Market Structure and the Dynamics of Retail Food Prices

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The effect of retail grocery market structure on the speed of adjustment of retail food prices to changes in producer prices, real wages, and the cost of energy was examined for SMSAs. Evidence failed to support the implication of the Mason-Bain paradigm that increased concentration reduces market efficiency as reflected in speed of retail price adjustment. Evidence of strong intertemporal relationships between change in producer prices and retail prices found for the categories meat, poultry, fish, eggs and cereal and baker products provide support to the hypothesis of cost-push inflation.

Introduction

Whether coordination occurs explicitly, or through a gaming process, information flow and communication have been recognized as crucial determinants of the success of output and price control. The presence of imperfect information, high degrees of product differentiation, and search costs observed in contemporary markets has motivated the so-called modern market theories of Schmalensee (1987), Salop and Stiglitz among others, which argue that concentration may enhance efficiency of information collection and processing leading to lower prices. These theories directly contradict the predictions of the Mason and Bain paradigm that concentration leads to higher profits and prices. In the absence of resolution at a theoretical level, the effect of market structure on market performance is clearly an empirical issue that deserves consideration.

The objective of this paper is to analyze the impact of market structure on a dimension of economic performance not typically considered: the speed of adjustment of prices to changes in cost structure as reflected by changes in input prices. By traditional perception, changes in input prices would lead a competitive market price to adjust through a process of entry-exit forced by changes in profits. In contrast, the presence of output and price control in a concentrated market of imperfectly competitive firms implies coordination and this may suggest efficiency in adjustment of prices to maintain desired profit levels. As Domberger noted:

> "The implications for dynamic pricing behavior are not difficult to rationalize in concentrated industries where costs of search and communication among sellers are relatively low, price adjustments can be effectively coordinated and equilibrium in the industry restored fairly rapidly. . . . The prediction which emerges (is) that the speed of price adjustment will tend to rise with the level of industrial concentration, ..."

This relation between speed of adjustment and market structure also emerges from the modern market theorists cited above. As Salop and Stiglitz have argued, in the presence of imperfect information, stability in sales may require rapid adjustment in prices to prevent consumers from searching for lower prices and switching stores. Price dispersion and, therefore, sales uncertainty are argued to increase with the cost of information. It follows from this modern market theory that when input prices vary over time, firms have a strong incentive to adjust prices quickly, and in harmony, to minimize resulting price dispersion, high search costs for consumers, and possible loss in sales.

Effects of market structure on speed of price adjustment are implied by a variety of

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theories, and their existence is suggested by evidence of a relation between market structure and prices and profits. Nonetheless, the nature of this relation is unresolved by either the traditional theories following Mason and Bain, or more recent modern market theories, leaving it an empirical question. The following sections will present evidence concerning the relationship between the market structure of retail grocery markets and the speed of adjustment of retail grocery prices to changes in inut prices as reflected by raw product prices and wages. The conceptual analysis has broader application to the question of dynamics of price adjustment and the role of market structure. Studies of the retail grocery industry have generally established a relationship between structure and either profit level (see Marion et al. or Hall et al.) or price level (see Lamm (1981) or Cotterill) however, the effect of structure on speed of adjustment has been left largely unexplored.

Model Specification

The relationship between market structure and price level or dynamics requires economic motivation. Consider a market equilibrium price function derived from a micro-economic theory of firm behavior for the imperfectly competitive case. For example, for the case where output pricing power exists, the first-order condition for the profit-maximizing of the j^{th} output choice by the h^{th} firm in the i^{th} market implies the following market specific price function for each output j:

$$(1) \qquad P^*_{ij} = [\partial C^h_i/\partial Q^h_{ij}][N_{ij}/(N_{ij} + \theta^h_{ij})]$$

where

$\partial C^h_i/\partial Q^h_{ij}$ is the marginal cost for the h^{th} firm, i^{th} market, j^{th} output,

$N_{ij} = (\partial Q_{ij}/\partial P_{ij})\,(P_{ij}/Q_{ij})$, and

$\theta^h_{ij} = (\partial Q_{ij}/\partial Q^h_{ij})\,(Q^h_{ij}/Q_{ij})$

For particular specifications for the cost function, market demand function, and reaction functions, (1) implies output price can be written as a function of the determinants of those functions. Since $\theta^h_{ij} = 0$ for the case of perfect competition, $\theta^h_{ij} = 1$ for monopoly, and $0 < \theta^h_{ij} < 1$ for other cases of pricing power, ϑ^h_{ij} reflects market power. Expression (1) can, therefore, be interpreted as establishing the relationship between market specific price levels and the structure of market power within that market. Appelbaum (1979, 1982) and Schroeter have recognized this in approaches to parametrically estimate θ^h_{ij} as a measure of market power. Oligopoly theories of Cournot, Chamberlain, and Stigler as well as modern market theories suggest market power is determined by a vector Z_{ij} of market structure characteristics suggesting that θ^h_{ij} may be written:

$$(2) \qquad \theta^h_{ij} = \theta^h(Z_{ij})$$

Both Lamm (1981) and Cotterill recognized the implied relationship between market specific price level and market power in studies of retail grocery pricing in SMSAs and local markets, respectively. In each case, they found aggregate food price index levels to be significantly and positively affected by market power as measured by concentration ratios, market shares, or critical levels of market share. Solution of reaction functions (1), together with (2) imply the existence of an industry level relationship between market specific prices and market structure:

$$(3) \qquad P^*_{ij} = P_{ij}(R_j, \phi_j, Z_{ij})$$

where

R_j and ϕ_j are arguments of C^h_i reflecting input prices and fixed factors.

By its derivation from (1), (3) implies instantaneous adjustment in output price to changes in its determinants. Modern market theories cited above suggest that the presence of imperfect information, costs of adjustment, and costs of coordination may imply the existence of inertia in adjustment, and variation in inertia across markets with different market structures. These differences imply the function $P_{ij}(\cdot)$ would be expected to vary across markets.

Specific dynamic adjustment models could be derived from specific theories of firm behavior, e.g. a general first-order differential might be implied (dropping subscripts):

$$(4) \quad dP_t = (1 - B)P_t = f(P^*_t - P_{t-1})$$

where

P^*_t is the equilibrium price given firm behavior.

The partial adjustment model consistent with minimization of costs of adjustment used by Griliches and adopted by Domberger is a special case of (4). The implication of (3) and (4) is that the observed market price P_t can be written in a dynamic reduced form determined by a vector of current and lagged determinants of P^*_t. To proceed, for each market area (i) and j^{th} retail product category, a generalized reduced form of (3) and (4) may be written:

(5) $P_{ij} = C_{ij}\beta_{ij}(L) + e_{ij}$
for $i = , \ldots n$ and $j = , \ldots m$

where

- e_{ij} is drawn from a covariance stationary univariate stochastic process with plim $e_{ij} = 0$, asymptotic variance-covariance $\sigma^2_{ij} I_T$, and plim $e_{ij}\, c_{ij} = 0$,
- P_{ij} is a $T \times 1$ vector of retail price observations,
- C_{ij} is a $T \times K$ matrix of the determinants of market specific price level, e.g., R_j, ϕ_j, Z_{ij} in (3), and
- $\beta_{ij}(L)$ is a $K \times 1$ vector of polynomials in the backward shift opertor L.

Equation (5) can be interpreted as a dynamic reduced form of a multiple product price adjustment process.

While the parameters of $\beta_{ij}(L)$ are considered fixed over time, their variation across markets provides the basis for investigation of the relation between market structure and the speed of adjustments of price as reflected by the form of $\beta_{ij}(L)$. In principle, the following approach will be taken. Dropping the product subscript j, define $S(\beta_i)$ as a $K \times 1$ vector function of the elements of β_i to be interpreted as a statistic summarizing the speed of price adjustment in market i to each respective determinant. In order to investigate the effects of market structure on $S(\beta_i)$, the following linear model is proposed:

(6) $$S(\beta_i) = Z_i\gamma + v_i$$

where

- Z_i is a $K \times P$ vector of measures of market structure in market
- γ is a $P \times 1$ vector of parameters, and
- v_i is identically, and independently distributed with plim $V_i = 0$, and plim $Z_i'V_i = 0$.

Definition of $S(\beta_i)$ results in characterization of the temporal distribution of adjustment. Four alternatives are considered. For the k^{th} determinant of retail price define:

Total Magnitude of Response:

$$\hat{S}^k_{1i} = \sum_{\tau=0}^{\infty} \hat{\beta}^k_{it-\tau}$$

Mean Lag of Response:

$$\hat{S}^k_{2i} = \sum_{\tau=0}^{\infty} \hat{\beta}_{it-\tau}/S^k_{1i}$$

Percentage of Response:

$$\hat{S}^k_{3i}(t - \tau) = \hat{\beta}^k_{it-\tau}/S^k_{1i}$$

Level of Response:

$$\hat{S}^k_{4i}(t - \tau) = \hat{\beta}^k_{it-\tau}$$

where

- $\beta^k_{t-\tau}$ is the coefficient of the $t - \tau$ order in the k^{th} polynomial in β_i.

Consistent estimates of $\beta^k_{it-\tau}$ follow from ordinary least squares (OLS) estimation of (5) and the assumed asymptotic properties of e_{ij}. Using Slutsky's Theorem, it follows that $\hat{S}^k_i$ for each definition are consistent estimates of S^k_i based on $\beta^k_{it-\tau}$. The following equation is implied for each definition of S^k_i:

(7) $$\hat{S}^k_i(\hat{\beta}_i) = S^k_i(\beta_i) + w_i,$$

where

- w_i is identically and independently distributed with plim (w_i) = plim $(w_i e_{ik})$ = plim $(w_i v_i) = 0$.

Finally, combining (6) and (7) single equation OLS provides consistent estimates of γ. Asymptotic efficiency of estimates of γ and β_{ij} follows directly from the assumed stochastic properties. In the simplest case where $S^k(\beta_i) = \beta^k_{it-\tau}$, (5) and (6) can be interpreted as a variation of the random coefficient model of Swamy. Results presented by Duncan for M-estimators for sequential estimation may also be used to establish consistency of these estimates.

Empirical Implementation

Empirical implementation of (5) and (6) for the retail food industry requires specification of sample characteristics such as definition of products, store type, and geographic market, and identification of appropriate and available data. The retail food industry is composed of supermarkets with highly diversified food and home products, grocery stores which are specialized in food products, and small specialty or convenience stores. Within any market area the composition of stores is relatively stable over time, while across markets greater differences would be expected to reflect differing consumer preferences. The focus of this study will be on the retail food industry as a whole rather than on a particular store type.

Food products are widely recognized to be marketed in unique local markets, the dimensions of which are defined by travel costs and, as implied by modern market theories, search costs. In this study the market area will be defined as an SMSA. This specification has also been adopted as an approximately accurate market area for retail foods by Lamm (1981), and Geithman and Marion.

The time period of 1969–81 was selected as a period in which input prices for food retailing were effected by significant general inflation processes. Monthly Bureau of Labor Statistics (BLS) data are available for twenty-three SMSAs for food product category CPIs and individual product prices. Products sampled by BLS vary across markets according to an observed base period sales volume. This variation led Geithman and Marion to conclude that use of BLS CPI indexes for cross-sectional study of structure-price relationships may result in a negative bias in estimates of any concentration-price relationship. For the current study, the BLS data is used to determine speed of adjustment prices over time, a use for which BLS prices were designed (Rothwell).

The extent that cross-sectional variation in base period product category composition is reflected in changes in prices over time is difficult to determine. Geithman and Marion argued that BLS sampling methods are biased toward high volume store brands leading these low-priced products to be included more frequently in concentrated than unconcentrated markets. By implication, as recognized by Lamm (1979, 1981), price index levels are negatively biased in concentrated markets. Whether this sampling bias would bias price adjustment measures depends on whether price adjustment varies by product volume. To the extent that high volume products are highly visible to the consumer, modern market theories would suggest price adjustment for these products would be faster. The conclusion could be drawn that BLS sampling procedures may bias price adjustment measures to reflect faster adjustment in concentrated markets than unconcentrated markets. To the extent that this bias occurs, results reported here would overstate actual relationships between speed of adjustment and market concentration.

A wide range of specifications of product types and levels of aggregation (both within and across type) would be both appropriate and of interest in the study of market structure effects on the speed of price adjustment. Lamm (1981) used a product category annual cost, while Cotterill and others have used store level market baskets. The present study will limit its focus on the three categories of products for which BLS category CPI indexes are available: (1) cereal and bakery products, (2) meats, poultry, and fish, and (3) fruits and vegetables. These categories account for important portions of U.S. farm production, yet they present variation in the extent of processing. By use of category indexes rather than the aggregate food price index, prices reflect variation of product categories across the sampled SMSAs. Respectively, in 1982 these categories accounted for 14%, 32%, and 16% of average weekly per person at home food expenditures for U.S. urban households.

The composition of the BLS categories is as follows: cereal and bakery products (CB) includes basic grains, milled and unmilled (e.g. wheat flour and rice), and bakery products such as bread, crackers, and cookies; meats, poultry, fish, and eggs (MPFE) focuses on whole and cuts of fresh meat with the exception of canned tuna and sardines and fresh eggs; fruits and vegetables (FV) includes fresh, frozen, and canned fruits and vegetables. Producer prices for CB and MPFE for the same categories were employed as available on a monthly basis from BLS. BLS changes the category definition between the CPI for FV and the PPI for FV where only fresh fruits and vegetables are sampled. Because wholesale or producer prices are determined in nation-wide markets rather than local SMSA markets, national aggregate producer prices were employed.

In addition to wholesale and raw product prices, prices of other inputs involved in retailing food products must be considered as determinants of the retail price level. Although omitted by Cotterill, German and Hawkes find wages and energy account for the predominant proportion of the cost of providing retail services. The roles of these input prices will be explored below. Labor costs were collected from county level unemployment insurance data available for ten SMSAs from BLS monthly employment and quarterly payroll reports for the retail food industry. Use of county level data allows a close approximation to SMSA labor markets. *Employment and Earnings* data was unavailable for the retail food industry on a monthly basis except at a national aggregate level. The monthly SMSA level CPI for energy based on the cost of gas and electricity was used for the price of energy.

Market Structure

Measurement of market structure by the elements of Z_i in (6) follow from theories of oligopoly (Cournot, Chamberlin, and Stigler) which suggest that the number of firms participating, disparity in size, and achievement of thresholds in output control affect the feasibility and effectiveness of price setting power. The implication of modern market theories is that imperfect information leads to coordination costs which increase with (1) the structure of sales distribution as measured by market concentration, (2) market growth, or the stability of market size, and (3) the existence of barriers to entry.

Measurement of market concentration requires choice of a statistic that reflects variation in the extent of market dominance or price setting power, as well as the distribution of that power among firms. Alternatives adopted in empirical studies have ranged from individual firm market shares (Kwoka 1979, 1981), and 4-, 8-, and 20-firm concentration ratios, to Stigler and Schmalensee's use of Herfindahl indexes to describe the distribution of firm level sales. Lamm (1981) found 2-, 3-, and 4-firm concentration ratios to be positively related to retail food price levels though use of firm market shares improved model fit marginally. More recently, Cotterill found the Herfindahl index, market shares, and market concentration ratios to effectively represent market structure as a determinant of retail food price levels in a cross-section of Vermont firms. However, Cotterill finds the Herfindahl index to outperform market share measures and market share measures to outperform market concentration ratios, a result also found by Shephard, Ravenscraft and Marion et al. In this study, data availability limits the enquiry to the use of firm concentration ratios. The 4, 8, and 20 firm concentration ratios (CR4, CR8, and CR20) are employed as reported for 1972 in *Grocery Retailing Concentration* (BLS). These ratios reflect different perspectives on the distribution of market power among firms. While the 4-firm measure focuses on the extent of highly concentrated market power, the 8-firm measure indicates the extent to which a larger group of firms may exist which could coordinate prices. In contrast to the 4-firm ratio, the 20-firm ratio indicates the extent to which market sales are claimed by a group of firms which is large enough that coordination costs may preclude expression of market power.

In addition to the sales structure of an industry, traditional oligopoly theory as well as more recent theories of entry deterrent pricing (e.g. Gaskins) suggest that the extent of barriers to entry influences the firm's pricing strategy and, thereby, an effect on the speed of adjustment would be expected. Kwoka (1981) has found a variety of measures of scale of plant size to be positively related to the level of the price-cost margin. In studies of retail food prices, Lamm (1981) has found average store size (measured by the number of market baskets sold per store) to be significant and negatively related to price level. This result suggests scale economies resulted in cost reductions rather than barriers to entry which led to increased prices. Cotterill also introduced store size as a determinant of firm level prices and found it insignificant in linear models, though significant when introduced in quadratic form. These results were interpreted as indicating an initial negative effect on prices as scale reduced costs; however, as scale continues to increase Cotterill argued store differentiation is achieved through expanded product line which establishes market power to set higher prices. In contrast, Salop and Stiglitz's modern market theory would predict the effect of scale on price level and speed of adjustment to be negative, and positive, respectively. In this study, firm size is measured by firm average square feet of retail sales area for 1972 as reported in *Grocery Retailing*.

Market growth has been argued to result in increased prices and profit levels as oligopolists attempts to ration available capacity. Although Kwoda (1979) among others have found significant positive relations between market growth and prices, others have found negative relations. Scherer, Schmalensee, and Parker and Connor found a significant and positive relationship in a price-cost margin model as well as one explaining the difference between store and manufacturer brand prices. Cotterill found an insignificant and negative relationship with firm level prices. Modern market theories would suggest that as market growth implies instability in past results of coordination, firms may be expected to reduce prices, and adjust them more quickly in response to exogenous factors. Market sales growth were employed as a measure of market growth as available from Bureau of Census, Retail Trade Division, Subject Statistics for 1972.

Empirical Results

Dynamic reduced forms such as (5) were estimated for the three product categories (CB, MPFE, and FV) for the twenty-three SMSAs for which CPI indexes are available. Prior to estimation, sufficient conditions for covariance stationarity were established for all price variables by first differencing in logarithms. For the MPFE CPI series, annual seasonality was also removed by differencing.[1] After prefiltering, the dynamic reduced forms were estimated as linear forms implying variables can be interpreted as rates of change and estimated parameters as coefficients of rates of change, or elasticities of the CPI to a change in the PPI.

The most general model including PPI, wages, and the price of energy was estimated for ten SMSAs. Results indicated that wages were consistently insignificant (out to sixteen lags) across products and SMSAs. The dynamic reduced forms were re-estimated without wages for the complete set of twenty-three SMSAs for which data are available. For all SMSAs for MPFE and CB, the lag structure for the PPI indicated statistically insignificant effects beyond the sixth period. Table 1 summarizes the estimated lag structures for the PPI. For the FV category, no statistically significant relation was found between the PPI and the CPI. In part, this may have been due

[1] Independent variables in time series are definitionally stochastic regressors. Covariance stationarity is required to ensure that the sample information matrix converges asymptotically to the underlying covariance matrix which characterizes the stochastic process which generated the sample. Sample specific prefiltering to achieve sufficient conditions for covariance stationarity is a prerequisite for establishing an asymptotic relation between estimated sample and population statistics, see Goldberger or Fomby, Hill, and Johnson.

Table 1a. Response of Percent Change in CPI to Percent Change in PPI by SMSA: Meat, Poultry, Fish, and Eggs Category

	Lag (months)								
City	0	1	2	3	4	5	6	R^2	D.W.
Atlanta	0.3983	0.2195	0.1997	—	—	—	—	.63	2.37
Baltimore	0.3140	0.2371	0.1675	—	—	—	—	.64	2.29
Boston	0.2644	0.2085	0.1562	—	0.0955	—	—	.57	2.34
Buffalo	0.4079	0.2058	0.2029	—	—	—	—	.60	2.09
Chicago	0.4155	0.2095	0.1293	—	0.0881	0.0857	—	.58	2.64
Cincinnati	0.4276	0.2825	0.1070	0.0753	—	—	—	.63	2.48
Cleveland	0.4381	0.3415	0.0892	—	—	—	—	.64	2.42
Dallas	0.4193	0.3459	0.0763	—	0.0921	—	—	.69	2.12
Detroit	0.4988	0.2617	—	—	0.1181	—	−0.0745	.65	2.31
Honolulu	0.0654	0.2175	0.1678	0.0567	—	—	0.0566	.51	1.98
Houston	0.4236	0.2793	0.1344	—	—	—	—	.63	2.19
Kansas City	0.3805	0.2885	0.1634	—	0.0874	—	—	.61	2.24
Los Angeles	0.3730	0.2926	0.0783	—	0.0857	—	—	.65	2.23
Milwaukee	0.3401	0.2500	0.1581	—	—	—	—	.61	2.12
Minneapolis	0.3467	0.2126	0.1488	0.0701	0.0698	—	—	.62	2.36
New York	0.3429	0.2040	0.1467	—	—	—	—	.63	2.34
Philadelphia	0.3821	0.2376	0.1112	—	0.1325	—	—	.60	2.41
Pittsburgh	0.3762	0.2607	0.1488	—	—	—	—	.55	2.38
St. Louis	0.4307	0.2475	0.1445	—	0.0941	—	—	.64	2.33
San Diego	0.4317	0.3019	—	—	0.0822	—	—	.67	1.99
San Francisco	0.3903	0.2878	0.0744	—	0.1199	—	—	.64	2.31
Seattle	0.3197	0.2607	0.1438	—	0.1395	—	—	.54	2.38
Washington, DC	0.3826	0.2811	0.1526	—	—	—	—	.56	2.41

Only parameter estimates significant at the 0.1 level (using two-tailed tests) are reported.

to differences in category definition. The CPI includes frozen and processed, while the PPI includes only fresh. Further analysis of the FV category was not pursued.

Because covariance stationarity was established before estimation it is not surprising that estimated residuals were found to be white noise in all cases. The estimated lag structure suggests that substantial adjustment of the MPFE product CPI occurs within two months, while for the CB product category the lag structure varies significantly across SMSAs. These results are consistent with those found by Lamm (1981) and Lamm and Westcott for national aggregate data.

The estimated lag structures for CPI adjustment to changes in PPI were employed to estimate five alternative SMSA specific speed of adjustment statistics. Substantial variation in each of these statistics was apparent across SMSAs. Interpretation of the speed of adjustment coefficients follows directly from the interpretation of estimated coefficients as elasticities. The magnitude, mean, and percentage response statistics represent the sum, mean, and period specific percentage of the elasticity of the CPI for change in the PPI over the lag structure identified and estimated. Estimates of equation (6) are reported in Table 2 and 3 for each of the speed of adjustment coefficients, and for alternative concentration ratios.

For the cereal and bakery product category, Table 2 indicates variation in total magnitude, and percentage responses in t, t-1, and t-3 appears related to market structure. In all cases, results are robust across different concentration ratios indicating that concentration is a significant positive determinant of total magnitude and percentage of response during the current period t and period t-3. A negative relation is indicated for period t-1. These results provide empirical evidence in support of modern market theories which hypothesize that concentration increases speed of adjustment through reduction in costs of and increase in feasibility of coordination among firms. As Stiglitz has noted, this implies an efficiency gained due to concentration over purely competitive structures where coordination is impossible, and information and search costs are high.

Barriers to entry as measured by firm average floor space was found a significant and negative determinant of total magnitude and percentage response in period t, a result consistent with that found by Lamm and Cotterill in models of price levels. A significant positive

Table 1b. Response of Percent Change in CPI to Percent Change in PPI by SMSA: Cereal and Bakery Products Category

	Lag (months)								
City	0	1	2	3	4	5	6	R^2	D.W.
Atlanta	—	0.3227	0.2028	0.2411	—	—	0.2607	.55	2.74
Baltimore	0.3631	—	0.3582	0.2381	−0.2332	—	—	.48	2.45
Boston	—	0.3795	0.2853	—	—	—	—	.56	2.46
Buffalo	0.2434	0.2214	0.2285	—	—	—	—	.50	2.61
Chicago	—	0.3956	0.2185	0.1962	—	—	—	.58	2.32
Cincinnati	0.5675	—	—	0.2232	—	—	0.2871	.47	2.48
Cleveland	0.2582	0.3434	—	0.2709	—	—	—	.37	2.29
Dallas	0.2483	0.2347	—	—	—	—	—	.53	2.50
Detroit	—	0.5634	—	—	—	—	—	.18	2.88
Honolulu	0.2052	—	0.5587	0.4153	—	0.2873	—	.56	2.10
Houston	0.1730	0.3900	0.3292	—	—	—	—	.57	2.29
Kansas City	0.1683	0.4869	—	—	—	—	—	.59	2.28
Los Angeles	—	0.2421	0.3398	0.1726	—	—	—	.62	2.09
Milwaukee	—	—	—	—	—	—	—	.19	2.82
Minneapolis	—	0.4500	—	0.2477	−0.2665	—	0.2385	.59	2.14
New York	0.2334	0.2847	0.1402	0.1891	−0.1873	—	—	.69	2.38
Philadelphia	—	0.3620	—	—	—	—	—	.56	2.31
Pittsburgh	0.2879	0.3745	—	0.1838	−0.2761	—	0.2195	.57	2.43
St. Louis	—	0.3647	—	—	—	—	—	.47	1.99
San Diego	—	0.3942	0.1930	—	—	—	—	.52	2.45
San Francisco	—	0.4968	0.2117	—	—	—	—	.68	2.13
Seattle	—	0.1771	0.2429	0.2073	—	—	0.2525	.57	2.38
Washington, DC	0.2674	—	—	0.3207	−0.3099	0.2658	—	.39	2.21

Only parameter estimates significant at the 0.1 level (using two-tailed tests) are reported.

Table 2. Market Power and the Speed of Adjustment of Retail Food Prices: Cereal and Bakery Products

Speed of Adjustment Statistic	Concentration Ratio	Intercept	Concentration Ratio	Barriers to Entry	Market Growth	R^2	Dep. Mean
Total Magnitude	4-firm	.69	.0044	−.000038	.11	.26	.74
		(2.49)	(1.35)	(1.80)	(.49)		
	8-firm	.76	.0029	−.000039	.079	.22	
		(2.56)	(.93)	(1.78)	(.33)		
	20-firm	.71	.0037	−.000044	.066	.21	
		(2.02)	(.85)	(1.89)	(.27)		
Mean Lag	4-firm	1.09	.015	.000051	.16	.10	2.26
		(.92)	(1.11)	(.56)	(.17)		
	8-firm	.85	.017	.000044	.14	.12	
		(.71)	(1.34)	(.49)	(.15)		
	20-firm	.39	.025	.000012	.084	.14	
		(.27)	(1.41)	(.13)	(.089)		
Percentage Response (t)	4-firm	.41	.0033	−.00005	−.020	.20	.26
		(1.12)	(.78)	(−1.79)	(.06)		
	8-firm	.46	.0022	−.00005	−.061	.18	
		(1.20)	(.54)	(1.79)	(.16)		
	20-firm	.45	.0024	−.00005	−.061	.18	
		(.99)	(.43)	(1.80)	(.20)		
Percentage Response (t − 1)	4-firm	.86	−.0098	.000042	−.38	.27	.41
		(1.93)	(1.89)	(1.21)	(1.02)		
	8-firm	1.16	−.013	.000062	−.31	.29	
		(2.16)	(2.03)	(1.76)	(.86)		
	20-firm	.91	−.0094	.000045	−.34	.28	
		(1.97)	(1.92)	(1.31)	(.93)		
Percentage Response (t − 2)	4-firm	−.070	−.0014	.000017	.44	.12	.21
		(−.17)	(−.28)	(.55)	(1.26)		
	8-firm	−.15	−.000099	.000017	.45	.11	
		(.34)	(.022)	(.54)	(1.34)		
	20-firm	−.29	.0021	.000014	.48	.12	
		(.56)	(.33)	(.41)	(1.41)		
Percentage Response (t − 3)	4-firm	−.21	.00094	−.000019	.031	.34	.15
		(.70)	−(2.67)	(.85)	(.12)		
	8-firm	−.14	.0072	−.000022	−.032	.24	
		(.42)	(2.04)	(.89)	(.12)		
	20-firm	−.14	.0074	−.000030	−.079	.15	
		(.34)	(1.43)	(1.11)	(.28)		

t-statistics are presented in parentheses.

relationship with percentage response in periods t-1 was found. Results were, in general, robust across models with different concentration measures. The results suggest that slower immediate and less total adjustment as scale increases. After one period, in t-1 larger scale firms adjust faster in the period. These results are consistent with increased scale introducing diseconomies into the speed of firm reaction to changes in cost. While increased concentration appears to enhance coordination and speed of adjustment among firms, increased size of firm appears to decrease the firm's speed of adjustment. Market growth appears to be a consistently insignificant determinant of speed of adjustment, a result also found by Cotterill in a model of price levels. Overall, the results appear to support the modern market theories, not the Bain and Mason paradigm.

Results for Meat, Fish, Poultry, and Eggs are reported in Table 3. For this product category, results vary substantially across speed of adjustment statistics. Results fail to indicate any relationship between percentage response in t, t-2, and t-3 and measures of market struc-

Table 3. Market Power and the Speed of Adjustment of Retail Food Prices: Meat, Fish, Poultry, and Eggs

Speed of Adjustment Statistic	Concentration Ratio	Intercept	Concentration Ratio	Barriers to Entry	Market Growth	R^2	Dep. Mean
Total Magnitude	4-firm	.86	−.0016	.000021	−.096	.20	.85
		(5.09)	(.84)	(1.60)	(.67)		
	8-firm	.87	−.0016	.000022	−.09	.20	
		(5.01)	(.89)	(1.65)	(.65)		
	20-firm	.92	−.00024	.000025	−.085	.21	
		(4.46)	(.95)	(1.81)	(.62)		
Mean Lag	4-firm	2.70	−.0064	.000038	−1.98	.49	1.93
		(6.40)	(1.33)	(1.18)	(3.38)		
	8-firm	2.75	−.0065	.000041	−1.17	.49	
		(6.36)	(1.42)	(1.27)	(3.39)		
	20-firm	2.89	−.0088	.000052	−1.15	.49	
		(5.62)	(1.39)	(1.53)	(3.33)		
Percentage Response (t)	4-firm	−.072	.00032	.000010	.18	.079	.11
		(.36)	(.139)	(.67)	(1.078)		
	8-firm	−.077	.00036	.000010	.18	.079	
		(.36)	(.16)	(.66)	(1.088)		
	20-firm	−.13	.0011	.0000087	.18	.086	
		(.51)	(.37)	(.53)	(1.12)		
Percentage Response (t − 1)	4-firm	.38	.0007	−.000012	.18	.41	.45
		(4.27)	(.76)	(1.75)	(2.43)		
	8-firm	.37	.00086	−.000012	.18	.42	
		(4.03)	(.88)	(1.81)	(2.45)		
	20-firm	.36	.0010	−.000013	.17	.47	
		(3.29)	(.75)	(1.88)	(2.39)		
Percentage Response (t − 2)	4-firm	.25	.00089	−.0000063	.072	.047	.297
		(1.56)	(.48)	(.51)	(.53)		
	8-firm	.21	.0013	−.0000069	.075	.068	
		(1.29)	(.78)	(.56)	(.57)		
	20-firm	.17	.0020	−.0000095	.072	.074	
		(.89)	(.84)	(.74)	(.55)		
Percentage Response (t − 3)	4-firm	.14	.00098	−.0000084	.039	.071	.16
		(1.03)	(.64)	(.82)	(.35)		
	8-firm	.19	−.000071	−.0000083	.020	.047	
		(1.43)	(.048)	(.79)	(.18)		
	20-firm	.26	−.0010	−.0000067	.014	.061	
		(1.58)	(.49)	(.61)	(.13)		

t-statistics are presented in parentheses.

ture. However, a strong significant relation is found for mean lag and percentage response in t-1 (the period where mean percentage response is largest). Results for total magnitude of response suggest a significant positive role of average firm size, opposite of that found for cereal and bakery products. Results are not invariant across speed of adjustment statistics. For the mean lag, market concentration is negatively related to speed of adjustment, a result that is robust across models with different concentration ratios. Barriers to entry as measured by average firm size is positively related to mean lag. In both cases, the concentration ratio and the barriers to entry measure are only marginally significant. Market growth, in contrast to findings for cereal and bakery products, is a significant, negative determinant. This result is consistent with the modern market theories which suggest that in expanding markets, firms adjust prices more rapidly as a strategy for maintenance of market share.

The distribution of speed of adjustment across periods is reflected in the percentage of response statistics. The model of this statistic for t-1 indicates that barriers to entry have a consistently negative significant effect on the within period adjustment. As firm size increases speed of adjustment slows. This same result was found for cereal and bakery products. Market growth has a positive and significant effect with period response in t-1, a result consistent with the finding that market growth reduces mean lag of response. One result found consistently across speed of adjustment coefficients was the absence of a role of concentration. Across all speed of adjustment coefficients and concentration ratios, the concentration ratio was insignificant except for the mean lag where it is only marginally significant.

Conclusions

Specific conclusions can be drawn from the results presented above. Variation in the rate of change of the retail food prices over the time period studied can be explained by variation in raw material prices measured by the PPI and energy prices. This confirms that raw material prices play an important role in retail food price inflation processes. This result suggests that effective agricultural price stabilization programs may lead to benefits realized through impacts on the retail food price inflation process. Wages did not appear to play a significant role; however, difficulty of measuring this variable may have led to this weak relationship. The speed of adjustment of retail prices was found by vary substantially across SMSAs and this variation was explained by market structure. Results appear invariant to the definition of the concentration ratio. Data availability precluded the use of other measures of market structure such as market shares or the Herfindahl index.

The absence of a role of concentration measures for the meat category (MPFE), and a generally positive relationship found for cereal and bakery products provide support for the modern market theory that concentration may enhance efficiency in price adjustment and, for the case of price adjustment, fails to support the implication that may be drawn from the Bain-Mason paradigm that concentration reduces the efficiency of market performance. Importantly, results support the conclusion that consumers may be benefited by concentration which leads to more rapid price adjustment.

Overall the results for these product categories provide evidence that is consistent with the hypothesis that market structure affects the speed of price adjustment in the retail grocery industry. The evidence does not reject this hypothesis; however, it can not be judged as strongly supporting the hypothesis. Importantly, the results fail to support the implications of the Mason-Bain paradigm and, instead, provide support for the modern market theories which suggest concentration may enhance some market performance characteristics.

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In Figure 2 on page 30, the cross-hatched area indicates where average date of last 28° freeze occurs after May 1.