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Structural Change and Competition in Seven U.S. Food Markets

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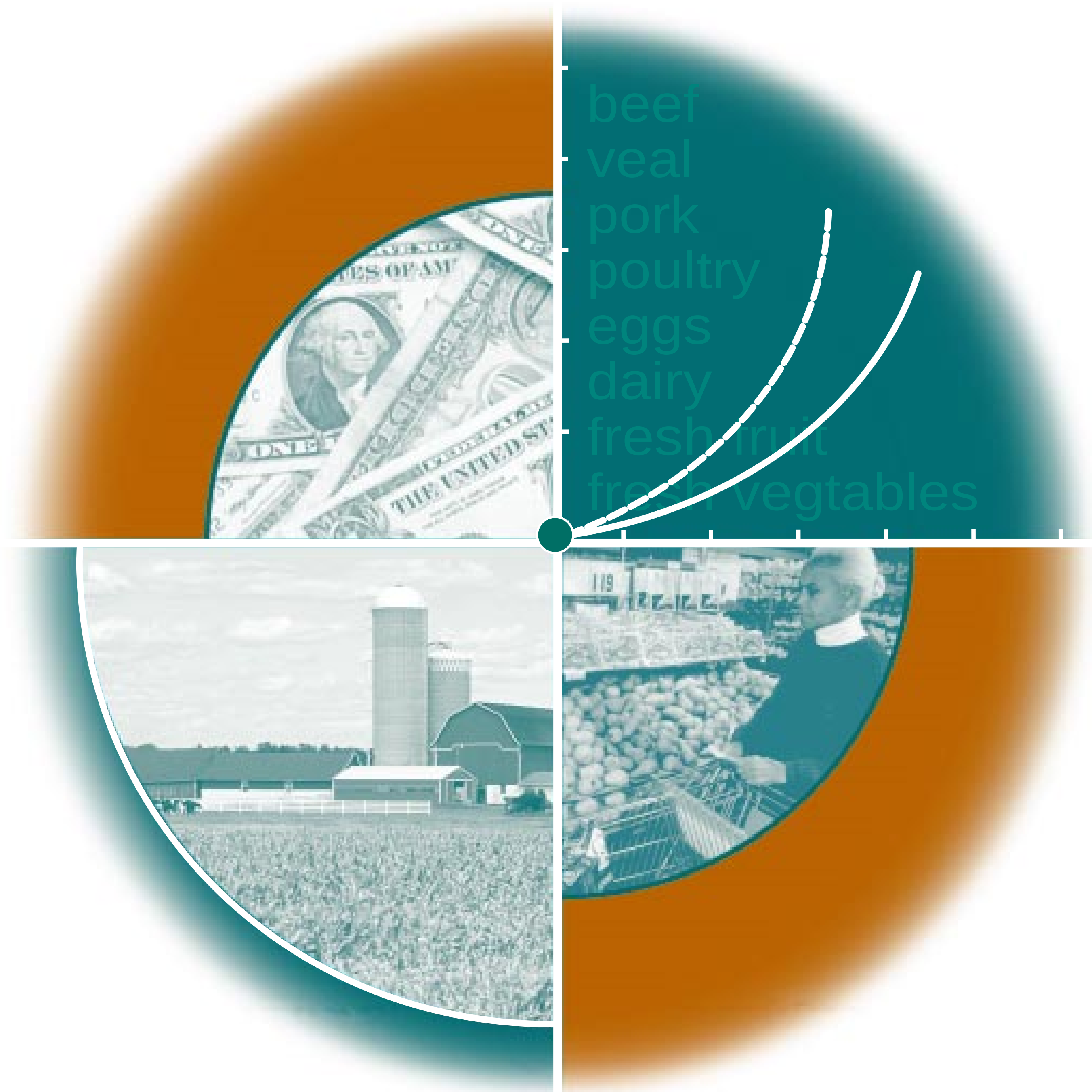

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Abstract

Recent trends in mergers and acquisitions in the U.S. food sector – food manufacturers, wholesalers, and retailers – raise concerns about market power. In the presence of market power, farmers may receive lower than competitive farm prices, and consumers may pay higher than competitive retail prices. This study presents empirical tests of market power at the national level for seven food categories: beef, pork, poultry, eggs, dairy, fresh fruit, and fresh vegetables. At the national level, our tests provide evidence of competitive conduct in both the sale of final food products and the purchase of farm ingredients.

Keywords: Retail food and farm prices, market power, structural change, cointegration.

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Summary

Recent trends in mergers and acquisitions in the U.S. food sector – food manufacturers, wholesalers, and retailers – raise concerns about market power. In the presence of market power, farmers may receive lower than competitive farm prices, and consumers may pay higher than competitive retail prices. This study presents empirical tests of market power at the national level for seven food categories: beef, pork, poultry, eggs, dairy, fresh fruit, and fresh vegetables.

The procedures employed in this study account for three important features of food markets. First, the procedures recognize that within a product category consumers prefer a variety of food items. Second, the procedures account for firm diversity by recognizing that different firms produce a variety of different products using different technologies. Firm and technological diversity seem to be particularly relevant to food industries because mergers and acquisitions may be feasible only if firms and technologies are diverse. Third, the procedures recognize that structural changes in food markets may be unpredictable. Some of the most significant impacts on food markets may have been unpredictable. For example, it was probably difficult to predict trends in consumer patterns following the entry of women into the labor force in the 1970's and 1980's. Unpredictability of consumer behavior poses considerable risk to food producers and can induce industrial reorganizations that spread this risk across stages of food production. Failure to account for product diversity and uncertainty has been shown to seriously affect tests of market power.

The empirical evidence presented in this bulletin is mostly consistent with competitive conduct. At the national level, our tests provide evidence of competitive conduct in both the sale of final food products and the purchase of farm ingredients. Also at the national level, the tests suggest that food industries pay competitive prices for farm commodities. The results do not rule out imperfectly competitive local markets, and they cannot be used to address questions of relative market power between specific stages of food production. Nevertheless, the broad implications of these results may be that it is the unpredictable nature of trends in consumer demand, and not imperfect competition, that may be responsible for observed trends in the industrialization and concentration of the U.S. food sector.

Structural Change and Competition in Seven U.S. Food Markets

A.J. Reed and J.S. Clark

Introduction

This bulletin presents empirical analyses of market structure and competition for seven major U.S. food markets: beef, pork, poultry, eggs, dairy, fresh fruit, and fresh vegetables. The analyses account for structural change. Our work emerges from a theory first developed by Heiner in which consumer demand and input supply mesh with the output supply and the input demands of nonidentical firms. While Chavas and Cox have recently generalized this theory, Wohlgenant (1989) and Wohlgenant and Haidacher (WH) were the first to extend and apply it to food markets. This bulletin builds on the work of WH by accounting for structural change.

We appeal to the economic theory found in WH and Wohlgenant (1989, 1996) to test for market power. However, the market data that we use to implement the tests appear to be driven by trends. In food markets, we hear about trends in technical change among firms (e.g., Clark and Youngblood), about trends in the industrial reorganization of industries (e.g., Martinez and Reed, MacDonald and Ollinger), and about trends in food consumption (Kinsey and Senauer). In this study, we define structural change as trends in market variables and account for different types of trends in tests of market power.

Unlike approaches that define structural change by changes in a model's parameters (e.g., Goodwin and Brester), our approach allows us to capture a key feature of some types of structural change: its changing trends are unpredictable. This view allows us to describe and to test for unpredictably changing trends in consumer and producer behavior. It allows us to test for market-clearing and to estimate longrun structural relationships despite such unpredictability. Finally, it allows us to test for market power in markets that may have undergone a series of permanent changes over time.

Our work sheds light on policy concerns associated with trends in the industrialization of food industries and in consumer behavior. Economic theory suggests such trends would be linked if markets clear (Engle and Granger). A *cointegrated* model links trends across variables and, in our case, provides a market-clearing representation. By failing to reject competition with a cointegrated model, our results suggest that trends in concentration and industrialization may be efficient solutions to unpredictable trends in consumer demand for food (Paul).

The Economic Model

Wohlgenant (1989) and WH introduced the agricultural economics profession to a market-clearing framework in which diverse firms demand farm and nonfarm inputs to produce the product mix of a composite industry. By relaxing the restriction of identical production functions across firms, they account for an industry's heterogeneous food items. They note that even if *each* firm produces its items in fixed-input proportions, because proportions vary across the diverse firms of a food industry, production of the entire industry is variable proportions.[1] The framework equips analysts with a tool for studying market relationships that is more general than the models derived from the traditional assumptions of fixed proportions and identical firms.

Market models of diverse firms generalize competitive market relationships. In particular, they can be used to reconcile the *concept* of competitive markets with the *observation* that: (1) increases in consumer food prices are not fully passed through to farmers (Wohlgenant, 1994); (2) nonfarm input prices and consumer food prices may move in opposite directions; (3) consumers may pay higher markups for higher priced products (George and King); and (4) competitive food industries may earn positive longrun rents (U.S. Department of Agriculture, September 1996). Market models based on fixed-proportions production must appeal to imperfect competition to explain these observations (Wohlgenant, 1999).

Tests of market power using market-level data, then, depend on assumptions concerning the nature of *industry* production. Retail-to-farm price spreads that exceed the marginal cost of transforming farm ingredients to final food products suggest market power, but the formulas used to compute spreads depend on the industry production function.[2] Studies based on fixed-proportions consistently reject the competitive model (e.g., Schroeter, Schroeter and Azzam, Azzam, Azzam and Park, and Koontz, Garcia, and Hudson), whereas WH and Wohlgenant (1989, 1996) just as consistently fail to reject the competitive model for U.S. food industries.[3]

This section presents an overview of the theory used in our empirical analyses. We refer readers unfamiliar with this theory to the cited studies of WH and Wohlgenant (1989, 1996) for a discussion that is more complete than that presented in this section. Readers familiar with the theory can skip to the next section.

At the core of the WH model are a pair of quasi-reduced-form retail and farm price equations for each market and a system of consumer-demand relationships linking the markets. Given a consumer demand schedule, the underlying structural model consists of two market-clearing conditions. The first states that the sum of food supply across the firms of the industry equals consumer demand for the industry output. The second states that the sum of farm ingredient demand across firms of the industry equals farm supply. The *critical feature of the WH model is that an industry's firms are not restricted to possessing identical production functions*. Within this general setup, WH assume that each industry faces an infinitely elastic supply of nonfarm inputs (exogenous nonfarm input prices), and a less-than-infinitely elastic supply of farm ingredients (endogenous farm prices). To simplify the model structure and isolate analysis on retail and farm prices, WH assume the food industry for a particular market consists of all the firms that manufacture, wholesale, and retail the industry's final food products.

[1] Wohlgenant (1999) formally shows that if one analyzes a competitive industry producing a heterogeneous mix of final consumer goods that we treat as a single composite (e.g., beef), the *observation* that retail-to-farm price spreads widen with increases in consumer food prices (e.g., George and King) implies input substitution. To explain this observation with a fixed-proportions based model for this heterogeneous industry, one must *rule out* the competitive model.

[2] Retail-to-farm price spread formulas used by ERS/USDA (Elitzak) are based on fixed-proportions production. Presumably, formulas based on variable-proportions would yield different magnitudes.

[3] These results are predicted by the theoretical results presented in Wohlgenant (1999).

Based on this structure[4] and on market clearing for farm ingredients and food output, Wohlgenant (1989, 1996) and WH derive the quasi-reduced-form

$$(1)\quad \begin{aligned} lnP_{rj} &= A_{rf}^{(j)}\, lnF_j + \boldsymbol{A_{rw}^{(j)}}\, \boldsymbol{lnW} + A_{rz}^{(j)}\, lnZ_j + e_{rj} \\ lnP_{fj} &= A_{ff}^{(j)}\, lnF_j + \boldsymbol{A_{fw}^{(j)}}\, \boldsymbol{lnW} + A_{fz}^{(j)}\, lnZ_j + e_{fj} \\ & \qquad j = 1,\ldots,J \end{aligned}$$

in which lnP_{rj} represents the natural logarithm (log) of the retail price in the *j*th market, lnP_{fj} is the log of the price of the farm ingredient used to produce output of the *j*th market, lnF_j is the log of the supply of the farm ingredient used to produce output of the *j*th market. In this study, lnF_j captures changes in domestic supply, and excludes changes in net exports and changes in private and government stocks of farm commodities. $\boldsymbol{lnW}$ is a vector of logged nonfarm input prices, and lnZ_j is a consumer demand shifter to be defined below. e_{rj} and e_{fj} are model errors on the retail and farm price equations.[5] These two retail and farm price equations are central to this bulletin.

In this framework, consumer demand defines the market, and the total consumer demand shift variable for the *j*th market, lnZ_j represents the effect of all variables that affect demand except the own-retail price for the product. For this bulletin, lnZ_j is derived as follows. Let

$$ln(Q_j/POP) = e_{jj}\, lnP_{rj} + e_{jy}\, ln\,(Y/POP) + \Sigma_{k\neq j}\, e_{jk}\, lnP_{rk} + u_j$$

be a per capita consumer demand relationship for the *j*th product in which $ln(Q_j/POP)$ is the log of per capita consumer demand for the output of the *j*th industry, lnP_{rj} is the log of the own-retail price, $ln(Y/POP)$ is the log of per capita disposable income, $lnP_{rk}\ (k\neq j)$ is the *k*th retail price of a gross substitute or complement to product category *j*, and u_j is an error term. Hence, the e_{jj} is the own-retail price elasticity of consumer demand, $e_{jk}\ (k\neq j)$ is a set of cross-price elasticities of demand, and e_{jy} is the income elasticity of demand for the *j*th good. Based on this relationship, the total demand shifter for the *j*th market is

$$(2)\quad lnZ_j = e_{jy}\, ln\,(Y/POP) + \Sigma_{k\neq j}\, e_{jk}\, lnP_{rk} + lnPOP$$

in which $lnPOP$ is the log of population. lnZ_j does not capture shifts in consumer demand caused by changes in the demand for food away from home, nor does it capture shifts caused by changes in the composition of the population.

The equations 1 are “quasi” reduced because they account for market-clearing in the *j*th market independent of market clearing in other markets.[6] Theory suggests four sets of expected signs on the quasi-reduced forms.

First, Heiner proves that for an industry of diverse firms, an increase (decrease) in the price of an input decreases (increases) an industry’s demand for the input. While this result is standard for an isolated firm and for an industry comprised of identical firms, Heiner’s proof applies to an industry comprised of firms with different longrun average costs. Heiner’s proof does *not* describe the negative slope of the sum of competitive firms’ input demand schedules holding output price constant. It describes instead the slope of industry input demand schedule as the sum of firms’ supply moves along a downward-sloping consumer demand schedule and, output price changes.[7] Braulke showed that Heiner’s proof applies to longrun equilibrium in which firms enter and exit the industry. In equations 1, $A_{ff}^{(j)}$ is the own-price flexibility of an

[4] A brief discussion of the relationship between the structural model and the quasi-reduced form is provided in the Appendix. The reader is referred to Wohlgenant (1989) or Wohlgenant and Haidacher for a more complete discussion.

[5] Constants and deterministic time trends are added to all of the empirical specifications below except the system of consumer demand relationships (Appendix). Only a constant term was added to the consumer demand system.

[6] The quasi-reduced-form equations derive from Heiner’s seminal work and the extensions of this work by Wohlgenant (1989) and WH. In the quasi-reduced-form representation, shifts in the market’s demand schedule are exogenous.

[7] For any single firm, Heiner found that the simultaneous change in the output price caused by the change in input price may trace out a positively sloped input demand for the firm. He found this positive relationship disappears when summing over all firms.

industry's demand for farm ingredients, and theory suggests $A_{ff} < 0$.[8]

Second, the industry's longrun quantity of food supply increases with its own-consumer food price. Heiner, Braulke, Panzar and Willig, and WH show that even if all input prices are exogenous to a competitive industry (flat input supplies), firm diversity implies that positive shifts in consumer demand trace an upward-sloping longrun industry supply function. Theory implies $A_{rz} > 0$.

Third, if firms are identical and farm ingredients are normal factors of production, a decrease in the supply of farm ingredients leads to a contraction of food supply and to an increase in consumer food prices. A *normal* factor of *industry* production is one in which the industry uses more of the factor to produce more output, while an inferior factor is one in which the industry uses less of the input to produce more output. The theory of diverse firms extends the neoclassical result that an increase in farm prices leads to increases in food prices only if farm ingredients are normal factors of industry food production. Since we expect that the aggregate farm ingredients specified here are normal, we expect $A_{rf} < 0$.

Fourth, if farm ingredients are normal factors of industry production and firms are diverse, positive shifts in consumer demand lead to longrun increases in farm prices. For that reason we expect $A_{fz} > 0$.

The theory of diverse firms does not unambiguously sign the response of consumer food prices to changes in nonfarm input prices. The reason is that a marketing input may be an *inferior* factor of production.[9] An increase in the price of an *inferior* factor raises a firm's average costs, but *reduces* its marginal costs. For a competitive industry comprised of identical firms, higher longrun average costs drive firms from the industry, reduce industry supply, and drive up consumer prices. The results may be different if an industry's firms are diverse.

Inframarginal firms are bestowed with firm-specific fixed assets that earn rent in the long run. Such firms are bestowed with firm-specific entrepreneurial capacity (Friedman) or location that provides them with a cost advantage over marginal firms (Panzar and Willig). One could argue that the entrepreneurial capacity of inframarginal pork-producing firms in the Southeast United States exceeds that of marginal producers in the Midwest. The cost advantage of inframarginal firms allows them to remain in the industry even as the longrun average cost of other firms is above market price. It follows that even in competitive markets, if the factor is inferior to inframarginal firms, an increase in its prices allows inframarginal firms to increase their supply even in the long run. The increase places downward pressure on output price. On the other hand, if the factor is inferior to marginal firms, their longrun average cost rises above output price. Marginal firms would exit the industry, thereby reducing industry supply and placing upward pressure on the market's average price of output. A negative sign on an element of $\boldsymbol{A_{rw}}$ suggests the associated factor is inferior to *industry* production, and that the positive supply response of inframarginal firms outweighs the negative response of marginal firms[10] (Panzar and Willig).

Theory provides a homogeneity condition. Since consumer demand is homogeneous of degree zero in retail (food) prices and income, and output supply and input demand are homogeneous of degree zero in farm and nonfarm input prices, the market-clearing price equations of (equations 1) are homogeneous of degree zero in farm and nonfarm input prices, retail prices, and income (e.g., WH, Wohlgenant [1989], Chavas and Cox).

The WH framework provides a test of the competitive model. The test is based on the notion that if a firm is

[8] Correspondingly, the retail price equation of (1) is a Heiner-type of industry-level output supply schedule.

[9] The example given here would not hold for the special case of only two inputs (e.g., one farm and one non-farm input). In this two-input case both factors must be normal, and increases in the price of either input raise the output price.

[10] Because firm-level production functions are not identical, a factor of production can be normal for some firms and inferior for others (e.g., older versus modern plants). Hence, a factor is normal (inferior) for an *industry* if the industry uses more (less) of the factor as it increases output. The weighted sums of individual firm-level elasticities determine whether a factor is normal or inferior (WH).

a price taker in both its purchase of inputs and its sales of output its profit function exists, and the symmetric second derivatives of its profit function define reciprocal relationships between a firm's output supply and input demands. Wohlgenant and WH derive an analogous symmetry condition for the group of diverse industry firms. Denoting $S_f^{(j)}$ as the cost share of farm ingredients for the *j*th industry, and to the coefficients in equations 1, WH show that symmetry at the *industry* level implies

$$A_{rf}^{(j)} = -S_f^{(j)} A_{fz}^{(j)}.$$

This condition states that if firms take farm and consumer prices as given, there exists a symmetric response of consumer and farm-level prices.

When studying retail and farm price relationships, analysts are often interested in the elasticity of transmission of farm prices to retail food prices. The *elasticity of price transmission* is the percent change in a retail food price induced by a 1-percent change in the farm price (George and King). Estimates of this elasticity reduce to the farm share if the food industry is competitive and if industry production exhibits constant returns with respect to farm ingredients. The assumption of fixed-proportions production (at the industry level) imposes constant returns with respect to *all* inputs, and therefore ensures transmission elasticities equal to the farm share. The WH model allows us to test whether the elasticity of price transmission equals the farm share within a variable-proportions framework.

Wohlgenant and WH show that in terms of the coefficients of equations 1, the *j*th industry's production displays constant returns with respect to the farm input if

$$A_{rz}^{(j)} = -A_{rf}^{(j)}$$
$$A_{fz}^{(j)} = -A_{ff}^{(j)}$$

hold. Constant returns for an industry imply zero industry profits in the long run. If both the symmetry and the constant returns restrictions hold, the elasticity of price transmission equals the farm share.

The model provides refutable hypotheses concerning oligopsony power. Policymakers often express concern that food producers exert market power when acquiring raw agricultural commodities from farmers. Some point out that captive supplies associated with new marketing arrangements may have changed the market structure so as to favor food producers and keep farm prices below competitive levels (U.S. Department of Agriculture, February 1996). Others counter that such voluntary arrangements may reflect the response to risk in a competitive market (Paul). The WH framework provides a test of the null that food producers acquire farm commodities competitively in national markets.

The test recognizes that if firms exert market power in acquiring farm commodities, a gap would exist between the farm price and the industry's demand for farm ingredients. Shifters on the farm supply function would define this gap.

At the level of the firm, the arguments are as follows. Let $P_{fj} = P_{fj}(F_j, \boldsymbol{S}_j)$ denote the inverse supply function for farm commodities facing the *j*th food industry, where $\boldsymbol{S_j}$ denotes a vector of shifters to this supply function. The first-order condition for profit maximization of a food-producing firm takes the form $MVP = P_{fj} + \lambda f\,(\partial P_{fj} / \partial F)$, where MVP is the marginal value product or firm-level demand for the farm commodity, and λ is a market power parameter. λ embodies the firm's conjecture about the effect its purchases of farm ingredients will have on the market (Bresnahan, p. 102-104). Note that the term $(\partial P_f / \partial F)$ in the above relationship is a function of $\boldsymbol{S_j}$. When $\lambda \neq 0$, the market level demand shifters, $\boldsymbol{S_j}$, enter the firm's optimization rule and define a gap between the market's farm price and the value of the marginal product for a competitive firm. Hence, when $\lambda \neq 0$, the marginal farm price – the firm's MVP – lies above the average farm price and firms restrict their demand for farm commodities. If $\lambda = 0$, price-taking firms recognize that their purchases impart no effect on the market, the farm price (or the value of the marginal product) equals the MVP as the industry level demand shifters ($\boldsymbol{S_j}$) do not enter firms' optimization conditions.

For a group of nonidentical firms of an industry, the arguments are similar. By eliminating F_j from equations 1, the two equations reduce to

$$(3)\quad P_{rj} = B_f^{(j)} P_{fj} + B_w^{(j)} W + B_z^{(j)} Z_j + v_r.$$

Equation 3 is an industry-level relationship similar to the first-order conditions of a price-taking firm. Under the null of no oligopsony power, the vector of supply shifters, $\boldsymbol{S_j}$, does not appear in equation 3. Under the alternative, S_j explains the gap and

(4) $P_{rj} = B_f^{(j)} P_{fj} + B_w^{(j)} W + B_z^{(j)} Z_j + \boldsymbol{B_s^{(j)\prime} S_j} + \boldsymbol{v_r}$

suggests the industry exerts oligopsony power in acquiring farm inputs. Equations 3 and 4 suggest that if industry j acquires farm commodities competitively, $\boldsymbol{B_s^{(j)\prime} = 0}$.

This concludes the review of the theory used to interpret the empirical results presented in the remainder of this report. Before we present these empirical results, however, we review the way in which structural change enters our empirical analyses.

Structural Change, Cointegration, and Market Clearing

In the previous section, the retail and farm price equations of 1 derive from market-clearing conditions for consumer food products and for farm ingredients (see Appendix). However, questions concerning the specification of equations 1 arise if markets have undergone structural change.

Regardless of the impact of structural change, market clearing means that excess supply (demand) would be of short duration and would equal zero on average. Furthermore, the effects of unforeseen shocks to market clearing would die out over time. In time series jargon, excess supply (demand) would be stationary. It is straightforward to show (using the market-clearing conditions shown in the Appendix) that the error terms, ε_{rj} and ε_{fj} , of equations 1 may represent excess supply (demand) variables for food and farm products. Stationary error terms imply market clearing, and market clearing would prevent the variables of equations 1 from moving too far apart.

Associated with structural change are market trends. Evidence of trends or changes in trends of market variables are often used as indicators of structural change. In time series jargon, variables that are characterized or generated by trends are non-stationary. However, even if each of the variables of equations 1 displays trends, the equations would still reflect market clearing if the excess supply variables or error terms are stationary. If each of the time series variables used in a model displays trends, and if model errors are stationary, the model is *cointegrated* (Engle and Granger). Tests of cointegration are tests of whether the data support the theory. In a cointegrated regression, some mechanism cancels or aligns the trends among the variables, and in equations 1 the mechanism is market clearing.

If the trends driving the non-stationary variables of equations 1 are not linked, excess supply would not die out, and the regression would be *spurious* (Granger and Newbold, p. 202-214, Hamilton p. 557-562). If the price equations of (1) are spurious, inter-temporal movements in one set of variables of equations 1 do not explain inter-temporal movements in other variables. A finding of a spurious regression would not support the theory.

In general, if market variables are driven by trends, the trends can be *deterministic*, stochastic, or a combination of both. A deterministic trend is a time trend defined in the usual way. Trends in demographics or predictable increases in real wages and productivity over the last century may drive the deterministic portion of trends in market variables. Market variables generated by deterministic trends pose few problems for statistical inference because with an infinite number of observations, such variables can be forecast from past observations with an arbitrary degree of accuracy. The second type of trend is a *stochastic* trend. Variables driven by stochastic trends are referred to as *unit root* or *integrated* series. For example, trends in real wages tied to unpredictable changes in the direction of inflation, unpredictable changes in the direction of consumer demand, technology, or the continual process of industrial reorganization, may be generating stochastic trends in market variables.[11] Unlike deterministic trends, stochastic trends change direction unpredictably. Integrated market variables pose special problems for statistical inference because even in infinite samples, optimal forecasts of these variables do not converge, but are continually revised as new observations become available.

More formally, the accuracy and reliability of forecasts of market variables depend on whether the variable is driven by a deterministic or a stochastic trend. As the forecast horizon grows, the forecast of a series generated by a deterministic trend converges to a time trend, and the mean squared error (MSE) of this forecast converges to the unconditional variance of the series (Hamilton, p. 438-42). Population, real wages, and real disposable income may be accurately and reliably forecast. On the other hand, tastes and preferences, technology, and the continual reorganization of an industry may be stochastic trends because changes in any of these may be impossible to predict. Unlike deterministically trending variables, the forecast of a unit root variable diverges with the length of the forecast horizon, and the MSE of the forecast increases without bound (Hamilton, p. 438-42).

Associated with each type of trend is a type of cointegration. A model constructed from deterministically trending variable series is *deterministically cointegrated* if the deterministic trends in the model's variables are linked. In practice, a regression model is

[11] That is, a series that is stationary around a deterministic trend.

deterministically cointegrated if a time trend variable appended to the model is not statistically different from zero. A model constructed from a set of stochastically trending series is *stochastically cointegrated* if the model errors are stationary. Just as market variables may reflect both a deterministic and a stochastic trend, a model may be both deterministically and stochastically cointegrated.

Using annual time series from 1958-97, we computed Dickey-Fuller and Phillips-Perron t-tests for the logged and deflated variables used in the seven sets of retail and farm price equations.[12] Both sets of tests are designed to refute the hypothesis that, conditioned on an AR(1) representation, a single unit root net of an intercept (or drift) or net of a deterministic time trend governs the series. The tests differ in the way they handle serial correlation of the error terms of the AR(1) specification.[13] Almost without exception, the two sets of tests suggest that both a stochastic and a deterministic trend drive most of the variable series used in equations 1.[14]

Given evidence of trends in the variables, the question is whether these equations are stochastically and deterministically cointegrated. The specification of equations 1 used throughout this report is as follows. The deterministic regressors include an intercept and a deterministic time trend. The stochastic regressors include a vector of (logged) nonfarm input prices (***ln W***), a (log) farm supply variable specific to market i ($ln\ Q_i$), and the total demand shifters ($ln\ Z_i$) for each of the consumer demand equations.[15] The vector ***ln W*** consists of (logs of) wages, the price of packaging, the price of transportation, and the price of energy. To satisfy homogeneity, all prices and income variables in equations 1 and 2 are deflated by the price of other nonfarm inputs (Elitzak). Hence, the tests for cointegration are based on a specification that includes six stochastic regressors, a constant, and a deterministic time trend for each retail and farm price equation. We compute two sets of tests.[16]

The first is based on model residuals, and specifically tests the null hypothesis of a spurious regression. Again, a model is spurious (or not *stochastically* cointegrated) if the model errors follow a unit process. Engle and Granger (1987) suggest applying Dickey-Fuller tests to model residuals. Phillips and Ouliaris confirm that the Dickey-Fuller and Phillips-Perron statistics can be used to test for spurious regressions. They find, however, that the critical values depend not on the number of observations, but on the number of stochastic regressors used in model specification, and whether the regression includes an intercept or a deterministic time trend.

Table 1 reports Dickey-Fuller and Phillips-Perron t-tests designed to refute the null that the equations 1 are spurious regressions. The statistics are based on Ordinary Least Squares (OLS) residuals. The Dickey-Fuller results presented in table 1 fail to reject the null of a spurious regression for each of the 14 equations, while the Phillips-Perron results fail to reject the null for 8 of the 14 equations.

[12] The sample series used to create the variables are discussed in the Appendix. All price and income variables are deflated by the price of other nonfarm inputs and logged prior to testing. The total demand shifters are computed directly from the estimates of the double-log level system of consumer demand equations. Prices and income are also deflated by the price of other nonfarm inputs prior to estimation. The test results were computed using Shazam.

[13] Dickey and Fuller control for the serial correlation in the error terms by adding lags to the AR(1) representation. Phillips and Perron compute estimates of the covariogram of the errors of an AR(1) process. For a concise comparison between the tests we refer the reader to Hamilton (p. 504-518). The simulations presented in Phillips and Perron (1988) reveal that neither test is universally more powerful than the other.

[14] The Dickey-Fuller and Phillips-Perron results are available upon request. Some results conflict. For example, the results suggest the farm prices for poultry and eggs may be stationary around a time trend (-3.95 and -4.49) but are unit root non-stationary around an intercept. The results also suggest the farm supply for beef may be stationary around a constant but non-stationary around a time trend. When unit root tests conflict, Holden and Perman spell out a multi-step procedure that may be useful in sorting out the results.

[15] The seven-equation demand system was also found to be stochastically cointegrated, and seemingly unrelated canonical cointegrating regression estimates with an intercept and no time trend and symmetry and homogeneity imposed are used to construct the demand shift variables used in equations 1.

[16] Because we found that a linear deterministic time trend variable was invariably statistically different from zero, we reject the null of deterministic cointegration for each price equation and included it in the model specifications for each industry. Hence, our tests of cointegration are tests specifically for stochastic cointegration.

The second set of tests is designed to examine the null that the regressions are stochastically cointegrated. The tests are based on the observation by Park (1990) that appending a set of integrated series to a stochastically cointegrated model yields a spurious regression. If the additional variables add no explanatory power to the regression, they are superfluous, and the original model specification is cointegrated. The technical problem of testing whether the variables are superfluous is that, in general, the model error terms are correlated with the first differences of the model's regressors. This correlation destroys the asymptotic normality of parameter estimates, and hence destroys the reliability of the usual chi-square tests. Park (1990) derives a transformation that accounts for this correlation, and uses it to transform each of the variables of a model. Chi-square tests based on the transformed regression represent valid tests of the null that the additional variables are superfluous, and the original model is stochastically cointegrated.

Table 1 — Residual-based tests of spurious regressions

	Dickey-Fuller	Phillips-Perron
Retail price equations		
Beef and veal	-3.61 (1)	-5.73** (1)
Pork	-3.18 (3)	-5.74** (3)
Poultry	-4.65 (0)	-4.44 (1)
Eggs	-3.94 (4)	-3.80 (1)
Dairy	-4.84 (1)	-7.52** (1)
Fresh fruit	-4.20 (0)	-4.12 (1)
Fresh vegetables	-4.22 (2)	-4.55 (2)
Farm prices (ln P_f)		
Beef and veal	-3.18 (0)	-3.15 (1)
Pork	-3.41 (0)	-3.49 (1)
Poultry	-3.41 (1)	-5.21* 1)
Eggs	-3.15 (1)	-4.99 (1)
Dairy	-3.72 (1)	-6.00** (1)
Fresh fruit	-4.02 (0)	-3.92 (1)
Fresh vegetables	-3.82 (1)	-5.23* (1)

Values are t-tests associated with the coefficient of the lagged OLS residuals in which the regression includes a constant and a time trend. Values in parentheses indicate the number of lagged-first differences in the autoregression (Dickey-Fuller) or the number of lags included in the error covariogram (Phillips-Perron).

*Reject the null of a spurious regression at (approximately) the 0.10 level. The result is based on a critical value of approximately -5.2. which is -0.5 plus -4.7. -4.7 is the critical value computed by Phillips and Ouliaris for a demeaned (constant) and detrended (one deterministic time trend) regression with five stochastic regressors (Phillips and Ouliaris, Table IIc). -0.5 is the increment associated with adding the sixth stochastic regressor (Ng).

**Reject the null of a spurious regression at (approximately) the 0.05 level. The result is based on a critical value of approximately -5.5 which is -0.5 plus -5.0. -5.0 is the critical value associated with a demeaned and detrended regression with five stochastic regressors and -0.5 is the increment associated with the sixth regressor (Ng).

To employ the test, however, one is faced with choosing a set of integrated and potentially superfluous variables to append to the model. Here, theory provides no clear guide. Hence, we follow Park's suggestion by appending higher order polynomials of deterministic time trends to the retail and farm price equations.

Table 2 reports chi-square estimates associated with Park's J_1 test (1990) of the null that the coefficients of the polynomial time trend terms (T^2, T^3 & T^4) are jointly zero (see table 2 for the specification of the polynomial terms). The p-values (in parentheses) suggest that at the 0.05 level of rejection, all equations are stochastically cointegrated except the retail price equation for poultry and the farm price equations for beef, poultry, and dairy.

The results in tables 1 and 2 provide somewhat mixed results. The Dickey-Fuller tests (table 2) suggest each of the 14 equations is spurious. The Phillips-Perron tests suggest that, at reasonable levels of rejection, 8 of the 14 price equations are cointegrated. Finally, Park's J_1 test suggests that, at reasonable levels of rejection, 10 of the 14 equations are cointegrated. In combination, the Phillips-Perron and Park's test results suggest that only the retail price equation for poultry and the farm price equation for beef and veal may be spurious.

Table 2 — Variable-addition tests of cointegration

Industry	Retail price equations	Farm price equations
Beef and veal	6.83 (.08)	9.52 (.02)
Pork	3.75 (.29)	0.59 (.90)
Poultry	18.23 (.00)	27.27 (.00)
Eggs	4.81 (.19)	3.16 (.37)
Dairy	7.72 (.05)	16.90 (.00)
Fresh fruit	1.28 (.73)	6.83 (.08)
Fresh vegetables	0.31 (.96)	2.92 (.40)

Entries are c2 with the restriction $a1=a2=a3=0$ in the general regression $P=\mathbf{X}\boldsymbol{\beta} + \mathbf{f(a,T)} + e$ in which $\mathbf{f(a,T)} = {}_1T^2 + a_2T^3 + a_3T^4$ and in which T = t/max(t) where P is either the logged and deflated retail or farm price, X includes the intercept, the linear time trend, and the six logged and deflated stochastic regressors. The values in parentheses are p-values, and represent the size of the rejection region associated with the restriction. For example, values greater than 0.05 fail to reject the null of stochastic cointegration at the 0.05 level of rejection.

Further Empirical Results

In the previous section, we argued that trends in market data reflect structural change, and found evidence of both deterministic and stochastic trends in key variables associated with seven U.S. markets. In addition, we found evidence that markets distribute these trends across consumers and food producers. Despite the claim that food markets may have undergone a sequence of permanent changes over time, we show in this section that market-clearing provides mostly stable longrun retail and farm price relationships.

The stochastic trends embedded in the variables of equations 1 require us to deviate from textbook estimation procedures. As stated above, such procedures fail to account for a non-zero correlation (at any lag) between the first difference of an explanatory variable (i.e., the fundamental error terms of the variables) and a cointegrated model's stationary error terms. This correlation is present in all but the simplest class of cointegrated models. While it does not destroy the consistency of parameter estimates, the correlation does destroy the asymptotic normality of the estimates and renders textbook formulas for the χ^2, F, and t tests invalid for inference.

Park (1990, 1992) and Park and Ogaki transform variables of cointegrated regressions based on this correlation. Their transformations reduce general, cointegrated regressions to the simple (or canonical) class of cointegrated regressions in which first differences of explanatory variables are *not* correlated with regression errors. The procedure is to first transform the variables of a cointegrated regression and then to apply textbook procedures to the transformed regressions. The canonical cointegrating regression (CCR) estimator applies OLS to a transformed single equation (Park 1990). We used the CCR estimator in the previous section to compute Park's variable addition test of the null of cointegration (Park 1992). In this section we again use the CCR estimator to compute a variable addition test of oligopsony power. In addition, we apply the seemingly unrelated regression (SUR) estimator to a transformed, seven-equation consumer demand system, and to the cointegrated quasi-reduced-form retail and farm price equations (i.e., equations 1) for each industry. This two-step estimator, termed the seemingly unrelated canonical cointegrating regression estimator (SUCCR), provides us with unbiased estimates of market structure and asymptotically-correct inference on tests of market power and constant returns in multiple equation systems (Park and Ogaki). We refer interested readers to Park, and to Park and Ogaki for details on the transformations that we use to compute the estimates presented in this section.

Consumer Demand

Kinsey and Senauer argue that changing trends in consumer behavior lead to a changing structure of the food sector. Cointegrated, market-clearing relationships would reflect the transmission of trends from consumers. To capture trends in consumer demand, we specify and estimate a consumer demand system for the seven industries. In this section, we discuss the specification of the seven-equation consumer demand system.

Appendix table 1 presents point estimates and t-values of the seven-equation system of composite per-capita demand. To construct the empirical consumer demand model, we used logged data on per-capita consumer disappearance as proxies for the seven dependent per capita consumption variables, and deflated all prices and income (explanatory variables) by the price of other nonfarm inputs (to ensure homogeneity of the market-clearing conditions). Based on Dickey-Fuller and Phillips-Perron tests, we could not refute the null that virtually all of the logged variables of the consumer demand system are unit root non-stationary around a deterministic trend. In addition, we found evidence that the individual consumer demand equations are cointegrated. Next, we imposed the symmetry and homogeneity restrictions (e.g., Deaton and Muellbauer, p. 43-46; Silberberg, p. 250-253) at the mean of the sample, and report the restricted point estimates in appendix table 1. The restricted estimates were then used to construct the demand shifters, $ln\ Z_j$, for each industry j (equation 2), and to compute full reduced-form price responses reported below.[17]

[17] We are aware of the problem with incorporating the adding-up condition on this double-log specification (Deaton and Muellbauer, p. 17), and are aware of the conceptual problem of using farm-level disappearance data as the dependent variable of the system (WH). Our purpose here is to compute only approximate values of the shifters on consumer demand.

Tests of Competition and Constant Returns

Table 3 reports the χ^2 and p-values associated with symmetry, constant returns or zero profits for the industry, and the joint restrictions of symmetry and constant returns for the seven industries. Failure to refute symmetry suggests food firms take both output and farm ingredient prices as given. Failure to reject constant returns for the *industry* suggests that free entry and exit of diverse firms result in zero longrun profits. Failure to refute the joint hypotheses of symmetry and constant returns suggests that, in the long run, a 1-percent increase in the price of a farm commodity results in an increase in the price of a composite food category by a percentage equal to the cost share of the farm commodity used in producing the food category.

The symmetry (only) and constant returns (only) test results provide evidence of longrun competition. In particular, the symmetry test fails to refute (at the 0.05 level) the longrun competitive model for the beef, dairy, eggs, fresh fruit, and fresh vegetable industries. The constant returns test fails to refute the longrun competitive model for poultry, fresh fruit, and fresh vegetables. It is worth repeating that this general finding of competitive markets takes into account the many permanent changes that may have occurred in these markets over time.

Table 3 — Tests of symmetry and constant returns

	Symmetry only	Constant returns only	Symmetry and constant returns
Beef and veal	0.4773	87.9159	97.2574
	(.490)	(0.00)	(.000)
Pork	94.9740	34.3050	98.0828
	(.000)	(0.00)	(.000)
Poultry	26.1533	4.0383	39.4573
	(3E-7)	(.133)	(1E-8)
Eggs	0.0381	223.52	255.7759
	(.845)	(0.00)	(0.00)
Dairy	0.9081	48.6256	49.1519
	(.341)	(0.00)	(.000)
Fresh fruit	2.1428	4.7536	4.7871
	(.143)	(.093)	(.188)
Fresh vegetables	1.7731	0.8505	6.0862
	(.183)	(.654)	(.107)

Values are chi-square statistics. Values in parentheses are p-values, or the size of the rejection region necessary to reject the null hypothesis.

Furthermore, our tests reject the joint restriction of symmetry and constant returns for all industries except fresh fruit and fresh vegetables. The results suggest that estimates of elasticities of farm price transmission to retail apply only to markets in which final products undergo a minimal amount of food processing.

The general finding of competitive markets is consistent with WH and Wohlgenant's (1989, 1994, 1996) findings. On the one hand, we expect our findings to be similar because the model structures and data are very much the same.[18] On the other hand we expect differences because of the different estimation procedures. While our approach exploits deterministic and stochastic trends in market data, the cited works remove stochastic trends through a first-difference transformation prior to estimation. In comparing the procedures using the same retail and farm price equations, Reed and Clark find that one fails to reject parametric restrictions more often using a first-difference specification of an econometric model.[19] The reason is if the explanatory variables are integrated, a first-difference transformation removes the dominant, longrun component of the variance of the variables. Hence, if the variables are integrated, a first-difference filter would inflate the variance of parameter estimates and could reduce the likelihood of rejecting any parametric restrictions. It is noteworthy that we reject both the symmetry and the joint restriction of symmetry and constant returns more often than the cited works.

Oligopsony Power

In a previous section, we reviewed the theory used to test for oligopsony power. If food firms exert oligopsony power in acquiring farm ingredients, a gap would exist between the farm price and the value of the marginal product of farm ingredients at the market level. Shifters on the farm supply associated with the jth market, $\boldsymbol{S_j}$, would explain this gap. Recall from above that under the null hypothesis of food producers taking

[18] We are aware that the model specifications for eight industries in Wohlgenant (1989, 1994) include only a single nonfarm input price. The model specification for beef and pork (only) used in his 1996 paper is similar to the specification used here, as it includes the same four nonfarm input prices. Furthermore, our work uses a different deflator to impose homogeneity.

[19] The study controls for differences in the data and model specifications.

farm prices as given, no gap exists and the retail-farm price relationship is

$$(3') \quad P_{rj} = B_f^{(j)} P_{fj} + B_w^{(j)} \boldsymbol{W} + B_z^{(j)} Z_j + v_r .$$

Under the alternative of oligopsony power, the retail-farm price relationship is

$$(4') \quad P_{rj} = B_f^{(j)} P_{fj} + B_w^{(j)} \boldsymbol{W} + B_z^{(j)} Z_j + \boldsymbol{B_s^{(j)\prime}} \boldsymbol{S_j} + v_r .$$

A test of whether firms acquire farm commodities competitively in national markets reduces to a test of the restriction $\boldsymbol{B_s^{(j)\prime}=0}$.

The usual chi-square tests of statistical significance of the $\boldsymbol{S_j}$ variables would be reliable only if the variables of equation 3′ and the $\boldsymbol{S_j}$ would be stationary. Because we found evidence that both sets of variables are integrated, we proceed as follows. Under the null hypothesis of price-taking, equation 3′ is cointegrated and its error terms are stationary, and the variables of equation 3′ are transformed to account for the correlation between first differences of explanatory variables and the model error terms (Park 1990, 1992). Under the null, the integrated (untransformed) $\boldsymbol{S_j}$ variables would be independent of the stationary error terms of equation 3′, and $\boldsymbol{B_s^{(j)\prime}=0}$. Under the alternative of oligopsony power, the error terms of equation 3′ would be integrated, and this price-taking relationship would be spurious. In this case, the integrated $\boldsymbol{S_j}$ variables and the integrated error terms of equation 3 would not be independent and in general $\boldsymbol{B_s^{(j)\prime} \neq 0}$. Rather than testing for the statistical significance of $\boldsymbol{S_j}$, chi-square tests of $\boldsymbol{B_s^{(j)\prime}=0}$ represent a test of whether the price taking model is cointegrated or correctly specified against the alternative that the oligopsony power relationship is correctly specified (Park 1992).

Table 4 reports the chi-square and p-values computed from the transformed regressions. The results report the statistics associated with one integrated, industry-specific farm supply shifter (S_1) and both integrated, industry-specific farm supply shifters (S_1 & S_2).[20] The results are based on a specification that includes a constant and a deterministic time trend. At reasonable levels of rejection, we fail to reject the null that in national markets, the seven food industries acquire farm ingredients competitively.

Table 4 — Tests of competition in acquiring farm commodities

	S_1	S_1 & S_2
Beef and veal	1.2377 (.266)	1.3832 (.501)
Pork	0.0443 (.833)	2.4686 (.291)
Poultry	0.0538 (.816)	0.2956 (.862)
Eggs**	0.1994 (.655)	
Dairy	0.1294 (.719)	0.2022 (.904)
Fresh fruit	2.5509 (.110)	
Fresh vegetables	0.2176 (.641)	

Entries are χ^2 values and values in parentheses are significance levels. The sets of supply shifters on farm supply are as follows (see Appendix for data series definitions): Beef: S_1 is steers, S_1 & S_2 are steers and corn price; Pork: S_1 is hog inventories, S_1 & S_2 are hog inventories and corn price; Poultry: S_1 is the price of soybean meal, S_1 & S_2 are the price of soybean meal and corn price; Eggs: S_1 is laying flock; Dairy: S_1 is cow numbers and S_1 & S_2 are cow numbers and price of soybean meal; Fresh fruit: S_1 is farm wages; Fresh vegetables: S_1 is farm wages.

**Sample interval is 1960-97.

Longrun Industry Structure

Table 5 presents SUCCR estimates of the parameters of equations 1 for the seven food markets. Although they account for a negative own-price consumer response, they are conditioned on shifters of consumer demand (i.e., *ln Z*). Hence, the estimates would not account for particular shifts in consumer demand induced by endogenous changes in relative retail prices among the seven composite markets. However, by controlling for such shifts, the ‘quasi’ reduced-form estimates of equations 1 provide information on industry structure. Given evidence of permanent change and cointegration presented above, the results in table 5 represent longrun estimates of industry structure.

Theory predicts a negatively sloped, longrun industry demand for farm ingredients. In terms of equations 1,

[20] The industry-specific shift variables on farm supply are defined in the appendix. For the test to be meaningful, the variables must be integrated. We could not refute the claim of integrated farm supply shifters. Furthermore, because the null hypothesis is that equation 3′ (and not equation 4′) is cointegrated, the $\boldsymbol{S_j}$ variables are not transformed.

Table 5 — Quasi-reduced-form estimates

Variable	Beef and veal*		Pork**		Poultry**		Eggs*		Dairy*		Fresh fruit*		Fresh vegetables*	
	Retail price	Farm price	Retail price	Farm price	Retail price	Farm price	Retail price	Farm price	Retail price	Farm price	Retail price	Farm price	Retail price	Farm price
Demand shifter (Z)	1.011	2.537	1.075	1.020	0.808	0.706	4.285	1.045	1.132	1.055	0.062	0.447	0.039	1.096
	(11.1)	(18.5)	(13.6)	(4.53)	(7.72)	(7.13)	(17.5)	(7.95)	(15.7)	(8.15)	(.32)	(1.71)	(.21)	(5.87)
Wage (W_1)	1.372	1.377	0.013	2.236	-0.461	1.229	0.079	2.037	0.172	1.792	-0.930	1.577	-0.115	1.746
	(15.6)	(5.38)	(.197)	(10.6)	(-2.99)	(4.85)	(.36)	(7.91)	(3.26)	(14.9)	(-5.1)	(3.92)	(-.90)	(6.86)
Packaging price (W_2)	0.446	0.354	0.115	-0.013	0.000	0.005	1.395	-0.080	0.010	-0.010	0.491	-0.108	0.308	0.235
	(6.29)	(1.69)	(1.63)	(-.06)	(.001)	(.017)	(5.48)	(-.29)	(.21)	(-.09)	(2.64)	(-.34)	(2.24)	(1.08)
Transportation price (W_3)	-0.608	-0.534	-0.123	-0.795	0.441	-0.261	-0.966	-1.362	-0.235	-0.468	0.224	-0.015	0.191	-0.251
	(-7.77)	(-2.27)	(-2.05)	(-3.84)	(2.90)	(-1.01)	(-4.03)	(-5.03)	(-4.82)	(-3.91)	(1.36)	(.04)	(1.52)	(-.98)
Energy price (W_4)	-0.157	-0.170	0.009	-0.100	-0.132	-0.101	-0.021	0.146	0.037	-0.152	-0.131	-0.224	-0.282	-0.255
	(-4.09)	(-1.48)	(.25)	(-.84)	(-1.47)	(-.69)	(-.15)	(.93)	(1.43)	(-2.55)	(-1.36)	(-1.31)	(-3.97)	(-2.09)
Farm supply (F_i)	-1.446	-1.970	-1.187	-2.107	-1.057	-0.897	-0.658	-2.125	-0.517	-1.593	-0.148	-0.984	-0.373	-1.342
	(-18.5)	(-8.72)	(-22.4)	(-17.2)	(-8.17)	(-9.69)	(-7.95)	(-21.9)	(-8.15)	(-15.3)	(-1.71)	(-3.72)	(-5.87)	(-7.96)
Constant	3.690	-0.096	0.257	8.223	7.300	2.551	-20.356	8.317	-0.999	4.941	0.593	4.318	2.924	3.210
	(9.83)	(-.11)	(.64)	(8.49)	(6.63)	(3.38)	(-13.5)	(14.4)	(-1.21)	(8.49)	(-.50)	(2.07)	(3.22)	(2.55)
Time trend	-0.004	-0.025	-0.000	-0.018	0.006	-0.020	-0.016	-0.023	-0.005	-0.018	0.021	-0.017	0.021	-0.018
	(-4.48)	(-10.1)	(-0.33)	(-7.27)	(1.01)	(-6.59)	(-5.28)	(-6.96)	(-6.70)	(-14.2)	(-4.33)	(-4.34)	(4.79)	(-7.01)

* Symmetry imposed.
**Unrestricted estimates.

theory predicts $A_{ff} < 0$. The negative estimates of A_{ff} for each of the seven markets are statistically different from zero. The estimates describe downward sloping, industry-level demand schedules for farm ingredients.

The theory of diverse firms in a competitive market predicts that positive shifts in the consumer demand function trace an upward sloping, longrun industry supply schedule. In terms of equations 1, theory predicts $A_{rz} > 0$. The estimates of A_{rz} are positive for all seven markets, and except for fresh fruit and fresh vegetables, they are statistically different from zero at reasonable levels of rejection.

The theory of competitive markets predicts that if farm ingredients are normal factors of production, a contraction in farm supply raises consumer food prices. In terms of equations 1, theory predicts $A_{rf} < 0$. The estimates of A_{rf} are negative for each of the seven industries, and are statistically different from zero. Theory also predicts that if farm ingredients are normal, $A_{fz} > 0$. The estimates of A_{fz} are positive for all seven markets and are statistically different from zero. Negative estimates of A_{rf} and positive estimates of A_{fz} suggest the aggregate farm ingredients are normal factors of industry production.

The estimates presented in table 5 suggest some marketing factors are inferior to a number of industries. Negative signs on elements of $\boldsymbol{A_{rw}}$ suggest the particular factor is inferior and that the supply response of inframarginal firms exceeds that of marginal firms. For example, the results suggest transportation is an inferior factor for the beef and pork industries. The estimates may indicate that for the U.S. pork industry, changes in vertical coordination have allowed the inframarginal firms in the Southeast United States to economize on the transportation of hogs. The estimates also suggest that labor is an inferior factor for the fresh fruit and fresh vegetable industries.

The results presented in table 5 also point to some nonfarm inputs that appear to be normal across industries. The positive signs on the A_{rw} coefficients associated with the price of packaging suggest that packaging is a normal factor for all seven industries. This may reflect the notion that consumers value the convenience associated with the packaging of food products, and suggests that consumers would be willing to pay more for packaging through higher food prices. Furthermore, labor appears to be a normal factor of production for four of the seven industries.

Input Substitution

The variety of consumer products and diversity of firms within a composite industry (e.g., fresh fruits) provide evidence of variable-proportions at the market level (e.g., WH, Wohlgenant [1999]). Refutation of restrictions implied by fixed-proportions analyses provides additional evidence.

The diversity of a composite industry's products suggests that production processes vary across firms. Meat products, for example, vary by the amount of processing. Manufacturers of processed meat products would, for example, utilize higher proportions of nonfarm ingredients (e.g., packaging, energy) than manufacturers of fresh meat products. An increase in the price of nonfarm inputs relative to farm inputs would, therefore, raise longrun average and marginal costs for manufacturers of processed products more than for manufacturers of fresh products.[21] In terms of meat industry supply, the quantity of manufactured products would fall relative to fresh products. In terms of input demand, the reduction in the supply of manufactured products relative to fresh products means that the ratio of nonfarm inputs to farm inputs demanded by the industry falls in response to an increase in the relative price of nonfarm inputs.

WH (p. 21) formally show that if firms are diverse, an industry's input price response can be decomposed into a substitution and an output effect in precisely the same manner as one could decompose the response of a single firm with a variable proportions production technology. Wohlgenant (1999) illustrates the presence of input substitution directly for a composite industry producing heterogeneous final food products. All that is required is that production functions differ across firms. These results imply that even if each firm in an industry produces its specific product in fixed-input proportions, if these proportions vary across firms, variable-proportions relationships apply at the market level.

Test results presented in table 6 suggest that market data do not follow the predictions of fixed proportions. In particular, fixed proportions at the market level imply that the own-price elasticity of an industry's

[21] The marginal costs could fall if nonfarm ingredients are inferior factors of industry production. The discussion here assumes they are normal factors.

Table 6 — Tests of fixed-proportions production

Industry	t-values
Beef and veal	110.744
Pork	15.063
Poultry	25.204
Eggs	233.344
Dairy	4.806
Fresh fruit	51.317
Fresh vegetables	-331.354

Values are Student t-values designed to test the restriction that the parameter A_{ff} equals the inverse of the demand for farm ingredients implied by fixed proportions. Values approximately equal to 2 (in magnitude) would reject the null of fixed proportions at the 0.05 level. The estimates of A_{ff} are found in table 5, and the estimates of the standard errors are the estimates used to compute the t-values in table 5. Point estimates of the own-price elasticities of consumer demand and the farm share for each market are found in appendix tables 1 and 2.

demand for farm ingredients equals the product of the own-price elasticity of consumer demand and the industry's cost share of farm ingredients (e.g., George and King). Since $A_{ff}^{(j)}$ is the inverse of the industry j's demand elasticity with respect to the jth farm price (from equations 1), e_{jj} is the own-price elasticity of consumer demand for the jth consumer product, and $S_f^{(j)}$ is the industry's cost share of farm ingredients, the restriction $A_{ff}^{(j)} = 1/(S_f^{(j)} e_{jj})$ would hold if an industry produced its composite mix of products in fixed-factor proportions. Based on the t-values reported in table 6, one can refute the null hypothesis of fixed proportions at the industry level for reasonable levels of rejection.[22]

Estimates of elasticities of input substitution can be computed from model parameters (Wohlgenant 1996) and used to measure the ease with which an industry varies its factor proportions. Table 7 reports estimates of Morishima elasticities of substitution (Blackorby and Russell) between the farm and the four marketing inputs when (in this case) changes in the farm-to-nonfarm price ratio are caused specifically by a

[22] The results are similar to those reported by Wohlgenant (1996). The tests are preliminary because they treat the estimate of the own-price elasticity of consumer demand as a parameter with no variation. A simulation procedure suggested by Ng, or a bootstrap procedure suggested by Li and Maddala may provide a more precise test.

Table 7 — Morishima elasticities of substitution

Industry	Nonfarm inputs			
	Labor	Packaging	Transport	Energy
Beef and veal	2.613	1.910	-3.084	-1.775
Pork	1.717	0.456	-1.470	-0.013
Poultry	4.356	1.146	-1.912	-1.220
Eggs	4.181	0.092	-10.45	2.807
Dairy	3.083	0.590	-2.193	-1.210
Fresh fruit	2.811	0.697	0.939	-1.227
Fresh vegetables	2.270	1.275	-0.218	-1.212

change in the farm price.[23] The larger the magnitude of the estimate, the easier it is for an industry to vary ratios of farm and marketing inputs. Positive estimates suggest the input pairs are substitutes (when the farm price changes), and negative estimates suggest they are complements. For example, the results in table 7 suggest that as farm prices rise, labor and packaging substitute for the farm ingredient in all seven industries. As in Wohlgenant (1996), the estimates suggest that significant substitution possibilities exist in U.S. food production.

Full-Reduced-Form Price Responses

Table 8 reports full-reduced-form estimates of percent changes in retail and farm prices induced by a 1-percent increase in the set of explanatory variables. Unlike the quasi-reduced-form estimates (table 5), the full-reduced-form estimates account for the effect of endogenous shifts in consumer demand when relative retail prices change. For example, increased wages may increase the retail price of both beef and pork products, but the magnitude of the responses would differ in the markets. The results presented in table 8 capture the effect of consumer responses to changes in relative retail prices on retail and farm prices.[24]

[23] Since Morishima elasticities of substitution are not symmetric, an estimate of response caused by a 1-percent change in the factor price ratio induced by a change in the nonfarm input price would differ from the estimates reported in table 7.

[24] The full-reduced-form estimates are computed using equations 19a and 19b (Wohlgenant 1991) or equations 30a and 30b (WH). Whereas theory provides expected signs on the coefficients of the quasi-reduced form, it does not provide expected signs on the parameter estimates of the full-reduced form. Technically, the reason is that unlike a single consumer demand equation, a system of consumer demand equations is not negative definite (Chavas and Cox). This is essentially why the quasi-reduced form can provide information on industry structure and the full-reduced form cannot.

Table 8 — Full-reduced-form estimates

Explanatory variable	Beef and veal Retail price	Beef and veal Farm price	Beef and veal Spread	Pork Retail price	Pork Farm price	Pork Spread	Poultry Retail price	Poultry Farm price	Poultry Spread
				Percent change					
Nonfood price	0.707	1.773	-1.066	0.749	0.711	0.038	-0.765	-0.669	-0.096
Beverage price	-0.138	-0.347	0.209	0.098	0.093	0.005	0.149	0.130	0.019
Sugar price	0.379	0.952	-0.572	-0.164	-0.155	-0.008	-0.788	-0.689	-0.099
Cereal price	-0.861	-2.160	1.299	-0.177	-0.168	-0.009	1.073	0.938	0.135
Income/capita	-0.043	-0.107	0.065	-0.221	-0.209	-0.011	0.604	0.528	0.076
Population	2.294	5.755	-3.461	2.736	2.596	0.140	0.970	0.848	0.122
Wage	2.292	3.684	-1.392	1.586	3.729	-2.143	-1.220	0.566	-1.786
Package price	0.592	0.721	-0.129	0.483	0.336	0.147	0.134	0.122	0.012
Transport price	-1.150	-1.893	0.743	-0.940	-1.570	0.631	0.687	-0.046	0.733
Energy price	-0.117	-0.069	-0.048	-0.039	-0.146	0.107	-0.179	-0.141	-0.038
Cattle supply	-2.037	-3.452	1.416	-1.329	-1.261	-0.068	0.477	0.417	0.060
Hog supply	-0. 577	-1.448	0.871	-1.577	-2.477	0.900	0.186	0.163	0.023
Poultry supply	0.096	0.240	-0.144	0.079	0.075	0.004	-1.159	-0.986	-0.173
Egg supply	-0.046	-0.116	0.070	-0.040	-0.038	-0.002	-0.041	-0.036	-0.005
Milk supply	-0.065	-0.163	0.098	-0.091	-0.086	-0.005	-0.133	-0.116	-0.017
Fruit supply	0.042	0.105	-0.063	0.035	0.033	0.002	-0.046	-0.040	-0.006
Veg. supply	0.063	0.157	-0.095	0.062	0.058	0.003	-0.068	-0.060	-0.009

Explanatory variable	Eggs Retail price	Eggs Farm price	Eggs Spread	Dairy Retail price	Dairy Farm price	Dairy Spread	Fresh Fruit Retail price	Fresh Fruit Farm price	Fresh Fruit Spread
				Percent change					
Nonfood price	2.477	0.604	1.873	1.143	1.065	0.078	-0.074	-0.535	0.460
Beverage price	0.054	0.013	0.041	0.138	0.129	0.009	0.019	0.138	-0.119
Sugar price	-0.915	-0.223	-0.692	-0.217	-0.202	-0.015	-0.052	-0.374	0.322
Cereal price	-0.801	-0.195	-0.606	0.270	0.251	0.018	0.111	0.796	-0.685
Income/capita	-1.449	-0.353	-1.096	-0.482	-0.449	-0.033	0.039	0.278	-0.239
Population	7.627	1.860	5.767	1.958	1.824	0.134	-0.084	-0.604	0.520
Wage	2.252	2.567	-0.315	0.496	2.093	-1.597	-1.073	0.547	-1.620
Packaging price	1.898	0.042	1.856	0.166	0.135	0.031	0.457	-0.350	0.808
Transport price	-2.010	-1.617	-0.394	-0.418	-0.639	0.221	0.307	0.576	-0.269
Energy price	-0.159	0.112	-0.271	0.001	-0.186	0.186	-0.131	-0.226	0.095
Cattle supply	-1.799	-0.439	-1.361	-0.314	-0.292	-0.021	0.110	0.794	-0.683
Hog supply	-0.680	-0.166	-0.514	-0.182	-0.170	-0.012	0.041	0.291	-0.251
Poultry supply	-0.407	-0.099	-0.308	-0.143	-0.134	-0.010	-0.026	-0.187	0.161
Egg supply	-0.733	-2.143	1.410	-0.038	-0.035	-0.003	0.006	0.043	-0.037
Milk supply	-0.321	-0.078	-0.243	-0.558	-1.632	1.074	0.007	0.050	-0.043
Fruit supply	0.082	0.020	0.062	0.011	0.010	0.001	-0.151	-1.009	0.858
Veg. supply	0.013	0.003	0.010	0.005	0.005	0.000	-0.007	-0.050	0.043

Explanatory variable	Fresh Vegetables Retail price	Fresh Vegetables Farm price	Fresh Vegetables Spread
		Percent change	
Nonfood price	-0.043	-1.197	1.155
Beverage price	0.005	0.151	-0.146
Sugar price	-0.023	-0.643	0.620
Cereal price	0.048	1.339	-1.291
Income/capita	0.021	0.589	-0.568
Population	-0.011	-0.319	0.307
Wage	-0.182	-0.146	0.037
Packaging price	0.298	-0.037	0.336
Transport price	0.224	0.698	-0.473
Energy price	-0.282	-0.249	-0.033
Cattle supply	0.050	1.404	-1.354
Hog supply	0.022	0.612	-0.590
Poultry supply	-0.012	-0.331	0.319
Egg supply	0.000	0.013	-0.012
Milk supply	0.001	0.037	-0.035
Fruit supply	-0.002	-0.060	0.058
Veg. supply	-0.375	-1.401	1.027

The effect of consumer substitution links the effects of changes in farm supply on retail and farm prices across the seven markets. For example, the first panel of table 8 suggests that increases in cattle supply depress both retail and farm prices for pork. Hence, increased cattle supply will lower both cattle prices (farm-level price) and consumer-level beef prices. Because our estimates suggest that consumers treat beef and pork as gross substitutes (appendix table 1), lower relative beef prices imply a reduction in consumer demand for pork. Because hogs are a normal factor of production in pork supply, hog prices (farm level) also fall. In this particular example, it is important to recall that estimates presented in table 8 exclude the effects of imports and exports of farm commodities.[25]

The column labeled "spread" summarizes the relative responses of retail and farm prices associated with a 1-percent increase in a particular explanatory variable. In particular, the estimates, computed in this double-log specification as the difference between the percent change in the retail and the percent change in the farm price, represent Gardner's response estimates of retail-to-farm price spreads. A positive (negative) sign implies that the market's retail price response exceeds (is less than) the response of the market's farm price. Because the estimates account for the response of a market's farm price and not the *retail equivalent* farm price based on variable proportions, the "spread" estimates reported in table 8 represent responses of spreads computed under the assumption of fixed-factor proportions.

Table 9 — Own-price elasticities of farm and consumer demand

Market	Own-farm price elasticity of industry demand for farm ingredients	Own-retail price elasticity of consumer demand
Beef and veal	-0.402	-0.065
Pork	-0.514	-0.745
Poultry	-1.074	-0.607
Eggs	-0.472	-0.064
Dairy	-0.627	-0.974
Fresh fruit	-1.020	-0.208
Fresh vegetables	-0.753	0.054

Table 9 compares the own-price elasticities of consumer demand with the full-reduced form, own-price elasticity of farm demand for farm ingredients for each market. Fixed-proportions production implies that an industry's input demand would be less own-price elastic than the own-price elasticity of retail demand. However, the results in table 9 suggest that for four of the seven markets, the industry's full-reduced-form demand for farm ingredients is more own-price elastic than consumer demand.[26] Such results call into question the validity of market price analyses based on fixed proportions.

[25] Attempts to interpret the results presented in this section are difficult because standard errors and t-tests are not computed. Such estimates might be computed using the methods suggested in footnote 22.

[26] The derived demand for farm ingredients reported in table 9 is as follows. If the full-reduced-form system (i.e., table 8) of farm price equations is represented in matrix notation as $\boldsymbol{P_f = \beta_1 F + \beta_2 X}$, where $\boldsymbol{F}$ is the vector of farm supplies, and $\boldsymbol{X}$ is the vector of all other explanatory variables, the estimates in table 9 are $\boldsymbol{\beta_1^{-1}}$. The own-price elasticities of derived demand for farm ingredients (accounting for consumer substitution) are the diagonal elements of $\boldsymbol{\beta_1^{-1}}$.

Conclusions

This report demonstrates that structural change can be incorporated directly into a very general and powerful set of market models. Early empirical findings of competitive U.S. food markets based on this general theory stood in sharp contrast to findings of market power based on the traditional and restrictive assumptions of fixed-factor proportions. By accounting for structural change, we have shown that food market data closely support this general theory and result in further evidence of competitive U.S. food industries.

In this report, we equate trends in market variables with structural change. We found that the most intriguing type of trend is the type that changes direction unexpectedly. Unpredictable changes in trends fit Kinsey and Senauer's description of structural change in food consumption and Paul's explanation of industry reorganization. Yet, it is precisely their unpredictability that poses obstacles to statistical inference. A main motivation of this report is to provide inference on competition, and we have relied on recent advances in econometrics to exploit such trends in market variables and test whether seven U.S. food industries operate competitively within an environment of structural change.

Although market variables may be driven by unpredictable changes in trends, it is necessary that models constructed from them do not. Our finding that retail- and farm-price relationships derived from this general theory are cointegrated supports the theory, as well as the claim, that changing trends in food consumption alter the structure of food industries. It also suggests that such models can be used to test for market power despite the many permanent changes in U.S. food markets that may have occurred over time.

Our finding of mostly competitive U.S. food industries offers important explanations of why concentrated food industries may be competitive. Paul argues that unexpected downward trends in consumer demand motivate firms to specialize and seek vertical arrangements that serve to spread the risk of downward trends in revenues. Our finding of cointegrated retail- and farm-price relationships suggests that competitive industries will reorganize to spread the risk of uncertain downward trends in consumer demand.

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Appendix

Structure of the WH Model

Underpinning the pair of quasi-reduced-form retail and farm price equations in the text equations 1 for each market is a general market-clearing condition for final industry output and a general market-clearing condition for farm ingredients. They are general because, within the same composite industry, firms' farm demand and firms' output supply will vary across firms. Within the same industry, one firm's production function may be different from other firms' production functions. Hence, the structure of this class of models provides an analytical framework for price analyses of market for which any one industry produces a variety of different final products. The comparative static results developed by Heiner, and extended by Wohlgenant (1989), do not depend on the restriction of identical firms. For a more detailed discussion of this structure, see WH or Wohlgenant (1989).

In the particular setup in the text, a food industry consists of all manufacturing, wholesaling, and retailing firms associated with a composite final demand product (e.g., beef). Each industry faces a final, market-level consumer demand schedule for its products and a market-level supply schedule for the farm ingredients. For tractability, firms in the same industry face the same composite market prices for output and all inputs.

In this setup, if *i* denotes an index associated with a firm in the industry, market clearing for final consumer goods could be generally written as

$$\Sigma_i S_r^i (P_r , P_f , W) = D_r (P_r, Z) \qquad (i)$$

where S_r^i is the supply of the ith firm, Pr the output price, P_f the farm price, W a vector of non-farm input prices, and Z is a consumer demand shifter. Condition (i) states that the sum of firm-level supply equals final consumer demand for the industry's output. Market clearing associated with farm ingredients is expressed as

$$\Sigma_i D_f^i (P_r , P_f , W) = F \qquad (ii)$$

in which *F* is farm supply, and D_f^i is the *i*th firm's demand for farm ingredients. Condition (*ii*) states that the sum of firm-level demands for farm ingredients equals the market-level supply of farm ingredients. Equations *i* and *ii* represent a general form of the structural model for each market examined in this study.

The structural parameters of equations *i* and *ii* define the A_r and the A_f of equations 1 in the text. The structural parameters of equations *i* and *ii* are found by totally differentiating equations *i* and *ii* and expressing the results in terms of partial elasticities of supply and demand. Totally differentiating equations *i* and *ii* and rearranging, gives

$$(e_{rr} - e)\, d\, lnP_r + e_{rf}\, d\, lnP_f = e_z\, d\, lnZ - e_{rw}\, d\, lnW$$
$$- e_{fr}\, d\, lnP_r - e_{ff}\, d\, lnP_f = e_{fw}\, d\, lnW - d\, ln\, F$$

where e_{rr}, for example, is the weighted sum of the elasticities of firm supply, i.e., $e_{rr} = \Sigma i\, e_{rr}^i (Q_r^i/Q_r)$. Wohlgenant and WH solve the two-equation system for $d\, lnP_r$ and $d\, lnP_f$ in terms of $d\, lnZ$, $d\, lnW$, *and* $d\, lnF$ and reveal the precise way in which the structural parameters define the coefficients of the quasi-reduced-form flexibilities (the *A* coefficients) in the text. The relationship between the structural parameters and the response coefficients are spelled out in the cited articles, and define the comparative static results discussed in the text.

Data

To the greatest extent possible, we followed the directions of Wohlgenant and Haidacher (WH) when constructing the economic variables used in this study. All of the variables of equations 1-4 in the text were constructed from annual data from 1958 to 1997. All of the potentially superfluous farm supply shift variables except the laying flock (for eggs) were also constructed from annual data from 1958 to 1997. The laying flock variable was constructed from 1960-97 data.

The retail prices are constructed from annual average Consumer Price Index (CPI) data, U.S. City Average, Not Seasonally Adjusted. The particular series used to construct retail prices are:

Beef and veal	Beef and Veal, CPI, U.S. City Average, Not Seasonally Adjusted
Pork	Pork, CPI, U.S. City Average, Not Seasonally Adjusted
Poultry	Poultry, CPI, U.S. City Average, Not Seasonally Adjusted

Eggs	Eggs, CPI, U.S. City Average, Not Seasonally Adjusted
Dairy	Dairy and Related Products, CPI, U.S. City Average, Not Seasonally Adjusted
Fresh fruit	Fresh Fruits, CPI, U.S. City Average, Not Seasonally Adjusted
Fresh veg.	Fresh Vegetables, CPI, U.S. City Average, Not Seasonally Adjusted
Nonfood	All Items Less Food, CPI, U.S. City Average, Not Seasonally Adjusted
Beverages	Non-Alcoholic Beverages and Beverage Material, CPI, U.S. City Average, Not Seasonally Adjusted
Sugar	Sugar and Sweets, CPI, U.S. City Average, Not Seasonally Adjusted
Cereal	Cereals and Bakery Products, U.S. City Average, Not Seasonally Adjusted

Farm level prices are constructed from non-seasonally adjusted annual Producer Price Index (PPI) data for farm products. The series used to construct farm price variables are:

Beef and veal	Slaughter Cattle, PPI, Farm Products, Not Seasonally Adjusted
Pork	Slaughter Hogs, PPI, Farm Products, Not Seasonally Adjusted
Poultry	Slaughter Poultry, PPI, Farm Products, Not Seasonally Adjusted
Eggs	Chicken Eggs, PPI, Farm Products, Not Seasonally Adjusted
Dairy	Fluid Milk, PPI, Farm Products, Not Seasonally Adjusted
Fresh fruit	Fresh Fruits and Melons, PPI, Farm Products, Not Seasonally Adjusted
Fresh veg.	Fresh Vegetables Except Potatoes, PPI, Farm Products, Not Seasonally Adjusted

The raw farm prices for beef and veal and pork (Pf*) are adjusted (as in Wohlgenant and WH) to account for the value of byproducts in the farm prices for slaughter cattle and slaughter hogs. We use the 1982 levels of the gross farm value (GFV) and the byproduct value (BPV), and hence the net farm value (NFV = GFV - BPV) for cattle and for hogs. These data are recorded by USDA/ERS. Using the 1982 values, data on byproduct allowances (Pb) for cattle and hogs are used to adjust the raw PPI farm price series. In general, the log of the adjusted farm price, $\ln P_f$ is computed as

$$\ln P_f = (GFV_{82}/ NFV_{82}) \ln P_f^* - (BPV_{82}/ NFV_{82}) \ln P_b$$

In particular, the adjusted farm price formula for beef is

$$\ln P_f = (155.5/141.1) \ln P_f^* - (14.4/ 141.1) \ln P_b$$

and is

$$\ln P_f = (94.3/87.0) \ln P_f^* - (7.3/87.0) \ln P_b$$

for pork.

All CPI and PPI price series were obtained from Bureau of Labor Statistics (BLS) web sites, and the GFV, BPV and byproduct allowances were obtained from ERS data sets with help from Lawrence Duewer.

The price of labor, packaging, transportation, energy, and other inputs from 1968-1997 were obtained directly from published reports (Elitzak). The series was extended back to 1958-1967 by overlapping the 1968-97 data set with a consistent and similar data set covering the 1958-67 period. The data were provided by Howard Elitzak.

The series used to construct farm supply are computed as the product of per capita food disappearance multiplied by U.S. Population, including armed services (July 1). The particular per capita disappearance data series used are:

Beef and veal	Beef plus Veal, Food Disappearance Per Capita, Carcass Weight (lbs)
Pork	Pork, Food Disappearance Per Capita, Carcass Weight (lbs)
Poultry	Total Chicken Plus Turkey, Food Disappearance Per Capita, Carcass Weight (lbs)

Eggs	Eggs, Food Disappearance Per Capita, Farm Weight (lbs)
Dairy	All Dairy Products, Food Disappearance Commercial Sales plus USDA donations (lbs of milk equivalent, milkfat basis)
Fresh fruit	Fresh Fruit (including melons), Food Disappearance Per Capita, Farm Weight (lbs)
Fresh veg.	Commercially Produced Fresh Vegetables Excluding Potatoes and Sweet Potatoes Minus Mushrooms, Commercial Disappearance Per Capita, Farm Weight (lbs)

The farm supply variables for beef and pork were converted from carcass weight to live weight by dividing the disappearance values (in carcass weight) by the annual average ratio of the dressed to live weight under federal inspection. The dressed and live weight data are reported in *Annual Livestock Slaughter*, the Agricultural Statistics Board, NASS, USDA, and were made available by Lawrence Duewer. The dependent variables used in the estimation of the system of consumer demand relationships were constructed using the same food disappearance variables, except no adjustment was made to convert the beef and pork disappearance data to live weight.

The data to compute shifters on consumer demand (other than prices) are:

Population	U.S. Population, including armed services, July 1
Income	Per Capita Personal Disposable Income, Current Dollars, multiplied by Population

The data used to form nonfarm input prices were provided by Howard Elitzak. They are:

Packaging	Food Marketing Cost Index (1982 = 100), packaging component
Transportation	Food Marketing Cost Index (1982 = 100), transportation component
Energy	Food Marketing Cost Index, energy component
Other	Food Marketing Cost Index, advertising, communications, rent maintenance and repair, business services, supplies, property taxes, short-term interest. This variable is used as the deflator in this bulletin.

Data to construct the shifters on the farm supply used to test for oligopsony power tests are:

Beef & veal	S_1: Steers, 1 year and older, January 1 S_2: Price of number 2 yellow corn, Chicago
Pork	S_1: U.S. hog inventory, all hogs and pigs S_2: Price of number 2 yellow corn, Chicago
Poultry	S_1: Price of soybean meal, 48 percent, Decatur S_2: Price of number 2 yellow corn, Chicago
Eggs	S_1: Laying flock, average annual per month
Dairy	S_1: Cows and heifers on farms, 2 years and older, January 1 S_2: Price of soybean meal, 48 percent, Decatur
Fresh fruit	S_1: Average hourly wages paid to all hired farm workers
Fresh vegetables	S_1: Average hourly wages paid to all hired farm workers

Estimates of Consumer Demand and Factor Shares

Appendix table 1 — System of consumer demand equations

Explanatory variable	Per capita consumer demand equation for: Beef	Pork	Poultry	Eggs	Dairy	Fresh fruit	Fresh vegetables
Beef price	-0.065 (-1.02)	0.574 (15.3)	-0.333 (-3.20)	0.191 (5.08)	0.089 (2.37)	-0.621 (-6.5)	-0.481 (-7.3)
Pork price	0.322 (15.0)	-0.745 (-20.0)	-0.085 (-1.2)	0.030 (.91)	0.069 (2.23)	-0.083 (-1.3)	-0.152 (-3.04)
Poultry price	-0.078 (-3.0)	0.029 (-.95)	0.607 (-6.0)	0.091 (2.60)	0.112 (3.36)	0.335 (5.02)	0.207 (3.70)
Egg price	0.045 (4.70)	0.012 (.79)	0.083 (2.31)	-0.064 (-2.18)	0.030 (1.50)	-0.101 (-2.90)	0.014 (.37)
Dairy price	0.053 (1.87)	0.085 (2.06)	0.303 (3.01)	0.084 (1.45)	-0.974 (-14.5)	-0.140 (-1.52)	-0.035 (-0.37)
Fresh fruit price	0.148 (-6.51)	-0.033 (-1.24)	0.327 (4.90)	-0.089 (-2.77)	-0.039 (-1.37)	-0.208 (-2.50)	0.135 (2.63)
Fresh vegetable price	0.093 (-7.22)	-0.050 (-2.89)	0.157 (3.59)	0.015 (.51)	-0.004 (-.17)	0.113 (2.66)	0.054 (.81)
Nonfood price	0.210 (1.93)	0.139 (2.00)	-1.170 (-6.67)	0.388 (5.5)	0.905 (13.8)	-0.008 (-.052)	-0.466 (-4.08)
Beverage price	-0.163 (-3.74)	0.164 (6.37)	0.093 (1.36)	0.012 (.47)	0.110 (4.78)	0.202 (3.12)	0.057 (1.32)
Sugar price	0.410 (4.51)	-0.367 (-6.50)	-0.701 (-4.48)	-0.195 (-3.3)	-0.101 (-1.9)	-0.455 (-3.0)	-0.253 (-2.53)
Cereal price	-0.668 (-4.56)	0.353 (3.57)	0.967 (3.53)	-0.129 (-1.2)	0.235 (2.51)	0.801 (2.92)	0.564 (3.19)
Income/ pop.	0.175 (5.33)	-0.103 (-3.23)	0.965 (12.2)	-0.335 (-9.0)	-0.430 (-13.8)	0.145 (2.10)	0.357 (5.66)
Intercept	1.005 (14.0)	-0.321 (-4.51)	6.291 (36.9)	2.797 (34.3)	5.406 (79.5)	5.010 (33.0)	5.344 (38.7)

Entries are elasticity estimates with symmetry and homogeneity imposed at the sample means. Values in parentheses are t-values. The shares of income used to impose the symmetry restriction at the point estimates are as follows: 0.316 (beef and veal), 0.0180 (pork), 0.008 (poultry), 0.0082 (eggs), 0.0241 (dairy), 0.0075 (fresh fruit), and 0.0062 (fresh vegetables).

Appendix table 2 — Factor shares for the seven industries

Industry	Labor	Packaging	Transport	Energy	Other	Farm
Beef & veal	0.1892	0.0731	0.0430	0.0215	0.1032	0.5700
Pork	0.2904	0.1122	0.0660	0.0330	0.1584	0.3400
Poultry	0.2156	0.0833	0.0490	0.0245	0.1176	0.5100
Eggs	0.1628	0.0629	0.0370	0.0185	0.0888	0.6300
Dairy	0.2244	0.0867	0.0510	0.0255	0.1224	0.4900
Fresh fruit	0.2948	0.1139	0.0670	0.0335	0.1608	0.3300
Fresh vegetables	0.2904	0.1122	0.0660	0.0330	0.1584	0.3400

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