Food Marketing Policy Center

Market Power and/or Efficiency:
An Application to U.S. Food Processing

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Preface

This article separates oligopoly-power and cost-efficiency effects of changes in industrial concentration and assesses their impact on output prices in 32 food-processing industries. Empirical results indicate that although concentration induces cost efficiency in one-third of the industries, oligopoly-power effects either dominate cost efficiency or reinforce inefficiency, resulting in higher output prices in most industries. The article also provides fresh econometric estimates of oligopoly power and economies of size for the industries in question.

Key words: industrial concentration, economies of scale, industrial organization, oligopoly power, food processing

JEL Codes: L00, L11, L13, L66

1. Introduction

In his review of the new empirical industrial organization (NEIO) literature, Bresnahan (1989) concludes that although NEIO models are useful in measuring market power, they are not as useful in guiding policy when market structure rather than conduct is the policy target. First, he argues that the narrow focus of NEIO studies on single and often heavily concentrated industries is too limited to be useful in mapping structure into conduct and performance. Cross-industry studies, on the other hand, provide information over a wider range of industries and, despite their well-known problems, continue to influence policy (Salinger, 1990). Second, Bresnahan also argues that since market power is imputed rather than observed in NEIO models, its relationship to observable structural and behavioral variables in often unclear to policy makers. Thus, to make NEIO findings more policy-relevant, studies should consider a wider range of industries and incorporate observable structural measures of interest to policy makers, such as industrial concentration.

This article develops an NEIO model that formally incorporates measures of industrial concentration and separates out the oligopoly-power from the cost-efficiency effects of concentration on output prices.[1] The model is the oligopoly analogue of Azzam's (1997) oligopsony model, which extends Appelbaum's (1982) model to formally include industrial concentration. The separation of the two effects is not only of academic interest but is also of public policy concern because the basic problem facing antitrust authorities is that of a tradeoff between efficiency and market power (Williamson, 1968; Perry, 1984; Bian and McFetridge, 2000). That is, whether or not concentration is in the public interest depends critically on whether or not the cost-efficiency gains through concentration offset the welfare losses from greater market power.

The model is applied to 4-digit SIC data on 32 U.S. manufacturing industries over the 1972-92 period. The econometric results provide fresh estimates of oligopoly power and economies of size in these industries and reveal that Cournot behavior is widespread. The empirical findings also indicate that although cost-efficiency effects from concentration are important in one-third of the industries, in nearly every case the oligopoly-power effects dominate or reinforce cost inefficiencies, resulting in higher output prices. The few exceptions where concentration is beneficial to buyers are in the fat-and-oil sector, which is characterized by strong economies of size and product homogeneity. Finally, this analysis shows that NEIO models that bridge the gap between conduct and market structure can be useful for policy making decisions, especially those targeting industries based on efficiency vs. the consumer's interest.

2. The Model

The starting point is an industry of N firms producing a homogeneous good Q requiring factors x_r for $r = 1, \ldots, k$ and facing a derived market demand curve

$$Q = f(p, z), \tag{1}$$

where p is output price and z is a vector of demand shifters. Profit maximization by the j[th] firm yields the supply relation

$$p = -\frac{s_j}{h}(1 + f_j) + \frac{\partial C_j(q_j, \underline{w}, t)}{\partial q_j}, \tag{2}$$

where $s_j = q_j / Q$ is the j[th] firm's market share, $h = (dQ/dP)(1/Q)$ is the semi-elasticity of demand $(h < 0)$,

$$f_j = d\sum\nolimits_{i \neq j}^{n} q_i / dq_j$$

is the j[th] firm's conjectural variation, $C_j(.)$ is the cost function, $\underline{w}$ is a vector of factor prices, and t is the state of technology. By Shephard's Lemma, the derived demand for the r^{th} factor by the j[th] firm is given by

$$x_{rj} = \frac{\partial C_j(q_j, \underline{w}, t)}{\partial w_j} \quad \text{for } r = 1, 2, \ldots, k. \tag{3}$$

[1] A large number of studies have tested the relationship between efficiency proxies and price-cost margins (e.g., Demsetz, 1973; Martin, 1988; Rosenbaum, 1994) while others have tested more explicitly the relationship between costs and concentration (Peltzman, 1977; Dickson, 1994) or the ad hoc relationship between price and market structure (e.g., Cotterill, 1986). A few recent studies have separated oligopsony power from cost efficiency using structural models (Azzam, 1997; Azzam and Schroeter, 1995) but studies of this type focusing on oligopoly power are lacking.

Following Olson and Shieh (1989), and Baffes and Vasavada (1989), the cost function is assumed to take the modified generalized Leontief form

$$C_j(q_j, \underline{w}) = q_j \sum_i \sum_j \alpha_{ij} w_i^{1/2} w_j^{1/2} + q_j t \sum_i \gamma_i w_i + q_j^2 \sum_i \beta_i w_i \quad , \qquad (4)$$

where α_{ij}, γ_I, and β_I are parameters. Multiplying through equations (2) and (3) by s_j, using (4), and summing across the industry yields, respectively, the industry-wide analogue of the supply relation

$$p = -\frac{H(1+\Phi)}{\eta} + \sum_i \sum_j \alpha_{ij} w_i^{1/2} w_j^{1/2} + t \sum_i \gamma_i w_i + 2HQ \sum_i \beta_i w_i \qquad (5)$$

and factor demand

$$\frac{X_r}{Q} = \sum_i \sum_j \alpha_{ij} \left(\frac{w_j}{w_i} \right)^{1/2} + t\gamma_i + HQ\beta_i \qquad (6)$$

for r=1,2,…k,

where

$$H = \sum_j s_j^2$$

is the Herfindahl index, Φ is the industry (weighted) conjectural variation and

$$X_r = \sum_j x_{rj}$$

is total industry employment of the r^{th} factor. By virtue of the expression for the semi-elasticity of demand, the demand function (1) takes the semi-logarithmic form

$$\ln Q = \delta_0 + \eta p + \sum_i \delta_i z_i , \qquad (7)$$

where η, δ_0 and δ_i are parameters. The first term on the right-hand side of the supply relation in (5) is the mark-up over marginal cost. Its magnitude depends on the level of concentration (H), the semi-elasticity of demand (η), and the type of market conduct (Φ). If conduct is competitive, then $\Phi = -1$ and the markup is zero. For Cournot, $\Phi = 0$ and the markup is $-H/\eta$. For conduct that is less competitive than Cournot, $0 < \Phi \leq (1/H) - 1$ and the upper bound on the markup is $-1/\eta$. However, for noncompetitive conduct, concentration affects both the markup and marginal cost. Note that the commonly-used conjectural variation elasticity (Appelbaum, 1982) can be defined as $\Phi^* = (1+\Phi)H$ which ranges between 0 and 1, the price elasticity of demand is given by $\eta^* = \eta P$, and the industry oligopoly power is defined by $L = -\Phi^* / \eta^*$.

Following Azzam (1997), Φ is treated as constant.[2] Thus, differentiating (5) with respect to H yields the decomposed effects of concentration on output price:

$$\frac{dp}{dH} = -\frac{(1+\Phi)}{\eta} + 2Q \sum_i \beta_i w_i , \qquad (8)$$

where the first term on the right-hand side is the oligopoly-power effect and the second is the cost-efficiency effect (or the effect of a rise in concentration on marginal cost).

A measure for the cost elasticity with respect to output is given by the ratio of industry marginal cost to average cost

$$e_{cy} = \frac{\mathrm{A} + 2HQB}{\mathrm{A} + HQB} , \qquad (9)$$

where

$$\mathrm{A} = \sum_i \sum_j \alpha_{ij} w_i^{1/2} w_j^{1/2} + t \sum_i \gamma_i w_i$$

and

$$B = \sum_i \beta_i w_i .$$

Note that e_{cy} depicts economies of size and is the inverse of the degree of returns to scale. If B=0, constant returns exist, and the only effect of rising concentration on price is through oligopoly power. If B>0, diseconomies of scale exist, and a rise in concentration raises prices through a rise in both oligopoly power and costs. When

[2] Following the work of Stigler (1964) and Stalhammar (1991), two additional specifications of Φ were tested. One was to let $\Phi = \theta_0 + \theta_1 \ln H$ in order to allow industry conduct to vary with concentration. Another test was the inclusion of imports and exports. Although these extensions provided some interesting insights in some cases, they deteriorated the results of interest. Given the focus and scope of this article, Φ was therefore treated as a constant.

economies of scale are present (B<0), the effect of a rise in concentration on price can be positive, negative, or zero, depending on whether the oligopoly-power effect is larger than, smaller than, or the same as the cost-efficiency effect.[3]

3. Empirical Procedures

The model is operationalized with data at the 4-digit SIC (1987 definitions) level for the 1972-92 period. The econometric model is based on equations (5), (6), and (7) depicting pricing behavior, input demand equations, and the output demand equation. Although (5) is the main equation of interest, the input and output demand equations impose stronger theoretical restrictions and assist in identifying the corresponding parameters in the pricing equations. The model assumes three variable inputs: materials (X_K), labor (X_L), and capital (X_K).[4] Thus, the estimating model consists of five equations: the pricing equation, three input demand equations, and the output demand equation. The latter is assumed to be a function of output price (P), income (y) and a trend variable (t), where price and income are deflated by the consumer price index (d).

The endogenous variables are Q, P, X_K, X_L, and X_K. The exogenous variables are W_K, W_L, W_K , y, t (=1 for 1972 through 21 for 1992), d, and H. The parameters to be estimated are $\Phi, h,$ the a_{ij}'s, the b_j's, the g_i's, d_0, and the δ_0 and δ_i's.[5]

[3] Note that these effects of concentration are for a constant level of output and that higher concentration leads to lower, higher, or no change in costs if there are increasing, decreasing, or constant returns to scale, respectively. By fixing output, a rise in the Herfindahl index implies a change in the distribution in output across firms, with more output being produced by the larger firms leading to lower industry cost in the presence of economies of scale. Note that the econometric model measures economies of scale on a market share-weighted industry cost function and that technical change (t) is assumed to affect the industry marginal cost intercept, not the slope.

[4] See Bartelsman and Gray (1996) for precise definitions of these variables. Following the U.S. Department of Labor (1997), capital services are assumed to be proportional to capital stocks. The derivation of the price of capital is discussed below. The material inputs include raw materials (agricultural, packaging, and certain services) as well as energy. The latter accounted for less than 2% of variable costs in all cases.

[5] The structural model contained 17 coefficients, which we estimated with 21 observations. Since the sample is small and the standard errors for nonlinear models are only approximately correct for small samples, the statistical significance of the coefficients should be interpreted with caution.

The main data source for prices and quantities of outputs and inputs was the online National Bureau of Economic Research database of Bartelsman, Becker and Gray (2000) on U.S. manufacturing industries.

Following the U.S. Department of Labor (1997, p. 107), we define the rental price of capital in dollars per unit of real capital stock. Due to lack of data on the price of capital at the 4-digit SIC level, all industries are assumed to face the same rental prices but each to have different levels of capital stock. Therefore, the rental price of capital was computed by dividing the cost of capital services (provided electronically by the Bureau of Labor Statistics) divided by total capital assets at the 2-digit SIC level. Due to data limitations, we use an instrumental variable for H for the years when it was not available (see Appendix for details).

Given the endogeneity of output quantity and the price of output, the system of five equations is estimated with non-linear 3SLS using the SHAZAM 7.0 software. The results for 32 food industries are presented below.

4. Empirical Results

4.1 Conduct, Demand, and Economies of Size

Table 1 presents selected estimated parameters: Φ, h, L and e_{cy}. The null hypothesis for conduct is $\Phi = -1$ (competitive behavior) and for the elasticity of cost with respect to output $e_{cy} = 1$ (constant returns to scale). An alternative conduct hypothesis tested is Cournot behavior $(\Phi = 0)$, given its common use in empirical analysis. As usual, one, two and three asterisks next to the estimated coefficients indicate significance at the one, five and 10 percent levels, respectively. Note that the indicated statistical significance for Φ an e_{cy} in Table 1 are relative to –1 and 1 (rather than the usual null hypothesis of 0), respectively.

The results in Table 1 indicate that the estimated Φs ranged from –0.996 for SIC 2062 (very close to the lower limit of –1) to 6.515 for SIC 2097 (well below its maximum limit of 63 given by $H^{-1} - 1$). Twenty of 32 industries (63%) had non-competitive industry conduct at the 5 percent level. It should be noted that the estimated Φs for 20 of the industries also rejected the null hypothesis of Cournot behavior $(\Phi = 0)$ at the 5% level. The derived conjectural variation elasticities ranged from 0.00068 for SIC 2062 to 0.2445 for 2075. The estimated oligopoly Lerner indexes (evaluated at mean values of *P*) ranged from 0.035 in SIC 2075 to 0.815 in SIC 2035. Twenty-seven of the 32 industries (84%) have mark-ups statistically different from perfect competition at the 5% level.

The number of industries with non-competitive mark-ups is much larger than those indicating non-collusive conduct from only estimated Φs. As shown by Nevo (2000) in the breakfast cereal industry, it is possible to have non-collusive behavior and yet have a strong degree of oligopoly power and level of mark-up. As a case in point, the soybean oil industry (SIC 2075) was the most collusive, based on the conjectural variation elasticities; however, this industry had the lowest mark-up as indicated by the Lerner index, suggesting that in some cases concentration is a way to survive when industries are operating on small margins. Nonetheless, the results are consistent with a myriad of studies that have found that most food processing industries are non-competitive.[6]

In terms of economies of size (calculated at mean values of data), the results reveal that approximately 12 industries (38%) have significant economies of size, 9 (28%) have significant diseconomies of size, and the remaining 11 (34%) did not reject the constant returns to size hypothesis at the 5% level. The economies of size parameters, which are critical for the evaluation of cost efficiency effects, appear to be reasonable and consistent with previous findings.[7]

4.2 Estimated Market Power and Efficiency Effects

Table 2 presents the results based on equation (8) for the separate effects of changes in the Herfindahl index on oligopoly-power, cost efficiency, and output price. These effects were calculated and tested at mean values of the data.

In terms of oligopoly power, at the 10% level, the results indicate that 26 of the 32 industries (81%) significantly increase oligopoly-power as concentration increases. For the remaining six industries, increases in concentration do not result in significant increases in oligopoly power.[8]

In terms of cost effects, the results indicate that 11 industries (34%) show significant gains (at the 10% level) in cost efficiency with concentration, while eight show a positive and significant increase in cost as concentration rises. The remaining 13 industries do not show a significant impact of concentration on cost efficiency.[9] The question then remains how much of these potential cost savings is being passed on and how much is being pocketed by the industries.

Combining oligopoly power and cost effects, concentration results in significant decreases in output prices in only three industries (9.4%), with 22 industries (68.7%) showing a net and significant increase in price as concentration rises, while seven (21.9%) show insignificant output price effects at the 10% level. The notable cases where concentration has a benign effect on prices are in the fat and oil industries: 2074, 2076 and 2077.

A number of industries show a strong trade-off between oligopoly-power and efficiency effects, where both effects are statistically significant and their magnitudes considerable. In this category fall SICs 2011 (meatpacking), 2022 (cheese), 2026 (fluid milk), 2044 (rice milling), 2074 (soybean oil), 2076 (vegetable oil), 2077 (animal fats and oil), 2097 (manufactured ice), 2098 (macaroni and spaghetti), and food preparations (2099).

5. Concluding Remarks

The purpose of this article is two-fold. First, it provides a theory-based empirical model that separates the oligopoly-power and cost-efficiency effects of industrial concentration on output prices. The econometric model adapts the oligopsony model of Azzam (1997) to the oligopoly-power case. A second contribution is that by applying the model to time-series data across U.S. food processing industries, information

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[6] See, for example, Bhuyan and Lopez (1997). In particular, the estimated Lerner index for the meatpacking industry (average=0.099) is between the ones estimated by Schroeter (1988) for beef (0.04) and by Schroeter and Azzam (1990) for beef and pork (0.14). Likewise, the mark-ups estimated by Morrison-Paul (2001) correspond to Lerner indexes between 0.05 and 0.20 between 1971 and 1991.

[7] The average estimate of Morrison-Paul (2001) for long run economies of size in meatpacking is 0.965 (cf. 0.95 in Table 1). The average degree of economies of size in Table 1 is 0.997, which is nearly CRS, somewhat above the 0.834 estimated by Bhuyan and Lopez (1997).

[8] The impacts are particularly large and significant for sausages and prepared meats (2013), poultry and egg processing (2015), fluid milk (2026), candy and confectionary (2064), prepared feeds (2048), animal and marine fats and oils (2077), canned fruits and vegetables (2033), manufactured ice (2097), macaroni and spaghetti (2098), and food preparations (2099).

[9] In descending order, the cost efficiency effects that are statistically significant at the 10% level are particularly pronounced in the following industries: animal and marine fats (2077), fluid milk (2026), cottonseed oil (2074), manufactured ice (2097), food preparations (2099), rice milling (2044), vegetable oil (2076), and meat packing (2011).

is provided on the potential impact of further concentration on oligopoly power and cost efficiency within each industry and across industries, which can be useful to policy makers concerned with market structure.

From the empirical results, we see some systematic patterns of the impact of industrial concentration in the U.S. food industries. Specifically, we find that further increases in concentration would: (1) significantly increase oligopoly power; (2) result in cost efficiency in one-third of the industries; and (3) increase output price in nearly every case. These patterns raise some interesting questions. What is and what determines, for example, the critical level of concentration beyond which concentration induces net inefficiency?

There are other issues that we do not address here but that are worthy of further attention. Among these are the role of foreign trade and possible simultaneous oligopsony effects on factor prices in certain markets. In spite of the pattern of concentration, there are strong differences among industries. It is timely and relevant to examine the sources and consequences of these differences more fully. Extending the analysis to incorporate any of these issues or applying it to other industries will provide further insights into the market power and cost-efficiency effects of changes in concentration. Finally, this article shows that NEIO analyses can be made more policy-relevant if they cover market structure and a wider range of industries, as suggested by Bresnahan (1989).

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Table 1: Selected Parameter Estimates, U.S. Food Processing Industries, 1972-92.

SIC	Industry	Φ	s.e.	h	s.e.	L	s.e.	$\bar{e}_{cy}$	s.e.
2011	Meat Packing	-0.603***	0.147	-0.175***	0.059	0.099***	0.011	0.950***	0.013
2013	Saus. & Prep. Meats	0.726	0.808	-0.243**	0.112	0.144***	0.017	0.998	0.015
2015	Poultry & Egg Proc.	-0.482	0.329	-0.135	0.083	0.106***	0.017	0.489	0.011
2022	Cheese	-0.865	0.129	-0.095	0.090	0.094***	0.015	1.011	0.011
2023	Dry/Cond.& Ev.Milk	0.099	0.412	-0.387***	0.173	0.197***	0.059	1.037	0.039
2024	Ice Cream & Fr. Desserts	-0.872	0.634	-0.049	0.241	0.097***	0.023	1.063**	0.026
2026	Fluid Milk	3.154**	1.528	-0.444***	0.157	0.200***	0.017	0.916	0.019
2032	Canned Specialties	-0.785***	0.073	-0.404***	0.101	0.125***	0.026	1.214***	0.032
2033	Canned Fruit & Veg.	-0.397	0.509	-0.168	0.139	0.118**	0.020	1.104***	0.023
2034	Dried Fruit & Veg.	-0.871***	0.196	-0.114	0.171	0.086***	0.023	1.195***	0.029
2035	Pickles, Sauces, etc.	0.063	0.687	-0.433	0.270	0.815***	0.046	0.884***	0.022
2043	Cereal Breakfast Foods	-0.863***	0.037	-0.259***	0.052	0.182**	0.033	1.553***	0.060
2044	Rice Milling	-0.495***	0.145	-0.191***	0.031	0.252	0.056	0.256**	0.062
2045	Prep. Flour & Doughs	-0.829***	0.153	-0.314*	0.188	0.063	0.041	1.310***	0.062
2046	Wet Corn Milling	-0.918***	0.039	-0.196***	0.045	0.090**	0.037	1.055	0.036
2048	Prep. Feeds	-0.422*	0.227	-0.096***	0.035	0.087***	0.018	0.991	0.019
2061	Cane Sugar	-0.752***	0.121	-0.121***	0.028	0.178**	0.071	0.918	0.070
2062	Cane Sugar Refining	-0.996***	0.011	-0.026	0.040	0.041	0.056	1.013	0.063
2063	Beet Sugar	-0.751***	0.098	-0.191***	0.040	0.215***	0.064	0.879*	0.074
2064	Candy & Conf. Prods.	0.512	0.342	-0.465***	0.077	0.272***	0.043	1.071	0.052
2066	Chocolate & Cocoa	-0.989***	0.018	-0.052	0.012	0.058	0.036	1.294***	0.046
2074	Cottonseed Oil Mill	-0.255	0.287	-0.332***	0.121	0.201	0.039	0.839**	0.027
2075	Soybean Oil Mill	0.993***	0.013	-0.023	0.034	0.035	0.024	0.997	0.022
2076	Vegetable Oil Mill	-0.636***	0.125	-0.221***	0.063	0.198***	0.042	0.841***	0.040
2077	An./Mar. Fats & Oils	-0.002	0.315	-0.095***	0.029	0.417***	0.049	0.732***	0.042
2082	Malt Beverages	-0.682***	0.116	-0.864***	0.186	0.090***	0.027	1.103***	0.029
2084	Wines & Brandy	-0.016	0.727	-0.433	0.317	0.204***	0.088	0.971	0.025
2087	Extracts & Syrups	-0.557***	0.141	-0.349***	0.105	0.286***	0.030	1.529***	0.058
2095	Roasted Coffee	-0.853***	0.098	-0.165***	0.025	0.147	0.095	1.077	0.118
2097	Manuf. Ice	6.515***	2.255	-0.876***	0.218	0.241***	0.035	0.908***	0.032
2098	Macaroni & Spaghetti	0.272	0.367	-0.354***	0.103	0.503***	0.042	0.799***	0.040
2099	Food Preparations	2.453***	0.921	-0.363***	0.081	0.329***	0.043	0.918*	0.045

Notes: Levels of statistical significance are represented by *(10%), **(5%) and ***(1%). The standard errors (s.e.) are indicated next to the estimated coefficients. The null hypothesis for Φ is -1 (perfect competition), while the null hypothesis for e_{cy} is 1 (CRS). These results are based on the joint estimation of equations (5), (6) and (7).

Table 2: Impacts of Industrial Concentration on Oligopoly Power, Cost and Output Price.

		Impact of H on					
SIC	Industry	Oligopoly Power	s.e.	Cost Efficiency	s.e.	Output Price	s.e.
2011	Meat Packing	2.286***	0.333	-2.244***	0.564	0.042	0.257
2013	Saus. & Prep. Meats	7.119***	0.815	-0.175	1.294	1.944***	0.630
2015	Poultry & Egg Proc.	3.864***	0.599	-0.759	0.715	3.105***	0.354
2022	Cheese	1.426***	0.217	0.290	0.298	1.716***	0.167
2023	Dry Cond. & Evap. Milk	2.848***	0.556	0.804	0.863	3.652***	0.479
2024	Ice Cream & Fr. Desserts	2.637***	0.615	2.863**	1.150	5.500***	0.643
2026	Fluid Milk	9.373***	0.790	-6.867***	1.531	2.506***	0.835
2032	Canned Specialties	0.530***	0.111	1.317***	0.193	1.847***	0.103
2033	Canned Fruit & Veg.	3.613***	0.612	5.170***	1.099	8.783***	0.588
2034	Dried Fruit & Veg.	1.136***	0.301	3.958***	0.554	5.094***	0.299
2035	Pickles, Sauces, etc.	2.459***	0.353	-1.383***	0.239	1.076***	0.182
2043	Cereal Breakfast Foods	0.531***	0.095	1.621***	0.159	2.152**	0.084
2044	Rice Milling	7.658	0.587	-2.680**	1.201	-0.022	0.740
2045	Prep. Flour & Doughs	0.550	0.362	3.957***	0.773	4.507	
2046	Wet Corn Milling	0.423**	0.170	0.448	0.292	0.871***	0.164
2048	Prep. Feeds	6.043***	0.899	-0.809	1.706	5.234***	0.864
2052	Cookies & Crackers	0.946*	0.540	-0.872***	0.881	0.074***	0.638
2061	Cane Sugar	2.063**	0.817	-1.721	1.453	0.342	0.771
2062	Can Sugar Refining	0.184	0.254	0.107	0.521	0.291	0.273
2064	Beet Sugar	1.311***	0.419	-1.423*	0.834	-0.112	0.413
2064	Candy & Conf. Prods.	3.254***	0.512	1.135	0.826	4.389***	0.413
2066	Chocolate & Cocoa	0.229	0.140	1.706***	0.252	1.935***	0.125
2074	Cottonseed Oil Mill	2.246***	0.375	-3.456***	0.565	-1.210***	0.303
2075	Soybean Oil Mill	0.329	0.224	-0.062	-0.405	0.267	0.203
2076	Vegetable Oil Mill	1.653**	0.635	-2.548***	0.635	-0.815*	0.354
2077	An./Mar. Fats & Oils	10.551***	1.226	-12.099***	1.909	-1.548**	0.780
2082	Malt Beverages	0.369***	0.108	0.719***	0.164	1.083***	0.074
2084	Wines & Brandy	2.276***	0.239	-0.388	0.333	1.888***	0.189
2087	Fd. Extracts & Syrups	1.276***	0.134	2.193***	0.224	3.469***	0.106
2095	Roasted Coffee	0.901	0.583	0.788	1.200	1.689**	0.677
2097	Manuf. Ice	8.587***	1.240	-5.136***	1.731	3.451***	1.185
2098	Macaroni & Spaghetti	3.446***	0.283	-1.862***	0.370	1.584***	0.164
2099	Food Preparations	9.536***	1.290	-3.415*	1.843	6.121***	1.102

Notes: Levels of statistical significance are represented by * (10%), ** (5%) and *** (1%). The standard errors (s.e.) are indicated next to the estimated coefficients. These results are based on equation (8).

Appendix: Herfindahl Indexes

The Herfindahl-Hirschmann index has only been published for 1982, 1987 and 1992. Additionally, the partial industrial concentration ratios (CR4, CR8, CR20, and CR50) have been published for (then Census years including) 1972 and 1977. Given this paucity of data, *H* had to be estimated for each food industry for much of the sample period. This process involved two steps: (1) the estimation of the 1972 and 1977 *H* indexes from partial concentration ratios; (2) interpolation of the *H* indexes for the inter-Census years.

The first step involved the application of entropy to estimate the market shares of the top 50 firms. Using the technique presented by Golan, Judge, and Perloff (1996), the market shares were forecasted for all firms encompassed in CR4, CR8, CR20 and CR50. It turns out that the maximum entropy solution for the size distribution of firms yields equal market shares of firms within intervals of these concentration ratios. After forecasting the market share of each firm in 1972 and 1977, we restricted the estimated distribution of market shares to a third-degree polynomial function in order to estimate the most probable market share for each firm. The average R-square of the estimated polynomial functions was 95%. From the individual market share, we then estimated the *H* index as the sum of the squares of market shares resulting from the polynomial distribution.

The second step involved regressing the Census-year *H* indexes on a set of instrumental variables that were available for the entire sample period (Chow and Lin, 1971) such as total number of employees, sales, payroll, and value of shipments per employee for which a complete time-series data were available. These regressions yielded an average R-square of 85%. Then the *H* indexes were estimated with the predicted values from these regressions. The forecasted values were quite smooth in relation to the *H* indexes for the Census years. The use of spline functions and exponential smoothing (using the CurveExpert software) did not significantly altered the results. Finally, it should be noted that food-processing industries for which the 1982 Herfindahl indexes were not available (due to reclassification of the SIC codes in 1987) were excluded from the sample.

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