markets spatially dispersed throughout the United States. For each brand, week, and market, data were collected for six separate measures—unit sales, average price per unit, expenditure, and three in-store merchandising variables. The merchandising measures control for the effect of feature ads and displays individually and jointly on a brand's market share. Given the availability of the data, the relevant product market was narrowly defined and so excludes information regarding complementary products such as pasta, Parmesan cheese, or breadsticks. Other demand shift variables, such as the seven calendar holiday dummies, three seasonality dummies, and time trends, were used to augment the brand data. Data on media advertising were not available in this analysis. Demographic variables were tested in the model, but were found to be statistically insignificant.[5] The distance, measured in miles, between a brand's manufacturing location and each selling market was estimated using the 1997 *Rand McNally Road Atlas* mileage chart. The average national retail gasoline price series measured in cents per gallon was taken from Standard and Poor's *Current Statistics* publication. A summary of the descriptive statistics for selected continuous variables in the econometric model is given in Table 1.

Empirical Results

Given the characteristics of the model presented in equations (1) to (6) and the need to generalize the results to other selling markets not present in the study, an EC3SLS estimator was chosen (Cornwell, Schmidt, and Wyhowski). With respect to the latter rationale, the 10 markets represent a sample of selling markets so an error-components model is required to make inferences about the population. The results from the EC3SLS estimator are also compared to a one-way fixed-effects three-stage least-squares (FE3SLS) estimator. The system is identified with respect to order and rank conditions (Bhargava).[6] The matrix of instruments was constructed according to Hausman and Taylor and maintains Baltagi's exogeneity properties; parameter estimates were obtained using generalized least-squares (Kinal and Lahiri). In the process of obtaining feasible EC3SLS estimators of market power and pricing conduct with computational methods developed by Kinal and Lahiri, several errors in the econometrics literature were identified and corrected (Vickner and Davies). The system weighted R^2 value for the FE3SLS

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[5] This result is not unexpected. The firm are astute marketers and have designed marketing plans with these demographic factors in mind. In the current model specification, the between version of the firms' marketing control variables captures the demographic effects sufficiently across selling markets, removing their explanatory power (Cornwell, Schmidt, and Wyhowski).

[6] In the EC3SLS model, the instrument set contains 15 merchandising variables, a linear time trend, and five transportation cost series. In the FE3SLS model, the instrument set contains the 15 merchandising variables, a linear time trend, and the fixed effects.

Table 1. Descriptive Statistics of Selected Continuous Variables in Econometric Model[a]

Variable	Mean	Standard Deviation	Coefficient of Variation	Minimum	Maximum
Market Share:[b]					
Ragu	35.62	9.12	25.61	15.52	71.90
Prego	35.58	8.73	24.55	13.00	64.45
Hunt's	5.54	2.52	45.53	0.35	19.00
Classico	16.94	8.55	50.49	2.05	43.52
Healthy Choice	6.32	2.19	34.70	2.24	20.83
Price:[c]					
Ragu	1.86	0.20	10.95	1.23	2.47
Prego	1.97	0.19	9.62	1.35	2.49
Hunt's	1.07	0.09	8.30	0.55	1.35
Classico	2.48	0.16	6.39	1.71	2.92
Healthy Choice	1.98	0.16	8.32	1.52	2.51
Log of Real Expenditure	12.20	0.73	6.01	10.70	13.89
Feature Merchandising:[d]					
Ragu	22.77	21.59	94.78	0.00	93.28
Prego	17.52	18.77	107.13	0.00	92.03
Hunt's	4.68	9.39	200.78	0.00	68.44
Classico	10.06	16.01	159.11	0.00	100.00
Healthy Choice	8.15	16.01	196.52	0.00	100.00
Display Merchandising:[d]					
Ragu	20.08	13.89	69.14	0.00	79.64
Prego	11.99	10.45	87.15	0.00	59.73
Hunt's	3.28	4.64	141.38	0.00	39.71
Classico	3.56	5.15	144.60	0.00	37.84
Healthy Choice	2.54	4.98	196.48	0.00	31.39
Feature & Display Merchandising:[d]					
Ragu	13.33	16.18	121.32	0.00	85.42
Prego	8.30	11.88	143.08	0.00	63.09
Hunt's	1.35	3.95	293.52	0.00	41.43
Classico	2.97	7.59	255.35	0.00	62.60
Healthy Choice	1.33	4.32	324.27	0.00	42.65
Fuel Price:[e]	115.07	4.24	3.68	108.00	132.30

[a] Calculations based on 1,040 observations in the panel data set.
[b] Dollar market share (percent).
[c] Dollars per 28-ounce equivalent unit.
[d] Percent of all commodity volume (ACV).
[e] Cents per gallon.

model was 73.51%. Given the estimation technique used for the EC3SLS model, a fit statistic is not available.

Homogeneity and Symmetry Restrictions

A standard F-statistic was employed to execute the tests of the linear homogeneity and symmetry restrictions in both estimators. In the FE3SLS model, we failed to reject the five homogeneity restrictions and the ten symmetry restrictions. In the EC3SLS model, we failed to reject homogeneity in the Ragu, Prego, and Healthy Choice demand equations. Also in the error-components model, we failed to reject four of the 10 symmetry restrictions.

Table 2. Point Estimates of Uncompensated Price Elasticities[a]

	Ragu	Prego	Hunt's	Classico	Healthy Choice
FE3SLS Model					
Ragu	−3.84***	1.35***	0.24***	0.86***	0.39***
Prego	1.38***	−3.41***	0.12***	0.72***	0.25***
Hunt's	1.52***	0.76***	−5.59***	1.84***	0.47***
Classico	1.78***	1.47***	0.60***	−5.74***	0.82***
Healthy Choice	2.16***	1.36***	0.40***	2.20***	−7.26***
EC3SLS Model					
Ragu	−3.92***	1.17***	−0.27**	1.12***	0.80***
Prego	1.28***	−3.16***	0.38***	0.85***	−0.15*
Hunt's	0.90***	0.36	−6.43***	0.28	0.94***
Classico	2.31***	1.62***	1.24***	−5.50***	0.12
Healthy Choice	2.32***	0.90***	0.81***	0.73***	−5.84***

[a] Elasticities are read from left to right. The uncompensated price elasticities are given by $\epsilon_{kl} = -\Delta_{kl} + (1/s_k)(\gamma_{kl} - \beta_k s_l)$. The Kronecker delta Δ_{kl} equals one for $k = l$ and zero otherwise. Average shares are used in the calculation.
Note: *** 1% significance level; ** 5% significance; * 10% significance level.

The four respective pairs of prices are Prego and Ragu, Classico and Ragu, Classico and Prego, and Healthy Choice and Hunt's. Those restrictions that we failed to reject were imposed on the model. This produced savings in degrees of freedom. Throughout the remainder of this section the results are reported with the respective restrictions imposed.

Price Elasticities

For the class of markets under investigation, calculations used to obtain measures of market power and pricing conduct require demand elasticities for each brand. Table 2 summarizes the point estimates of the uncompensated own-price and cross-price elasticities for the five estimated demand equations. The price elasticities for the LA/AIDS model are calculated using average shares (s_k) and, for each demand equation, are read from left to right in the table. The uncompensated elasticity formula employs Chalfant's assumption (e.g., $\partial \log P^*/\partial \log p_l = s_l$) and is given by $\epsilon_{kl} = -\Delta_{kl} + 1/s_k(\gamma_{kl} - \beta_k s_l)$. The Kronecker delta Δ_{kl} equals one for $k = l$ and zero otherwise. Alston, Foster, and Green found this elasticity measure to perform well relative to alternatives in Monte Carlo experiments.

The own-price elasticities are found along the diagonal in both sections of Table 2. They are statistically significant ($p < 0.01$) and negative and, not surprisingly, the demand for each brand is elastic. For example, in the case of the FE3SLS estimator, a 1% increase in the price of Ragu leads to a 3.84% decrease in the quantity sold of Ragu. There are several explanations for the elastic demands. As noted above, a jar of spaghetti sauce is a durable good because its storable life exceeds the time between price changes, which implies that each product is an intertemporal substitute for itself (Tirole). Hence, consumers can stockpile sauce when it is on sale and avoid purchases at the regular price. The end result is a heightened state of consumer price sensitivity. This finding is entirely consistent with marketing research literature examining the adverse effects of frequent price discounting (Papatla and Krishnamurthi).

Another explanation for the magnitude of own-price elasticities stems from the disaggregated data used in this study. Many empirical demand studies for food use yearly, quarterly, or monthly data for broad groups of related products. Because of the long-run nature of the data and the masking effect induced by brand and market aggregation, own-price elasticities obtained in these studies are usually found to be inelastic. In the present study,

Table 3. Point Estiamtes of Price Reaction Elasticities[a]

	Ragu	Prego	Hunt's	Classico	Healthy Choice
FE3SLS Model					
Ragu	—	0.47***	0.08***	0.30***	0.14***
Prego	0.45***	—	0.05**	0.25***	0.03
Hunt's	0.33***	0.16***	—	0.40***	0.10**
Classico	0.32***	0.26***	0.11***	—	0.22***
Healthy Choice	0.27***	0.17***	0.06*	0.39***	—
EC3SLS Model					
Ragu	—	0.48***	−0.10*	0.23***	0.33***
Prego	0.32***	—	0.08*	0.26***	0.08*
Hunt's	0.14***	0.05	—	0.04	0.21***
Classico	0.58***	0.37***	0.25***	—	0.09
Healthy Choice	0.47***	0.09***	0.30***	0.23***	—

[a] Elasticities are read from left to right.
Note: *** 1% significance level; ** 5% significance; * 10% significance level.

weekly brand-level scanner data was used to measure the consumer response to price changes in a narrowly defined market. Results of other studies based on this micro-level data corroborate this finding (Guadagni and Little; Capps and Nayga; Cotterill; Seo and Capps). For example, despite being biased and inconsistent, the average uncompensated own-price elasticities of demand in the Seo and Capps study are similar to those in Table 2 for Prego, Ragu, Classico, and Hunt's sauces for the 10 IRI markets used in this study.[7]

The cross-price elasticities are found off the diagonal in both sections of Table 2. For the FE3SLS estimator, all cross-price elasticities are statistically significant ($p < 0.01$) and, consistent with a-prior expectations, positive. Thus, the brands represent economic substitutes and constitute a well-defined, relevant product market. For example, a 1% increase in the price of Prego leads to a 1.35% increase in the quantity sold of Ragu. In the case of the EC3SLS estimator, all but five of 20 cross-price elasticities are statistically significant ($p < 0.01$) and positive. There are two negative cross-price elasticities, albeit at marginal levels of statistical significance, and three positive but statistically insignificant cross-price elasticities. The common store location of this market facilitates price comparisons and logically leads to the general finding of economic substitutes. For similar studies reporting cross-price elasticities, the products were generally found to be economic substitutes (Cotterill; Seo and Capps).

[7] The Seo and Capps study actually used *store-level* data for a sample of 1,500 spatially dispersed stores. The authors then aggregated the data by store into 40 "markets." These data should not be confused with the standard IRI Infoscan *market-level* data used in our study. This data includes not only the 1,500 stores, but also the other stores scanned in each market (e.g., it includes the entire population of scanned stores in each market). Hence, the IRI *market-level* data is more comprehensive and appropriate for the empirical objective of this paper (Cotterill).

Price Reaction Elasticities

The calculations used to obtain measures of market power and pricing conduct also require information regarding a firm's price reaction, or response, to rivals' pricing decisions. Table 3 summarizes this information regarding the strategic interaction among the firms. Read from left to right, the table presents the point estimates of the price reaction elasticities for each of the five price reaction equations in the econometric system. Given the double log specification, the price-reaction elasticities, or ϵ_{kl}^{pr}, are given by the estimated ϕ_{kl} parameters from equation (6).

The empirical results are consistent with

Table 4. Point Estimates of Measures of Market Power and Pricing Conduct

Brands	Non-Following Elasticity[a] (i)	Observed Elasticity[b] (ii)	Fully Collusive Elasticity[c] (iii)	Rothschild Index (iii)/(i)	O Index (iii)/(ii)	Chamberlin Quotient 1-(ii)/(i)
FE3SLS Model						
Ragu	−3.84	−2.77	−1.00	0.26	0.36	0.28
Prego	−3.41	−2.51	−0.94	0.28	0.37	0.27
Hunt's	−5.59	−5.19	−1.01	0.18	0.19	0.07
Classico	−5.74	−4.28	−1.07	0.19	0.25	0.25
Healthy Choice	−7.26	−6.41	−1.14	0.16	0.18	0.12
EC3SLS Model						
Ragu	−3.92	−2.57	−1.10	0.28	0.43	0.35
Prego	−3.16	−2.24	−0.80	0.25	0.36	0.29
Hunt's	−6.43	−6.14	−3.96	0.62	0.64	0.05
Classico	−5.50	−4.46	−0.20	0.04	0.05	0.19
Healthy Choice	−5.84	−4.76	−1.08	0.19	0.23	0.19

[a] Uncompensated own price elasticities, or ϵ_{kl} for $k = l$, from Table 2.
[b] Observed elasticity is given by $\epsilon_k^o = \epsilon_{kk} + \Sigma_{l \neq k} \epsilon_{kl} \epsilon_{lk}^{pr}$, where ϵ_{lk}^{pr} are the price reaction elasticities from Table 3.
[c] Fully collusive elasticity is given by $\epsilon_k^{fc} = \Sigma_l \epsilon_{kl}$.

theory, as the estimated price reaction functions are generally upward-sloping (Shapiro). For example, under the FE3SLS estimator, the Ragu price reaction function (in the row labeled Ragu) shows that a 1% increase in the price of Prego results in a 0.47% increase in the price of Ragu. Nineteen of the 20 price reaction elasticities for the FE3SLS estimator are statistically significant ($p < 0.10$) and, consistent with a-priori expectations, positive. Only one price reaction elasticity is positive and statistically insignificant. Sixteen of the 20 price reaction elasticities for the EC3SLS estimator are statistically significant ($p < 0.10$) and positive. Three price reaction elasticities are positive and statistically insignificant. In the Ragu price reaction function, the elasticity for the price of Hunt's sauce is negative and marginally significant ($p < 0.10$).

Cotterill found similar results in the direction, magnitude, and statistical significance for leading brands in the carbonated soft drink industry (e.g., the Pepsi price reaction elasticity in the Coke empirical price reaction function was 0.51, while the Coke price reaction elasticity in the Pepsi empirical price reaction function was 0.56). Both estimates were statistically significant ($p < 0.01$). Consistent with Liang, price coordination appears to be more common within a market segment than between segments. For example, in the EC3SLS model, the respective pairs of price reaction elasticities are greater in magnitude for Ragu and Prego (within the traditional segment) than for either Hunt's and Prego (between the canned and traditional segments) or Healthy Choice and Classico (between the health-conscious and premium segments).

Measures of Market Power and Pricing Conduct

Combining the results from Tables 2 and 3 yields the various measures of market power and pricing conduct that address the empirical objective of this paper. Table 4 summarizes these point estimates. The market power formulas are restated in the footnotes to the table for convenience. The second column in Table 4 contains the non-followship elasticities (US DOJ and FTC Horizontal Merger Guidelines). These are simply the uncompensated own-price elasticities from Table 2. As the name implies, this elasticity measures the sensitivity in quantity sold that a firm faces when it raises price but no rivals follow. For this reason, the non-followship elasticity represents a unilateral measure of market power. The findings are very similar across both estimators. Ragu and Prego, the two largest brands in the market, maintain the greatest degree of unilateral market power relative to the other competing brands. However, in an absolute sense, none of the brands in this market maintains much unilateral market power as each has a sensitive non-followship elasticity. The average non-followship elasticity in the Cotterill study of -1.53 is much less elastic than those obtained for the brands in this study.

The third column in Table 4 contains the observed elasticities. These are derived following Baker and Bresnahan and are given by $\epsilon_k^o = \epsilon_{kk} + \Sigma_{l \neq k} \epsilon_{kl} \epsilon_{lk}^{pr}$. Following Cotterill, the elasticities of demand (e.g., ϵ_{kk} and ϵ_{kl}) are taken from Table 2 and the empirical price reaction elasticities (*e.g.,* ϵ_{lk}^{pr})[8] are taken from Table 3. In the presence of imperfect tacit price collusion (*e.g.,* $0 < \epsilon_{lk}^{pr} < 1$), the observed elasticities will be smaller in absolute value than the non-followship elasticities. In some cases, the observed elasticities are a full percentage point lower in absolute value than the non-followship elasticities. This is further evidence of tacit price collusion in the spaghetti sauce market and the pattern persists across both estimators. Based on the observed elasticities, Ragu and Prego again, in a relative sense, maintain the greatest degree of market power. As was the case for the non-followship elasticities, in an absolute sense, none of the brands maintain much market power based on the observed elasticities. The average observed elasticity in the Cotterill study of -1.45 is much less elastic than those obtained for the brands in this study.

The fourth column in Table 4 contains the fully collusive elasticities. These are given by $\epsilon_k^{fc} = \Sigma_l \epsilon_{kl}$. In the presence of imperfect tacit price collusion, the fully collusive elasticities will be smaller in absolute value than the observed elasticities.[9] With the exception of Hunt's and Classico in the EC3SLS model, the fully collusive elasticities are within one-fifth of a percentage point of being unitary elastic.[10] Thus, the brands in the market acting in concert have the potential to substantially reduce the elasticity of their own respective demand curves. In this scenario, each firm in the market exercises a high degree of absolute market power. The average fully collusive elasticity in the Cotterill study of -0.94 is fairly close to those obtained for the brands in this study.

The fifth and sixth columns in Table 4 contain two indexes of market power. The Rothschild index (*RI*) measures the existence of market power as the ratio of the fully collusive elasticity to the non-followship elasticity (Greer). Under perfect competition, the latter elasticity converges to negative infinity, driving the ratio to zero. Under monopoly, the two elasticities coincide, resulting in an index value of one. The closer the index is to one the greater the degree of market power, and it is bounded as $0 \leq RI \leq 1$.

The O index (*OI*), introduced by Cotterill, also measures the existence of market power. The *OI* is given by the ratio of the fully collusive elasticity. Similar to the *RI*, it is bounded as $0 \leq OI \leq 1$. Again, the closer the index is to one the greater the degree of market power. The relationship between the two indexes is $0 \leq RI \leq OI \leq 1$. In a relative sense, Ragu and Prego tend to exercise more market power than the other brands. One exception to this is Hunt's in the EC3SLS model. In an absolute sense, comparing each index value to its maximum possible value of unity shows that no firm in the market exercises much market power. Additionally, the average *RI* and *OI* values in the Cotterill study of 0.67 and 0.72, respectively, exceed those obtained for the brands in this study. Again, this indicates less

[8] The subscripts on the price reaction elasticity are intentionally reversed to indicate the appropriate *column* of data from Table 3.

[9] The relationships among the three elasticities in Table 4 are as follows:

$$\epsilon_k^o = \epsilon_{kk} \quad \text{for } \epsilon_{lk}^{pr} = 0 \quad \text{and}$$

$$\epsilon_k^o = \epsilon_{kk} \quad \text{for } \epsilon_{lk}^{pr} = 0.$$

[10] If preferences were homothetic (e.g., $\beta_k = 0\ \forall k$), the fully collusive elasticities would be exactly equal to unity with homogeneity imposed on the model.

market power exists in the spaghetti sauce market than in the carbonated soft drink market.

The last column in Table 4 contains the values of the Chamberlin quotient (*CQ*). The *CQ*, introduced by Cotterill, is given by one less the ratio of the observed elasticity to the non-followship elasticity. Thus, it quantifies the fraction of market power, as measured by the observed elasticity, derived from tacit price collusion. It is bounded as $0 \leq CQ \leq 1$, where higher *CQ* values indicate higher levels of tacit price collusion. Under both estimators the pattern of results is similar. Ragu and Prego each derive roughly 30% of their market power from tacit price collusion. Classico and Healthy Choice derive less market power from tacit price collusion than the two leading brands, while Hunt's derives the least. Juxtaposing this result with those based on the *RI* and *OI*, Hunt's appears to possess market power due to its positioning in a niche segment, not collusion. This result is consistent with the brewing industry (Baker and Bresnahan), where the three largest players were found to exercise market power in the absence of price collusion. Finally, the average *CQ* value for the two leading brands in the Cotterill study (e.g., Coke and Pepsi) of 0.15 is less than most of those obtained for the brands in this study, indicating more tacit price collusion exists in the spaghetti sauce market than in the regular carbonated soft drink market.

Transportation Costs

In the FE3SLS model, the λ_k parameters are not estimable as the transportation costs, or log c_{ki} variables, do not vary by week and, hence, are correlated with the fixed-effects for the markets. Given the generality of the EC3SLS framework (Cornwell, Schmidt, and Wyhowski), it was possible to include time-invariant effects in the design matrix of the model. Consequently, transportation costs were built into each price reaction function to empirically separate their effect on a firm's price from that of rivalrous pricing conduct. Table 5 summarizes the point estimates of the transportation cost elasticities for the five brands represented in the econometric system. The first version of the EC3SLS model assumes transportation costs are defined to be the product of shipping distance and fuel price, while the second version of the EC3SLS model uses shipping distance only.

Table 5. Point Estimates of Transportation Cost Elasticities[a]

	FE3SLS Model	EC3SLS Version 1	EC3SLS Version 2
Ragu	—	0.02***	0.02***
Prego	—	−0.11***	−0.13***
Hunt's	—	−0.01***	−0.01***
Classico	—	0.003	0.004*
Healthy Choice	—	−0.01**	−0.01**

[a] Transportation costs are defined to be the product of distance and fuel price in Version 1 and distance only in Version 2.

Note: *** 1% significance level; ** 5% significance; * 10% significance level.

The results for both versions of the EC3SLS model are similar. The Ragu transportation cost elasticity across both models is statistically significant ($p < 0.01$) and positive. Thus, a 1% increase in the transportation cost leads to a 0.02% increase in the price of Ragu. The elasticity for Classico is also positive for both models, but is statistically significant ($p < 0.10$) in the second version only. The positive coefficient is evidence that the firms may be practicing basing-point pricing (Greenhut, Norman, and Hung). Across both models, the transportation cost elasticities for Prego, Hunt's, and Healthy Choice are statistically significant ($p < 0.05$) and negative. However, the elasticities are very small in magnitude in the case of Hunt's and Healthy Choice. Based on these parameter estimates and the relationship among the firms in the industry, it may be conjectured that Prego is more concerned with being obedient to avoid a punishment strategy administered by Ragu than to price its products according to transportation costs.

Summary

Point estimates of measures of market power and pricing conduct were obtained at the brand level in a representative product-differentiated, oligopolistic output market using a simultaneous-equations panel data model. The $1.3 billion domestic spaghetti sauce industry was chosen as the case study for this paper. The empirical model builds on Baker and Bresnahan and Liang, and extends the approach taken by Cotterill. The empirical findings augment the existing pool of information available to business strategists and antitrust policy makers for this class of markets. Evidence of market power was found in the domestic spaghetti sauce industry, albeit to a lesser extent than in the carbonated soft drink industry. However, a higher percentage of market power was derived from tacit price collusion in the former than in the latter. Like the ready-to-eat breakfast cereal industry, a higher degree of tacit price collusion was found among brands within a specific market segment than between market segments. Similar to the three largest firms in the brewing industry, one firm in this study was able to maintain market power in the absence of tacit price collusion, likely due to its positioning in a niche segment of the market.

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