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U.S. Food and Nutrient Demand and the Effects of Agricultural Policies

by

Jeffrey T. LaFrance

New Economic Approaches to Consumer Welfare and Nutrition
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Jeffrey T. LaFrance

ABSTRACT

An econometric model of annual per capita U.S. food and nutrition demand is developed. The model is a flexible, full rank two Gorman polar form. It is strictly aggregable across income, demographic variables, and variations in micro preference parameters. Parametric conditions for global quasi-concavity of the (quasi-)utility function are derived. The model is implemented with annual time series data on U.S. per capita food consumption for the sample period 1918-1994. A battery of new test statistics are developed for and applied to the following hypotheses: (1) strict exogeneity of income or total expenditures; (2) global symmetry and negative semidefiniteness of the Slutsky substitution matrix; (3) parameter stability in a multivariate, nonlinear regression model based on within sample residuals; and (4) weak separability of food items from all other goods in the representative consumer's preference function. The empirical results are very encouraging with respect to the strictures of economic theory, heretofore a virtually unheard of outcome. The model is used to analyze the food and nutrient consumption and consumer welfare impacts of the U.S. dairy program.

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U.S. Food and Nutrient Demand and the Effects of Agricultural Policies

1. Introduction

Farm and food policy in the United States is undergoing a major transformation. Most, though not all, farm-level price and income support programs are being replaced by cash payments and a move toward an open market. At the same time, welfare, food stamps, Women, Infants and Children (WIC), Aid to Families with Dependent Children (AFDC), and school lunch programs are being reduced in scope at the federal level and replaced by block grants to states. It almost goes without saying that these changes will influence the prices paid for and quantities consumed of food items and nutrients, as well as incomes and food expenditures of U.S. consumers. Exactly how much and in which directions these effects will be realized, however, is much more of an open question.

There are many reasons why it is not altogether clear what impacts these policy changes will have on the economic well-being, food consumption patterns, or nutritional intakes of U.S. consumers. One important reason is simply that we do not yet fully understand the joint influences of past policies on these matters, much less what will happen once the new policies begin to take effect. As an illustrative example, consider the joint economic impacts of the food stamp program and the U.S. dairy program. Food stamps provide direct in-kind subsidies for food consumption. The goal of the food stamp program is to increase the food consumption and nutritional status of the poor. The current food stamp program acts essentially as an income transfer mechanism.[1] On the other hand, price discrimination in federal milk marketing orders increase the retail price of fresh milk and lower the prices of manufactured dairy products (Heien; Ippolito and Masson; LaFrance and de Gorter; LaFrance 1992, 1993).[2] This creates incentives to substitute away from fresh, healthy foods and towards processed, less healthy foods.

As a second example, target prices for feed corn increase prices received by farmers, increasing the supply of corn. To clear these additional supplies from the market, prices paid by demanders of feed corn, chiefly hog and cattle feedlot operators, are lower than they otherwise would be.[3] The resulting decreases in input costs to the livestock sec-

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[1] That is, recipients currently do not have to pay for the food stamps received and nearly all recipients spend more on food than the value of food stamps. This implies that food stamp recipients are not at a "corner solution" on their budget constraint and the value of stamps received is equivalent to an income transfer of the same dollar amount.

[2] Many federal marketing orders and agreements for fruits, nuts, and vegetables also contain regulations that lead to higher prices for fresh products and lower prices for manufactured products (Jamison).

tor has the effect of increasing supplies of livestock to slaughterhouses, thereby reducing the market prices paid for red meat by consumers. The resulting increase in red meat consumption may be contrary to sound nutrition or health policy. It is commonly argued by nutritionists and healthcare professionals that foods that are high in animal fat, cholesterol (e.g., butter, cheese, prime rib, and bacon), salt, simple sugars, or chemical additives (e.g., canned and processed foods) do less to promote good nutrition and health than foods that are low in these factors and high in fiber, vitamins, and minerals (e.g., fresh fruits and vegetables, reduced fat fresh milk, yogurt, or fish).

The upshot is that, by and large, farm level policies (i.e., price supports and other commodity programs) create consumer incentives that directly oppose those created by food subsidy programs. What, then, can we say about the joint impact of domestic U.S. farm and food aid policies on food and nutrition consumption, health, and economic welfare of the U.S. population? At this juncture, very few unequivocal judgments can be reached. For example, while food aid recipients spend more on food, they probably eat less healthy foods due to relative price distortions. Certainly from a nutritional perspective, it is unclear whether this group is better or worse off with the combination of farm and food programs. It is not even altogether clear whether or not they are better off economically than might be the case with no government intervention in the farm and food sector. On the other hand, people that are neither farmers nor food aid recipients pay higher taxes to finance farm and food subsidies. This lowers disposable incomes, food expenditures, and economic welfare. In addition, under the scenarios described above, policy-induced price distortions create incentives to consume a less healthy mix of foods for members of this group. But very little is really known about the size of the net economic costs or the impacts on nutrition and health.

As a first cut at answering these important and interesting questions, this paper presents a model of U.S. food and nutrition consumption. The model is estimated econometrically using annual time series data for per capita U.S. food consumption and nutritional intake over the period 1919-1994. This empirical model is then applied to an analysis of consumers' economic welfare effects and the impacts on food and nutrition of the U.S. dairy program. The dairy program is chosen for the policy application for two reasons. First, it is one of a few commodity programs left largely unchanged in the most recent farm bill. Second, the structure of dairy policy permits a straightforward identification and econometric estimation of the impacts of the program at each level of the marketing channel (LaFrance and de Gorter; LaFrance 1993a, 1993b).

I offer several innovations to the theoretical and empirical analysis of food demand and policy. The theoretical model exploits household production theory (Becker; Lancaster 1966, 1971; Lucas; Michael and Becker; and Muth) to obtain a direct link be-

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[3] However, nonrecourse loans administered by the Commodity Credit Corporation place a floor on the market price determined by three- to five-year rolling averages of past market prices.

tween food and nutrition consumption which accommodates the existence of tradeoffs between nutrition and taste in food preferences (LaFrance 1983). A second theoretical contribution is the concept of a more general and plausible concept of aggregation, which I call strict aggregation. Simply put, strict aggregation is aggregation across individuals' incomes, demographics, and micro-level preference parameters to market-level demand equations which are consistent with the theory of consumer choice.[4] For the specific econometric model chosen for the empirical application, I also present a plausible solution to the issues posed by this concept.

Three econometric innovations are presented. First, explicit parametric restrictions which are necessary and sufficient for global quasi-concavity of preferences are derived and implemented. To my knowledge, this is the first successful derivation and implementation of such a coherent econometric structure on a large-scale demand model involving twenty-one food items and seventy-six time series observations. Second, a procedure based on the generalized methods of moments principle is derived for testing the hypothesis of strict exogeneity of income or total expenditure in a set of demand equations.[5] This procedure is closely related to, but more general and simpler to apply than, the standard Durbin-Wu-Hausman test. Third, a within sample, multivariate diagnostic test for model stability is derived. This diagnostic test is particularly useful in situations where there is a large number of parameters relative to the number of observations, so that split sample Chow tests or tests based on sequential post-sample recursive residuals (Brown, Durbin, and Evans; Harvey 1990, 1993; Hendry) are infeasible. In this paper, the most general formulation of the empirical model has 1596 observations, 615 structural parameters, and 232 parameters associated with the error covariance matrix and a common autocorrelation coefficient. Thus, I have less than two degrees of freedom per unknown parameter in the unrestricted model specification. This highly parameterized model (Zellner) precludes the use of these other diagnostic techniques.

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[4] ***Strict aggregation*** allows for different preferences across individuals in addition to those that arise from measurable factors such as demographics. This concept of aggregation is more general than, and consequently more limited in interpretation and application, than that of ***exact aggregation***, i.e., aggregation across income and demographics to the market level. See Stoker (1993) for an excellent recent survey of exact aggregation. Under exact aggregation, preferences of micro units are recovered from macro level demand equations. In contrast, under strict aggregation, a set of sufficient statistics are obtained for micro preferences from the macro level data, while individual micro-level preference functions can not be completely recovered.

[5] ***Strict exogeneity*** is the property of statistical independence between a right-hand-side regressor and the error term in a regression equation (Engle, Hendry, and Richard). When the regressor and the error term are normally distributed, strict exogeneity is equivalent to zero correlation.

With regard to the agricultural policy literature, this paper demonstrates that careful and comprehensive economic and econometric analyses which seek to coherently combine economic theory with empirical practice are both practical and informative. The empirical results are encouraging with respect to the parameter restrictions associated with economic theory - heretofore a virtually unheard of outcome. As a result, we are able to draw reasonable and logically consistent inferences based on the model estimates regarding the year-to-year and cumulative U.S. farm policy impacts on food consumption, nutritional intakes, and consumer welfare.

The organization of the paper is as follows. The next section considers several theoretical and econometric issues associated with the modeling problem. Section three characterizes the specific econometric model and its properties. Section four discusses the data, empirical results, hypothesis tests, and a battery of model diagnostics. The fifth section reports the results of the application of the econometric model to the analysis of the impacts of U.S. dairy policy over the period 1949-94 on food consumption, nutritional intakes, and consumer welfare. The final section summarizes and concludes.

2. Modeling Food Demand

It is reasonable to assume that food is eaten for two fundamental reasons — for its contribution to health due to nutritional intake and for its contribution to pleasure through flavor, odor, appearance, texture, and other qualities of the foods consumed. The relationship between nutrient intake and food consumption can be represented linearly. That is, "twice as much meat yields twice as much protein and twice as much fat, hence the technology must be homogeneous of degree one. Further, the amount of protein contained in an egg is not dependent of the amount of meat consumed, so the technology is additive" (Lucas, p. 167). This specification is independent of the household's welfare function for nutrients, and therefore does not relate to such findings from nutrition studies as (Dantzig; Hall; Foytik; Smith; and Stigler):

1. After certain levels of intake, additional quantities of nutrients yield decreasing (and sometimes eventually negative) returns to health.
2. The optimum quantity of any nutrient depends on the level of intake of the other nutrients.
3. Purely nutritional requirements appear to have at most a small effect on food expenditures.

Thus, let $\boldsymbol{z}$ denote an m-vector of nutrients important to the health status of the household, let $\boldsymbol{x}$ denote an n_x-vector of food items, and let $\boldsymbol{N}$ denote an $(m \times n_x)$ matrix of nutrient content per unit of food. Let the relationship between food consumed and nutrient availability be $\boldsymbol{z} = \boldsymbol{N}\boldsymbol{x}$. Also, let $\boldsymbol{y}$ denote an n_y-vector of all other goods, let $\boldsymbol{s}$ be a k-vector of demographic variables and other demand shifters, and write the consumer's utility function as $u(\boldsymbol{x}, \boldsymbol{y}, \boldsymbol{z}, \boldsymbol{s})$. The objective of the consumer is to

$$\text{(2.1)} \qquad \underset{\boldsymbol{x},\boldsymbol{y},\boldsymbol{z}}{\text{maximize}} \left\{ u(\boldsymbol{x}, \boldsymbol{y}, \boldsymbol{z}, \boldsymbol{s}) : \boldsymbol{x} \geq \boldsymbol{0},\ \boldsymbol{y} \geq \boldsymbol{0},\ \boldsymbol{p}_x'\boldsymbol{x} + \boldsymbol{p}_y'\boldsymbol{y} \leq m,\ \boldsymbol{N}\boldsymbol{x} = \boldsymbol{z} \right\},$$

where $\boldsymbol{p}_x$ is the vector of prices for $\boldsymbol{x}$, $\boldsymbol{p}_y$ is the vector of prices for $\boldsymbol{y}$, and m is income.

There is substantial empirical evidence that food is separable from non-food items in consumer preferences (deJanvry; numerous others). This is equivalent to separability of the utility function in the partition $\{(\boldsymbol{x}, \boldsymbol{z}), \boldsymbol{y}\}$,

$$(2.2) \qquad u(\boldsymbol{x}, \boldsymbol{y}, \boldsymbol{z}) = \tilde{u}(u_x(\boldsymbol{x}, \boldsymbol{z}), \boldsymbol{y}).$$

Separability permits us to focus on the maximization of the food sector sub-utility function, $u_x(\mathbf{x}, \mathbf{z})$, subject to a food expenditure budget constraint, $\mathbf{p}_x'\mathbf{x} = m_x$, where $\boldsymbol{p}_x$ is the vector of prices for $\boldsymbol{x}$ and m_x is total expenditure on food items, and the nutrient equations, $\boldsymbol{N}\boldsymbol{x} = \boldsymbol{z}$. Separability substantially reduces the dimension of the parameter space. In this paper, I consider (2.2) to be the model structure of interest, but nest the separability hypothesis (2.2) within the larger paradigm (2.1) along the line of inquiry suggested by LaFrance (1985), LaFrance and Hanemann, and Gorman (1995b) based on a price index for non-food items.

Now let $\boldsymbol{p} = [\boldsymbol{p}_x' \;\; \boldsymbol{p}_y']' \in \mathbb{R}_+^n$, where $n = n_x + n_y$, denote the vector of market prices for all goods and let the utility-maximizing conditional mean vector of quantities demanded given prices, income, demographics, and the nutrient content matrix be written as $E(\boldsymbol{x}|\boldsymbol{p}, m, \boldsymbol{s}, \boldsymbol{N}) \equiv \boldsymbol{h}^x(\boldsymbol{p}, m, \boldsymbol{s}, \boldsymbol{N})$. In the current framework, separability of $(\boldsymbol{x}, \boldsymbol{z})$ from $\boldsymbol{y}$ is equivalent to the demands for $\boldsymbol{x}$ having the structure

$$(2.3) \qquad \boldsymbol{h}^x(\boldsymbol{p}, m, \boldsymbol{s}, \boldsymbol{N}) \equiv \tilde{\boldsymbol{h}}^x(\boldsymbol{p}_x, \mu_x(\boldsymbol{p}, m, \boldsymbol{s}, \boldsymbol{N}), \boldsymbol{s}, \boldsymbol{N}),$$

where

$$(2.4) \qquad \mu_x(\boldsymbol{p}, m, \boldsymbol{s}, \boldsymbol{N}) \equiv \boldsymbol{p}_x'\boldsymbol{h}^x(\boldsymbol{p}, m, \boldsymbol{s}, \boldsymbol{N}) \equiv E\left(\boldsymbol{p}_x'\boldsymbol{x}|\,\boldsymbol{p}, m, \boldsymbol{s}, \boldsymbol{N}\right)$$

is the conditional mean of expenditure on $\boldsymbol{x}$ given prices, income, and demographic variables (Gorman 1995a; Blackorby, Primont, and Russell).[6]

The remainder of this section is devoted to three specific issues associated with the empirical implementation of (2.3) and (2.4) using aggregate time series data. First, I address the question of strict aggregation across individuals to coherent, theoretically consistent market level demand equations when income, demographics, and the microparameters of the individual's underlying utility function $u(\cdot)$ all vary across consumers. Second, I consider the empirical consequences of the fact that the conditional mean of total food expenditure, $\mu_x(\cdot)$, is a latent variable, while observed food expenditure is endogenous (Attfield 1985, 1991; Blundell 1986, 1988; Deaton 1975, 1986; Theil; LaFrance 1991). Last, I develop a robust (i.e., distribution independent), multivariate,

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[6] To see this, simply substitute $\boldsymbol{N}\boldsymbol{x}$ for $\boldsymbol{z}$ in $u(\cdot)$ to obtain the neoclassical utility maximization problem

$$\max_{x,y}\left\{u(\boldsymbol{x}, \boldsymbol{y}, \boldsymbol{N}\boldsymbol{x}, \boldsymbol{s}) : \boldsymbol{x} \geq \mathbf{0},\ \boldsymbol{y} \geq \mathbf{0},\ \boldsymbol{p}_x'\boldsymbol{x} + \boldsymbol{p}_y'\boldsymbol{y} \leq m\right\}.$$

within sample residuals based diagnostic test for model stability. In the following section, the solutions I propose for these three questions are dominating influences in the econometric application.

2.1 Strict Aggregation

There are several reasons to consider the effects of aggregation from micro units to macro, market level data in demand analyses. First, the effects of any policy vary across individuals. Eligibility for the food stamp program is based on income, household size, and total assets. Non-recipients share the cost of the program through income taxes, which vary with income. Second, it is highly likely that preferences differ across individuals. Some of this variation may be predictable with observable demographics like ethnicity, gender, or age characteristics of household members (Pollak and Wales). But available empirical evidence from cross-section studies suggests that variation in preferences across individuals remains after measurable influences have been accounted for. Finally, the theory of consumer choice applies to individual decision-makers, not to aggregate behavior. Although the economic rationality of the representative consumer is an interesting empirical hypothesis, without aggregation across economic agents there is no a priori reason to expect this property to hold.

But tracing the economic consequences of farm and food policies on prices, quantities traded, and so forth requires market-level data and analyses. At best, we might hope that a market-level analysis can provide basic answers to the following two questions: “What are the crude economic forces at work?” and “What are the aggregate economic gains and losses at issue?” We therefore seek a consumer choice model which could be consistent with rationality at the individual consumer level (even if this level of behavior is not observed or measured), accommodates empirical testing of aggregate economic rationality (i.e., of the representative consumer), and, barring rejection of the latter based on the available empirical evidence, provides a basis for reasonable estimates of the aggregate consumption and welfare effects of U.S. farm and food policy.

Let $\boldsymbol{d} = (m, \mathbf{s}')' \in \mathbb{R}^{k+1}$ denote the vector of income and other measurable demographic characteristics that distinguish between household types, let $\theta \in \mathbb{R}^{r}$ be the vector of micro parameters that vary across households, let $\Omega \subset \mathbb{R}^{k+1} \times \mathbb{R}^{r}$ be the set of household characteristics and micro parameters, and consider each household type $\omega = (\boldsymbol{d}, \theta)$ as an element of the set Ω. For notational simplicity, rewrite the mean demands for food items given prices, income, demographics, and micro-parameters in the form $\bar{\boldsymbol{x}}(\boldsymbol{p}, \omega)$. Also, let the compensating variation for the welfare effects due to a change in market prices from $\boldsymbol{p}^0$ to $\boldsymbol{p}^1$ be denoted by $cv(\boldsymbol{p}^0, \boldsymbol{p}^1, \omega)$, where

$$u^0 \equiv v(\boldsymbol{p}^0, m, \boldsymbol{s}, \theta) \equiv v(\boldsymbol{p}^1, m - cv(\boldsymbol{p}^0, \boldsymbol{p}^1, \omega), \boldsymbol{s}, \theta), \tag{2.3}$$

and $v(\boldsymbol{p}, m, \boldsymbol{s}, \theta)$ is the indirect utility function for problem (2.1) above, omitting the dependence on $\boldsymbol{N}$ for notation simplicity. Finally, let $(\Omega, \mathcal{F}, \psi)$ be a measure space, with $\psi: \Omega \to \mathbb{R}_+$ a finite, countably additive measure on $\mathcal{F} \equiv \sigma(\Omega)$, the smallest sigma algebra for the Borel subsets of Ω. Without loss in generality let $\psi(\Omega)=1$. Assume that ω, $cv(\boldsymbol{p}^0,\boldsymbol{p}^1,\omega)$, and $\bar{\boldsymbol{x}}(\boldsymbol{p},\omega)$ are ψ-integrable $\forall\ \boldsymbol{p},\boldsymbol{p}^0,\boldsymbol{p}^1\in\mathbb{R}^n_+$. The mean compensating variation with respect to the probability measure $\psi(\cdot)$ is defined by

(2.4) $$E\left[cv(\boldsymbol{p}^0,\boldsymbol{p}^1,\omega)\right]=\int_\Omega cv(\boldsymbol{p}^0,\boldsymbol{p}^1,\omega)d\psi(\omega).$$

Mean quantities demanded for foods are defined analogously. I require the following.[7]

Definition. *The demands for* $\boldsymbol{x}$ *are strictly aggregable if,* $\forall\ \boldsymbol{p},\boldsymbol{p}^0,\boldsymbol{p}^1\in\mathbb{R}^n_+$,

(a) $E\left[\boldsymbol{x}(\boldsymbol{p},\omega)\right]=\mathbf{x}\left[\boldsymbol{p},E(\omega)\right]$ *and*

(b) $E\left[cv(\boldsymbol{p}^1,\boldsymbol{p}^0,\omega)\right]=cv\left[\boldsymbol{p}^1,\boldsymbol{p}^0,E(\omega)\right]$.

Remark 1. Linearity of the nutrient equations, $\boldsymbol{z}=\boldsymbol{N}\boldsymbol{x}$, implies that nutrient demands are strictly aggregable if and only if food demands are strictly aggregable. Also note that strict aggregation is stronger than exact aggregation across a single function of income (Gorman 1953, 1961; Muellbauer, 1975, 1976) or across income and demographic variables (Stoker), since strict aggregation requires aggregation jointly across income, demographics, and individual-specific micro-parameters. The essential conditions required for strict aggregation are that the elements of ω enter $\bar{\boldsymbol{x}}(\boldsymbol{p},\omega)$ linearly and that the elements of $\boldsymbol{d}$ are uncorrelated in the joint distribution determined by $\psi(\cdot)$ with elements of θ that interact with $\boldsymbol{d}$. Thus, if the micro-parameters of the preference function vary across individuals, to be able to aggregate across consumers to the market level, we require the same linearity condition with respect to $\boldsymbol{d}$ as for exact aggregation, an additional linearity condition with respect to the variable micro-parameters, and at most bilinearity between $\boldsymbol{d}$ and θ with zero correlation between demographics and idiosyncratic micro-parameters.

Remark 2. A crucial aspect of strict aggregation is the fact that both quantities demanded and welfare measures must aggregate. A simple example motivates and illustrates the main issues involved. Let the indirect utility function be a full rank three Quadratic Expenditure System (Howe, Pollack, and Wales; van Daal and Merkies) of the form,

$$v(\boldsymbol{p},m)=-\frac{\sqrt{\boldsymbol{p}'\boldsymbol{B}\boldsymbol{p}}}{(m-\alpha(\boldsymbol{s})'\boldsymbol{p})}+\frac{\gamma'\boldsymbol{p}}{\sqrt{\boldsymbol{p}'\boldsymbol{B}\boldsymbol{p}}}.$$

[7] We could just as easily focus on the equivalent variation welfare measure, $ev(\boldsymbol{p}^0,\boldsymbol{p}^1,\omega)$, where

$$u'\equiv v(\boldsymbol{p}',m,\boldsymbol{s},\boldsymbol{\theta})\equiv v(\boldsymbol{p}^0,m+ev(\boldsymbol{p}',\boldsymbol{p}^0,\omega),\boldsymbol{s},\boldsymbol{\theta}).$$

However, equivalent variation is strictly aggregable if and only if compensating variation is.

By an application of Roy's identity, we have

$$\bar{\boldsymbol{x}}(\boldsymbol{p},\omega) = \alpha(\boldsymbol{s}) + \left(\frac{m - \alpha(\boldsymbol{s})'\boldsymbol{p}}{\boldsymbol{p}'\boldsymbol{B}\boldsymbol{p}}\right)\boldsymbol{B}\boldsymbol{p} + \left[\boldsymbol{I} - \frac{\boldsymbol{B}\boldsymbol{p}\boldsymbol{p}'}{\boldsymbol{p}'\boldsymbol{B}\boldsymbol{p}}\right]\gamma\,\frac{(m - (\boldsymbol{s})'\boldsymbol{p})^2}{\boldsymbol{p}'\boldsymbol{B}\boldsymbol{p}},$$

while the compensating variation for the price change $\boldsymbol{p}^0 \to \boldsymbol{p}^1$ is

$$\overline{cv}(\boldsymbol{p},\omega) = m - \alpha(\boldsymbol{s})'\boldsymbol{p}^1 - \left\{\frac{\sqrt{(\boldsymbol{p}^1)'\boldsymbol{B}\boldsymbol{p}^1/(\boldsymbol{p}^0)'\boldsymbol{B}\boldsymbol{p}^0} \times (m - \alpha(\boldsymbol{s})'\boldsymbol{p}^0)}{1 + \left[\frac{\gamma'\boldsymbol{p}^0}{\sqrt{(\boldsymbol{p}^1)'\boldsymbol{B}\boldsymbol{p}^1 \cdot (\boldsymbol{p}^0)'\boldsymbol{B}\boldsymbol{p}^0}} - \frac{\gamma'\boldsymbol{p}^1}{(\boldsymbol{p}^0)'\boldsymbol{B}\boldsymbol{p}^0}\right] \times (m - \alpha(\boldsymbol{s})'\boldsymbol{p}^0)}\right\}.$$

Now assume that $\alpha(\boldsymbol{s}) \equiv \alpha_0 + \boldsymbol{A}\boldsymbol{s}$, α_0 and $\boldsymbol{A}$ are uncorrelated with ω, and $\boldsymbol{B}$ is constant across individuals, while E(γ) = 0 and γ is (stochastically) independent of all other micro-parameters and demographic variables. Then quantities demanded aggregate to a model that is linear in income. However, given the other conditions, compensating variation aggregates if and only if γ = 0 with probability one. Thus, if we are interested in using an estimated aggregate demand model for consumer welfare analysis, then we must restrict attention to (at most) a rank two demand model.

2.2 Testing for Strict Exogeneity of Food Expenditures

Consider the empirical sub-system of demand equations

(2.5) $$\boldsymbol{x}_t = \boldsymbol{h}^x(\boldsymbol{p}_t, m_t, \boldsymbol{s}_t) + \varepsilon_t,\; t = 1,\ldots,T,$$

where ε_t is a vector of stochastic error terms. I continue to omit reference to the nutrient content matrix since it plays no role in the developments of this subsection. Assume that $\{\varepsilon_t\}$ is a multivariate martingale difference sequence, so that $E(\varepsilon_t) = \boldsymbol{0}\ \forall\ t$. Assume further that $E(\varepsilon_t\varepsilon_t') = \Sigma_t$ is a finite, positive definite $n_x \times n_x$ matrix ∀ t.[8] Given separability of $\boldsymbol{x}$ from $\boldsymbol{y}$, let observed food expenditures be defined by $m_{xt} \equiv \boldsymbol{p}_{xt}'\boldsymbol{x}_t$. Then we have

(2.6) $$m_{xt} = \mu_x(\boldsymbol{p}_t, m_t, \boldsymbol{s}_t) + \upsilon_t,$$

where $\upsilon_t \equiv \boldsymbol{p}_{xt}'\varepsilon_t \sim (0, \boldsymbol{p}_{xt}'\Sigma_t\boldsymbol{p}_{xt})$. Standard empirical practice would be to estimate a complete system of conditional demand equations for foods as a function of food prices, observed food expenditures, and demographic variables,

(2.7) $$\boldsymbol{x}_t = \tilde{\boldsymbol{h}}^x(\boldsymbol{p}_{xt}, m_{xt}, \boldsymbol{s}_t) + \tilde{\varepsilon}_t,$$

where

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[8] Note that, since foods comprise only a subset of all goods consumed, the budget identity does <u>not</u> imply that Σ_t is singular.

(2.8) $$\tilde{\varepsilon}_t \equiv \varepsilon_t + \tilde{\boldsymbol{h}}^x(\boldsymbol{p}_{xt}, \mu_{xt}(\boldsymbol{p}_t, m_t, \boldsymbol{s}_t), \boldsymbol{s}_t) - \tilde{\boldsymbol{h}}^x(\boldsymbol{p}_{xt}, m_{xt}, \boldsymbol{s}_t)$$

is the vector of conditional demand residuals.

It is clear from (2.8) that observed expenditure must be related to the residuals in the conditional demand equations in a non-trivial manner. The following lemma, which follows from Jensen's inequality (LaFrance 1991), shows that the simple but statistically indispensable property that the conditional error terms have vanishing means has important consequences for the model's structure.

Lemma 1. *$E(\upsilon_t | \boldsymbol{p}_t, m_t, \boldsymbol{s}_t) = 0$ and $E(\tilde{\varepsilon}_t | \boldsymbol{p}_t, m_t, \boldsymbol{s}_t) = \mathbf{0}$ if and only if*

$$\frac{\partial \tilde{\boldsymbol{h}}^x(\boldsymbol{p}_{xt}, m_{xt}, \boldsymbol{s}_t)}{\partial m_{xt}} \equiv \beta(\boldsymbol{p}_{xt}, \boldsymbol{s}_t).$$

The upshot is that a consistent stochastic specification between the conditional demand equations for food and total food expenditure restricts the functional form of the conditional demand model in the same way and for essentially the same reason that exact aggregation in income restricts the functional form of the unconditional demands.[9]

Given lemma 1, the relationship between the conditional and unconditional demand residuals is

(2.9) $$\tilde{\varepsilon}_t \equiv \left[\mathbf{I} - \frac{\partial \tilde{\boldsymbol{h}}^x(\boldsymbol{p}_{xt}, \mu_{xt}, \boldsymbol{s}_t)}{\partial m_{xt}} \boldsymbol{p}'_{xt} \right] \varepsilon_t,$$

so that the correlation between group expenditure and the conditional demand residuals is

(2.10) $$E(\tilde{\varepsilon}_t \upsilon_t | \boldsymbol{p}_t, m_t, \boldsymbol{s}_t) = \left[\mathbf{I} - \frac{\partial \tilde{\boldsymbol{h}}^x(\boldsymbol{p}_{xt}, \mu_t, \boldsymbol{s}_t)}{\partial m_{xt}} \boldsymbol{p}'_{xt} \right] \Sigma_t \boldsymbol{p}_{xt}.$$

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[9] Although it is always possible to modify the stochastic specification to construct a model with, say, budget shares on the left-hand-side and nonlinear functions of expenditure on the right-hand-side, a result analogous to lemma 1 applies to these cases as well. A coherent statistical model restricts our attention to (at most) rank two demand systems linear in a single nonlinear function of expenditure (Edgerton). Moreover, the nature of the available aggregate data on income or expenditure dictates the nature of the analogue to lemma 1 that must be applied. For example, the geometric mean of the population distribution of income (expenditure) requires a PIGLOG model, while a mean of order ρ, say, requires a PIGL specification. In this study, the income variable is per capita disposable income and the food expenditure variable is per capita food expenditure. Therefore, the focus throughout this paper is on demand models that are linear in income and food expenditure.

This leads to another lemma giving a necessary and sufficient condition for the strict exogeneity of food expenditure in the system of conditional food demand equations. A detailed proof of this result can be found in LaFrance (1997a).

Lemma 2. *If* $\Sigma_t \boldsymbol{p}_{xt} \neq \boldsymbol{0}$ *then* $E(\tilde{\varepsilon}_t \upsilon_t | \boldsymbol{p}_t, m_t, \boldsymbol{s}_t) = \boldsymbol{0}$ *if and only if*

$$\frac{\partial \tilde{\boldsymbol{h}}^x(\boldsymbol{p}_{xt}, \mu_{xt}, \boldsymbol{s}_t)}{\partial m_{xt}} \equiv \left(\boldsymbol{p}_{xt}' \boldsymbol{S}_t \boldsymbol{p}_{xt}\right)^{-1} \boldsymbol{S}_t \boldsymbol{p}_{xt} = \left(\boldsymbol{p}_{xt}' \Sigma_t \boldsymbol{p}_{xt}\right)^{-1} \Sigma_t \boldsymbol{p}_{xt}.$$

Where
$$\boldsymbol{S}_t = \frac{\partial \boldsymbol{h}^x(\boldsymbol{p}_t, m_t, \boldsymbol{s}_t)}{\partial \boldsymbol{p}_{xt}'} + \frac{\partial \boldsymbol{h}^x(\boldsymbol{p}_t, m_t, \boldsymbol{s}_t)}{\partial m_t} \boldsymbol{h}^x(\boldsymbol{p}_t, m_t, \boldsymbol{s}_t)'$$

is the $n_x \times n_x$ *sub-matrix of Slutsky substitution terms for food items.*

To construct a test of strict exogeneity of expenditure based on the generalized method of moments principle (Hansen), first note that lemma 2 defines a set of n_x moment conditions, so that

(2.11) $$\varepsilon_t \varepsilon_t' \boldsymbol{p}_{xt} = \varepsilon_t \upsilon_t = -\varphi_t \boldsymbol{S}_t \boldsymbol{p}_{xt} + \boldsymbol{u}_t,$$

where $E(\boldsymbol{u}_t) = \boldsymbol{0}\ \forall\ t$, $E(\boldsymbol{u}_t \boldsymbol{u}_t') = \Phi_t$, say, $\{\boldsymbol{u}_t\}$ is a multivariate martingale difference sequence, and $\varphi_t > 0$ is defined by

(2.12) $$\varphi_t \equiv \varphi(\boldsymbol{p}_t, m_t, \boldsymbol{s}_t) \equiv -\left(\boldsymbol{p}_{xt}' \Sigma_t \boldsymbol{p}_{xt}\right)^{-1} \left(\boldsymbol{p}_{xt}' \boldsymbol{S}_t \boldsymbol{p}_{xt}\right).$$

Define $\boldsymbol{z}_t = \varepsilon_t \upsilon_t$, $w_{it} = \sum_{j=1}^{n_x} s_{ijt} p_{x_j t}$, $\boldsymbol{w}_t = [w_{1t} \ \cdots \ w_{n_x t}]'$, and assume that $\boldsymbol{\Phi}_t$ is uniformly bounded $\forall$ t. Then estimate φ_t for each t by ordinary least squares (OLS),[10]

(2.13) $$\hat{\varphi}_t = -\left(\boldsymbol{w}_t' \boldsymbol{w}_t\right)^{-1} \boldsymbol{w}_t' \boldsymbol{z}_t.$$

The OLS residuals from this estimation procedure can be written as

(2.14) $$\hat{\boldsymbol{u}}_t = \left[\boldsymbol{I} - \boldsymbol{w}_t \left(\boldsymbol{w}_t' \boldsymbol{w}_t\right)^{-1} \boldsymbol{w}_t'\right] \boldsymbol{u}_t = \mathbb{M}_t \boldsymbol{u}_t,$$

and we have $E(\hat{\boldsymbol{u}}_t \hat{\boldsymbol{u}}_t') = \mathbb{M}_t \Phi_t \mathbb{M}_t$. Next, for each t define the within period average residual by $\bar{\hat{u}}_{\cdot t} = \sum_{i=1}^{n_x} \hat{u}_{i,t} / n_x$. Similarly, define the overall average residual by $\bar{\hat{u}} = \sum_{t=1}^{T} \bar{\hat{u}}_{\cdot t} / T$. In LaFrance (1997a), I show that a robust, asymptotically standard normal t-statistic to test for strict exogeneity of food expenditure is then equal to

[10] In practice, neither $\boldsymbol{z}_t$ nor $\boldsymbol{w}_t$ are observed. However, consistent estimates can be obtained readily. Since this does not alter any of the asymptotic results, for notational brevity this is ignored in the discussion.

(2.15) $$t = \frac{T\bar{\hat{u}}}{\sqrt{\sum_{t=1}^{T}\left(\bar{\hat{u}}_{\cdot t}\right)^2}}.$$

This test is closely related to, but different from the Durbin-Wu-Hausman (DWH) procedure to test for endogenous regressors. In the present context, the standard approach to the DWH test would be to estimate (2.6), so that

(2.16) $$m_{xt} = \hat{\mu}_x(\boldsymbol{p}_{xt}, \boldsymbol{p}_{yt}, m_t, \boldsymbol{s}_t) + \hat{\upsilon}_t,$$

then include $\hat{\upsilon}_t$ as well as m_{xt} in the conditional demand equations,

(2.17) $$\boldsymbol{x}_t = \alpha(\boldsymbol{p}_{xt}, \boldsymbol{s}_t) + \beta(\boldsymbol{p}_{xt}, \boldsymbol{s}_t)m_{xt} + \gamma\hat{\upsilon}_t + \boldsymbol{e}_{xt},$$

and calculate an F-test for $\gamma = \mathbf{0}$. But this is not a valid test of $E(m_{xt}\tilde{\boldsymbol{\varepsilon}}_{xt}) = 0$ because it ignores the (non-constant) effects of the terms $-\beta(\boldsymbol{p}_{xt}, \boldsymbol{s}_t)\upsilon_t$ in (2.7), and we do not obtain consistent parameter estimates.[11] Moreover, although properly estimating

(2.18) $$\boldsymbol{x}_t = \alpha(\boldsymbol{p}_{xt}, \boldsymbol{s}_t) + \beta(\boldsymbol{p}_{xt}, \boldsymbol{s}_t)\hat{\mu}_{xt} + \gamma\hat{\upsilon}_t + \boldsymbol{e}^*_{xt},$$

gives consistent parameter estimates, $\gamma = \mathbf{0}$ does not imply $E(m_{xt}\tilde{\boldsymbol{\varepsilon}}_{xt}) = 0$. The GMM alternative to the DWH test proposed here is straightforward to implement and invariant to distributional assumptions and alternative structures of the variance-covariance matrix for the demand residuals. It is a specific test of the necessary and sufficient condition that expenditure is uncorrelated with the conditional demand residuals. Thus, a rejection of this hypothesis also implies rejection of strict exogeneity.

2.3 Testing for Model Stability

As mentioned in the introduction, the sample period for the empirical application is 1919-1994. This period covers the Roaring Twenties, the Great Depression, World War II, the OPEC Oil Embargo, and all times in between these notable sub-periods. *Ex post*, it might stretch one's imagination to assert that the structure of food demand remained constant

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[11] The problem with the standard DWH procedure is that the explanatory part of the right-hand-side of (2.17) does not relate to a proper conditional expectation. In particular, if $\varepsilon_t \sim N(\mathbf{0}, \Sigma_t)$, then we have

$$E(\boldsymbol{x}_t | \boldsymbol{p}_{xt}, m_{xt}, \boldsymbol{s}_t) = \alpha(\boldsymbol{p}_{xt}, \boldsymbol{s}_t) + \beta(\boldsymbol{p}_{xt}, \boldsymbol{s}_t)m_{xt} + \left[(\boldsymbol{p}'_{xt}\Sigma_t\boldsymbol{p}_{xt})^{-1}\Sigma_t\boldsymbol{p}_{xt} - \beta(\boldsymbol{p}_{xt}, \boldsymbol{s}_t)\right]\upsilon_t.$$

Thus, to properly construct a DWH-type test, we need to jointly estimate all of the parameters associated with the term in the square brackets and perform an F-test of the restriction that the bracketed term vanishes identically across *t*. In the empirical model, depending on the degree that the model is restricted to meet the requirements of economic theory, this involves inverting the asymptotic variance-covariance matrix for either 434, 462, or 672 parameters. This is considerably more difficult to implement than the simple t-test in (2.15) above.

throughout this period. On the other hand, it is an interesting econometric question whether or not this may be true. The relevance of this issue is heightened by the very large number of parameters in a system of twenty-one demand equations and the corresponding need for as many time series observations as feasible and plausible.

Although many diagnostic procedures for testing model stability have been developed over the past two decades for single equation models, fewer are readily available for systems of equations. Moreover, in the present problem, the full sample period only provides about three degrees of freedom for each structural parameter estimated in the unrestricted model. This precludes the use of recursive one-step ahead forecast residuals or Chow tests based on sample splits to analyze the stability of the model parameters. Therefore, I need a test procedure which can be applied to within sample estimated errors. This is the focus of this subsection. The interested reader can find detailed derivations and arguments in LaFrance (1997b).

Let Σ_t be factored into $\boldsymbol{L}_t\boldsymbol{L}_t'$, where $\boldsymbol{L}_t$ is lower triangular and nonsingular. Define the random vector ξ_t by $\varepsilon_t = \boldsymbol{L}_t\xi_t$. In addition to our previous assumptions on the stochastic error terms $\boldsymbol{\varepsilon}_t$, we now also need to assume that $E(\varepsilon_{it}^4) < \infty \;\forall\; i, t$. The hypothesis of interest is that the $\boldsymbol{\varepsilon}_t$ are innovations in a stable econometric model, so that

$$H_o\colon\; \Sigma_t = \Sigma \;\forall\; t.$$

Under H_o, estimate the within period average sum of squared standardized residuals by

(2.19) $$\hat{\upsilon}_t = \tfrac{1}{n}\hat{\xi}_t'\hat{\xi}_t = \tfrac{1}{n}\hat{\varepsilon}_t\hat{\Sigma}^{-1}\hat{\varepsilon}_t,$$

where $\hat{\varepsilon}_t$ is the vector of residuals from the estimated system of demand equations for period t and $\hat{\Sigma} = \sum_{t=1}^{T}\hat{\varepsilon}_t\hat{\varepsilon}_t'/T$ is the estimated error covariance matrix. Then estimate the asymptotic variance of the $\hat{\upsilon}_t$ by

(2.20) $$\hat{\sigma}_\upsilon^2 = \frac{1}{T}\sum_{t=1}^{T}\left(\hat{\upsilon}_t^2 - 1\right).$$

Finally, for any $z \in [0,1]$ let $[zT]$ be the integer part of zT. In LaFrance (1997b), I offer a proof of the following.

> ***Proposition 1:*** $B_T(z) = \dfrac{1}{\sqrt{T}\hat{\sigma}_\upsilon}\cdot\sum_{t=1}^{[zT]}(\hat{\upsilon}_t - 1) \xrightarrow[T\to\infty]{D} B(z)$, *uniformly in z, where B(z) is a standard Brownian bridge on* [0,1].

For all $z \in [0,1]$, $B(z)$ has a Gaussian distribution, with mean zero and standard deviation $\sqrt{z(1-z)}$ (Bhattacharya and Waymire). For a given z - i.e., to test for a break point at a known date - an asymptotic 95% confidence interval for $B_T(z)$ is $\pm 1.96\sqrt{z(1-z)}$. To test for an unknown structural break, the statistic

(2.21) $$Q_T = \sup_{z \in [0,1]} \left| B_T(z) \right|$$

has an asymptotic 5% critical value of 1.36 (Ploberger and Krämer). It is important to note that sequential values of $B(z)$ are stochastically dependent, with correlation equal to $\sqrt{z_1(1-z_2)/(1-z_1)z_2}$ for $0 \le z_1 \le z_2 \le 1$. Thus, the values of $B(z)$ for adjacent time periods in the interior of the sample period are highly correlated. This is because B(z) is a Brownian bridge, i.e., $B(z) = W(z) - zW(1)$, where $W(z)$ is standard Brownian motion. Thus, one should not expect $B_T(z)$ to behave anything at all like "white noise."

3. The Econometric Model

For the empirical application in the following section, I use a simplified version of (2.1) above based on the concept of weak integrability (LaFrance 1985; LaFrance and Hanemann). It is only possible to recover part of the preference map from a subset of demands (Epstein; Hausman 1981; LaFrance 1985; LaFrance and Hanemann) and as shown by LaFrance and Hanemann, there is little loss in generality from aggregating over all nonfood items into essentially a Hicks composite commodity for total nonfood expenditures. Thus, for the remainder of this paper, let y be a scalar representing nonfood expenditures, and let $\pi(\boldsymbol{p}_{yt})$ denote a known, increasing, linearly homogeneous, and concave price index for nonfood items. Then the (quasi-)utility function for foods, nutrients, and nonfood expenditures for this study is defined by

(3.1) $$\begin{aligned} u(\boldsymbol{x}, y, \boldsymbol{z}, \boldsymbol{s}) = {} & \tfrac{1}{2}(\boldsymbol{x} - \alpha_1(\boldsymbol{s}))' \boldsymbol{B}_{xx} (\boldsymbol{x} - \alpha_1(\boldsymbol{s})) + \tfrac{1}{2}\beta_{yy}(y - \alpha_2(\boldsymbol{s}))^2 \\ & + \tfrac{1}{2}(\boldsymbol{z} - \alpha_3(\boldsymbol{s}))' \boldsymbol{B}_{zz} (\boldsymbol{z} - \alpha_3(\boldsymbol{s})) + (\boldsymbol{x} - \alpha_1(\boldsymbol{s}))' \beta_{xy} (y - \alpha_2(\boldsymbol{s})) \\ & + (\boldsymbol{x} - \alpha_1(\boldsymbol{s}))' \boldsymbol{B}_{xz} (\boldsymbol{z} - \alpha_3(\boldsymbol{s})) + (y - \alpha_2(\boldsymbol{s}))\beta'_{yz}(\boldsymbol{z} - \alpha_2(\boldsymbol{s})). \end{aligned}$$

Two desirable properties of this model specification are that it is a second-order flexible functional form for direct preferences that generates demands for foods and nutrients that are linear in income. Furthermore, we have the following result on strict aggregation.

> ***Proposition 2.*** *The preference function (3.1) is strictly aggregable if and only if*
>
> *(a)* $\alpha_i(\boldsymbol{s}) = \alpha_{i,0} + \boldsymbol{A}_i\boldsymbol{s}$, *i = 1,2,3;*
>
> *(b)* $\boldsymbol{B}_{xx}$, β_{yy}, $\boldsymbol{B}_{zz}$, β_{xy}, *and* β_{zy} *are constant across individuals; and*
>
> *(c)* $E(\boldsymbol{A}_i\boldsymbol{s}) = E(\boldsymbol{A}_i) \cdot E(\boldsymbol{s})$, *i = 1,2,3, in the joint distribution for* ω.

Proof: Substitute $\boldsymbol{Nx}$ for $\boldsymbol{z}$ in (3.1) and consider the problem of maximizing $u(\boldsymbol{x},y,\boldsymbol{Nx},\boldsymbol{s})$ with respect to $(\boldsymbol{x},y)$ subject to $\boldsymbol{p}_x'\boldsymbol{x} + \pi(\boldsymbol{p}_y)y \le m$. At an interior solution, the demands for $\boldsymbol{x}$ can be written as

(3.2) $$\boldsymbol{h}^x(\boldsymbol{p}_x, \pi, m, \boldsymbol{s}) = \alpha_x(\boldsymbol{s}) + \left(\frac{m - \alpha_x(\boldsymbol{s})'\boldsymbol{p}_x - \alpha_y(\boldsymbol{s})\pi}{\boldsymbol{p}_x'\boldsymbol{C}_{xx}\boldsymbol{p}_x + 2\boldsymbol{p}_x'\gamma_{xy}\pi + \gamma_{yy}\pi^2} \right) \cdot \left(\boldsymbol{C}_{xx}\boldsymbol{p}_x + \gamma_{xy}\pi \right),$$

where

$$\boldsymbol{C} = \begin{bmatrix} \boldsymbol{C}_{xx} & \gamma_{xy} \\ \gamma'_{xy} & \gamma_{yy} \end{bmatrix} = \begin{bmatrix} \boldsymbol{B}_{xx} + \boldsymbol{B}_{xz}\boldsymbol{N} + \boldsymbol{N}'\boldsymbol{B}_{zx} + \boldsymbol{N}'\boldsymbol{B}_{zz}\boldsymbol{N} & \beta_{xy} + \boldsymbol{N}'\beta_{zy} \\ \beta'_{xy} + \beta'_{zy}\boldsymbol{N} & \beta_{yy} \end{bmatrix}^{-1},$$

and

$$\begin{bmatrix} \alpha_x(\boldsymbol{s}) \\ \alpha_y(\boldsymbol{s}) \end{bmatrix} = \begin{bmatrix} \boldsymbol{C}_{xx} & \gamma_{xy} \\ \gamma'_{xy} & \gamma_{yy} \end{bmatrix} \begin{bmatrix} \boldsymbol{B}_{xx} + \boldsymbol{N}'\boldsymbol{B}_{zx} & \beta_{xy} + \boldsymbol{N}'\beta_{zy} & \boldsymbol{B}_{xz} + \boldsymbol{N}'\boldsymbol{B}_{zz} \\ \beta'_{xy} & \beta_{yy} & \beta'_{zy} \end{bmatrix} \begin{bmatrix} \alpha_1(\boldsymbol{s}) \\ \alpha_2(\boldsymbol{s}) \\ \alpha_3(\boldsymbol{s}) \end{bmatrix}.$$

Similarly, the compensating variation for a change in prices from $\boldsymbol{p}_x^0$ to $\boldsymbol{p}_x^1$ is defined by

$$cv(\boldsymbol{p}_x^0, \boldsymbol{p}_x^1, m, \boldsymbol{s}) = \boldsymbol{m} - \alpha_x(\boldsymbol{s})'\boldsymbol{p}_x^1 - \alpha_y(\boldsymbol{s})\pi(\boldsymbol{p}_y) \tag{3.3}$$

$$- \left(\boldsymbol{m} - \alpha_x(\boldsymbol{s})'\boldsymbol{p}_x^1 - \alpha_y(\boldsymbol{s})\pi(\boldsymbol{p}_y)\right) \cdot \sqrt{\frac{(\boldsymbol{p}_x^1)'\boldsymbol{C}_{xx}\boldsymbol{p}_x^1 + 2(\boldsymbol{p}_x^1)'\gamma_{xy}\pi(\boldsymbol{p}_y) + \gamma_{yy}\pi(\boldsymbol{p}_y)^2}{(\boldsymbol{p}_x^0)'\boldsymbol{C}_{xx}\boldsymbol{p}_x^0 + 2(\boldsymbol{p}_x^0)'\gamma_{xy}\pi(\boldsymbol{p}_y) + \gamma_{yy}\pi(\boldsymbol{p}_y)^2}}.$$

The proposition follows immediately from the definition and (3.2) and (3.3). **Q.E.D.**

Simply stated, this result shows that strict aggregation has empirical content, while nevertheless admitting a flexible functional form for the aggregate food demand model.

Due to the adding up condition, $\boldsymbol{p}_x'\boldsymbol{x} + m_y \equiv m$, heteroskedasticity considerations lead naturally to an empirical specification with deflated expenditures, rather than quantities, as the left-hand-side variables. Abusing notation slightly, then, the actual empirical model is

$$\boldsymbol{e}_x \equiv \boldsymbol{P}_x\boldsymbol{x} = \boldsymbol{P}_x\alpha_x(\mathbf{s}) + \left(\frac{m - \alpha_x(\mathbf{s})'\boldsymbol{p}_x - \alpha_y(\boldsymbol{s})}{\boldsymbol{p}_x'\boldsymbol{C}_{xx}\boldsymbol{p}_x + 2\gamma'_{xy}\boldsymbol{p}_x + \gamma_{yy}}\right)\boldsymbol{P}_x\left(\boldsymbol{C}_{xx}\boldsymbol{p}_x + \gamma_{xy}\right) + \varepsilon_x, \tag{3.4}$$

where m and $\boldsymbol{p_x}$ have now been deflated by $\pi(\boldsymbol{p_y})$ and $\boldsymbol{P}_x \equiv \text{diag}(p_{xi})$. The budget identity then implies $\iota'\varepsilon_x + \varepsilon_y \equiv 0$, where ι is an n_x-vector of ones and ε_y is the residual for total expenditures on nonfood items.

The estimation procedure is nonlinear seemingly unrelated regressions equations (SURE) with one iteration on the residual covariance matrix. This produces consistent, efficient, and asymptotically normal parameter estimates under standard conditions (Rothenberg and Leenders), while avoiding spurious overfitting of a subset of equations, which can result from iterative SURE methods.[12]

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[12] The reason for this numerical result with iterative SURE in small samples with numerous shared parameters across equations can be best understood by writing the estimated covariance matirx, say $\boldsymbol{S}$, at a given iteration in factored form as $\boldsymbol{S} = \boldsymbol{Q}\Delta\boldsymbol{Q}'$, where $\boldsymbol{QQ}'$ =

3.1 Parameter Restrictions and Test Procedures

First, I make note of the fact that the right-hand-side of (3.2) is zero degree homogeneous in the elements of $\boldsymbol{C}$, so that a normalization is required for identification. A useful choice is $\gamma_{yy} = 1$, which replaces $\boldsymbol{C}$ with $-\boldsymbol{C}$ and fixes the lower diagonal element at unity. This results in the Gorman polar form representation of indirect preferences. This is convenient for deriving the parameter restrictions for global quasi-concavity.

Second, in this model, separability of foods from nonfood expenditures, which in turn is necessary and sufficient for separability of foods from all other goods (LaFrance and Hanemann) is equivalent to the restrictions $\boldsymbol{\gamma_{xy}} = 0$.

The $n_x \times n_x$ submatrix of Slutsky substitution terms for food items is

$$(3.5) \qquad \boldsymbol{S} = \left(\frac{m - \boldsymbol{\alpha}_x(\mathbf{s})'\boldsymbol{p}_x - \alpha_y(\boldsymbol{s})\pi}{\boldsymbol{p}_x'\boldsymbol{C}_{xx}\boldsymbol{p}_x + 2\boldsymbol{p}_x'\boldsymbol{\gamma}_{xy}\pi_x + \pi^2} \right) \left[\boldsymbol{C}_{xx}' - \left(\frac{(\boldsymbol{C}_{xx}\boldsymbol{p}_x + \boldsymbol{\gamma}_{xy}\pi)(\boldsymbol{p}_x'\boldsymbol{C}_{xx} + \boldsymbol{\gamma}_{xy}'\pi)}{\boldsymbol{p}_x'\boldsymbol{C}_{xx}\boldsymbol{p}_x + 2\boldsymbol{p}_x'\boldsymbol{\gamma}_{xy}\pi + \pi^2} \right) \right].$$

Hence, $\boldsymbol{S}$ is symmetric if and only if $\boldsymbol{C}_{xx}$ is also and global symmetry is accomplished with $\frac{1}{2}n_x(n_x\text{-}1)$ parameter restrictions on $\boldsymbol{C}_{xx}$.

However, symmetry of $\boldsymbol{S}$ only guarantees the existence of the direct and indirect preference functions, it does not ensure the proper curvature associated with utility maximization. The necessary and sufficient condition for the observed demands to be consistent with utility maximization is quasi-concavity. Quasi-concavity of the (quasi-) utility function in $(\boldsymbol{x}, y)$, in turn, implies that at least n_x eigen values of $-\boldsymbol{C}$ must be negative (Lau). Hence, at least n_x of the eigen values of $\boldsymbol{C}$ must be positive for quasi-concavity. However, given separability, the (quasi-)utility function (3.1) is additively separable in $\boldsymbol{x}$ and y. Under this condition, quasi-concavity requires that preferences must be concave either in $\boldsymbol{x}$ or in y (Gorman 1995c).[13] Treating foods and nonfood items sym-

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$\boldsymbol{Q}'\boldsymbol{Q} = \boldsymbol{I}$, and $\Delta = \text{diag}(\delta_i)$ is the diagonal matrix of eigen values. If one or more of the δ_i is "small" relative to all others, then since $\boldsymbol{S}^{-1}$ is held fixed during the next iteration on the structural parameters, the linear combination of the ε_t's associated with that eigen value will carry a "large" weight relative to all others in the sum of squares criterion. The associated linear combination of the residuals will tend to become closer and closer to a perfect fit, creating an artificial singularity (i.e., one that only results in finite samples).

[13] For strict quasi-concavity, this can be easily demonstrated as follows. Strict quasi-concavity requires

$$[d\boldsymbol{x}' \quad dy]\begin{bmatrix} u_{xx} & \mathbf{0} \\ \mathbf{0}' & u_{yy} \end{bmatrix}\begin{bmatrix} d\boldsymbol{x} \\ dy \end{bmatrix} < 0 \ \forall \begin{bmatrix} d\boldsymbol{x} \\ dy \end{bmatrix} \neq \begin{bmatrix} \mathbf{0} \\ 0 \end{bmatrix} \ni [d\boldsymbol{x}' \quad dy]\begin{bmatrix} u_x \\ u_y \end{bmatrix} = 0.$$

Setting $dy = 0$ implies that $d\boldsymbol{x}'u_{xx}d\boldsymbol{x} < 0 \ \forall \ d\boldsymbol{x}'u_x = 0$, so that the sectoral utility function for foods must be strictly quasi-concave. But if $\boldsymbol{u_{xx}}$ is indefinite (has exactly one positive metrically in this regard implies that the eigen values of $\boldsymbol{C}_{xx}$ must all be positive. This is straightforward (although admittedly tedious and numerically very intensive) to implement. Specifically, let $\boldsymbol{C}_{xx} = \boldsymbol{LL}'$, where $\boldsymbol{L}$ is a lower triangular matrix with nonzero diagonal elements, so that $\boldsymbol{C}_{xx}$ is positive definite. These explicit, nested parameter restrictions ensure that the (quasi-) utility function (3.1) is globally weakly integrable (LaFrance and Hanemann).

In the complete sample, there are twenty-one equations ($N = 21$) and seventy-six usable time series data points ($T = 76$), for a total of 1596 observations. The unrestricted model (where $\boldsymbol{C}_{xx}$ is neither symmetric nor positive definite) has 615 parameters ($K = 615$), which leaves 981 degrees of freedom. There are 210 parameter restrictions associated with symmetry of $\boldsymbol{C}_{xx}$ ($G = 210$) and 238 restrictions associated with symmetry and positive definiteness of $\boldsymbol{C}_{xx}$ ($G = 238$).[14] In large demand models such as this, the classical Wald (W), likelihood ratio (LR), and Lagrange multiplier (LM) asymptotic χ^2 test statistics are well-known to be substantially biased towards rejecting a true null hypothesis too often (Laitinen; Meisner; Bera, Byron and Jarque). W is largest and most likely to reject a true null, while LM is smallest and therefore least likely to reject. Careful examination of the Monte Carlo results of Bera, Byron and Jarque also reveals the simple, intuitively appealing degrees of freedom correction $(NT-K)/(G\cdot NT)$, combined with critical values from the $F(G, NT\text{-}K)$ distribution to under-correct W and LR and to over-correct LM.

The approach that I take to this issue is to construct an approximate F-test based on the Lagrange multiplier principle. This test at least partially overcomes the problems of the classical χ^2 tests, while mitigating the tendency of the simple degrees of freedom adjustment to over correct the LM test. Let a "^" denote unrestricted estimates, let a "~" denote restricted estimates, and let the variance-covariance matrix for $\boldsymbol{\varepsilon}_x$ be denoted by $\boldsymbol{\Sigma}$. Given a first round estimate for $\boldsymbol{\Sigma}$, say $\boldsymbol{S}$, the least squares criterion for the SURE estimates is

$$s(\boldsymbol{S}) = \sum_{t=1}^{T} \varepsilon_{xt}' \boldsymbol{S}^{-1} \varepsilon_{xt}. \tag{3.6}$$

---

eigen value) and $u_{yy} > 0$, then the sign condition for quasi-concavity of u jointly in $\boldsymbol{x}$ and y fails.

[14] The maximal number of independent parameters associated with a Choleski factorization $\boldsymbol{LL}'$ that is positive semidefinite with rank n_x-g for $0 \le g \le n_x$ has the elements of the lower right triangle of $\boldsymbol{L}$ all equal to zero. This block has seven rows and columns in the quasi-concave model, which is the same number of negative eigen values that appear in the symmetry restricted model. Setting $l_{ii} = 0.01$ and $l_{ji} = 0$ for $j > i$ in those rows and columns ensures positive definiteness. This generates ½g(g+1) = 28 parameter restrictions for the binding curvature constraints, in addition to the symmetry restrictions.

Denote the first round estimate of Σ obtained with the unrestricted specification by $\hat{\Sigma}$, the corresponding estimate of Σ obtained with the restricted model by $\tilde{\Sigma}$, the second round unrestricted sum of squares given $\hat{\Sigma}$ by $\hat{s}(\hat{\Sigma})$, the unrestricted sum of squares given $\tilde{\Sigma}$ by $\hat{s}(\tilde{\Sigma})$, and the restricted sum of squares given $\tilde{\Sigma}$ by $\tilde{s}(\tilde{\Sigma})$. Minimization implies that $\hat{s}(\hat{\Sigma}) \leq NT$, with equality if and only if the convergent iterative SURE estimates (maximum likelihood under normality) just happens to occur at the end of the second round of the unrestricted estimation procedure. The *F*-statistic is calculated as

$$F(G, NT-K) = \frac{\left(\tilde{s}(\tilde{\Sigma}) - \hat{s}(\tilde{\Sigma})\right)\big/G}{\hat{s}(\hat{\Sigma})\big/(NT-K)}. \tag{3.7}$$

The numerator converges in distribution to a $\chi^2(G)/G$ random variable. It is calculated using the first round variance-covariance matrix obtained from the restricted model specification. This is consistent with the Lagrange multiplier principle and is well-known to have the smallest empirical size among W, LR, and LM. Under normality of the residuals, the denominator converges in distribution to a $\chi^2(NT-K)/(NT-K)$ random variable. It is calculated using the unrestricted model for both the first and second round estimates. The test is asymptotically valid even if the errors are not normally distributed. In that case, the denominator converges to one, while $F(G, NT-K) \xrightarrow[T\to\infty]{D} \chi^2(G)/G$.

3.2 Data

The data set consists of annual time series observations over the period 1918-1994. Per capita consumption of twenty-one food items and corresponding average retail prices for those items were constructed from several USDA and Bureau of Labor Statistics (BLS) sources. The quantity data are aggregates taken from the USDA series *Food Consumption, Prices and Expenditures*. Estimated retail prices corresponding to the quantity data were constructed as follows. Detailed disaggregated retail price estimates that are available for 1967 were used along with the respective quantity observations to construct an average retail price per pound in 1967 for each food category (e.g., beef). For all other years, the fixed 1967 quantity weights, together with either consumer price indices for disaggregate food items or average retail food prices were combined to construct a consistent retail price series for each commodity. The consumer price index (CPI) for all nonfood items is used for the "price" of nonfood expenditures.

The demographic factors included in the data are the first three moments (mean, variance, and skewness) of the empirical age distribution for the U.S. population and proportions of the U.S. population that are Black and neither White nor Black. The estimated age distribution is based on ten-year age intervals, plus categories for children less than five-years old and adults that are sixty-five years old and older. The ethnic variables are linearly interpolated estimates of Bureau of Census figures reported on 10-year intervals. I also allow for habit formation by including lagged quantities as elements of $\boldsymbol{s}$. This reduces the effective sample period to 1919-1994, with 1918 required for initial conditions, a total of 76 annual time series observations.

The income variable is per capita disposable income. Initially, no corrections are made for either the Great Depression or World War II. In so far as I am attempting to understand and explain consumer food demand choices, rather than equilibrium prices or producer supply responses, it seems appropriate to "let the data speak for itself" during these relatively volatile periods. For all versions of the model, stability of the model parameters is tested using the procedure described in the previous section.

With the generous aid of Nancy Raper, Shirley Gerrior, and Claire Zizza of the Human Nutrition Information Service (HNIS), annual estimates of the percentages of the total availability of seventeen nutrients from each of the twenty-one food categories were compiled for the period 1948-1992. These percentages were multiplied by the respective total supply of nutrients per capita and divided by the respective per capita consumption of each food item to obtain year-to-year estimates of the average nutrient content per pound of each food item - e.g., the number of grams of protein per pound of beef.

Several difficulties were encountered with these year-to-year nutrient content estimates. First, there are only slight annual changes in these data over the entire period from 1948-1992. A non-constant $\boldsymbol{N}$ matrix makes the model parameters time-varying. In principle, a time-varying $\boldsymbol{N}$ matrix would permit the separate identification and estimation of the preference parameters associated with both nutrition and taste. This is impossible with an essentially constant $\boldsymbol{N}$ matrix. Second, the construction of the annual nutrient content matrices creates a simultaneity problem. That is, the elements of $\boldsymbol{x}$ are used to calculate the elements of $\boldsymbol{N}$ each year, so that quantities demanded tacitly end up on both sides of the demand equations. Third, the percentage contribution estimates are reported with only two or three significant digits. This generates errors in variables, and exaggerates the changes in $\boldsymbol{N}$ over time. As a result, on the advice of the HNIS, the nutrient content matrix is assumed constant across years using the average of the 1952-1983 annual estimates for $\boldsymbol{N}$. This 32-year period represents the longest time frame available with a completely consistent set of disaggregated percentage contribution estimates. Table 1 presents the estimated average nutrient content for food items matrix.

4. Empirical Results

Table 2 presents the complete set of model diagnostics for two samples: 1919-94, including World War II; and 1919-41 and 1947-94, which excludes World War II plus 1946 to re-initialize the difference equations due to the presence of habit formation. The rationale for this can be explained by the results of the model stability tests at the bottom of table 2 and the plots of the partial sums $B_T(z)$ in figure 1. In the figure, the top panel shows the case where the full sample is included, while the bottom panel depicts the case where the years 1942-46 are excluded from the sample period. It is noteworthy that the unrestricted model does not show evidence of a structural break over the complete sample, while both the symmetry and quasi-concavity restricted specifications clearly suggest that a break occurred during the War years. On the other hand, the F-test for the symmetry restriction does not reject this hypothesis even at the 20 percent level of significance. Hence, if we conclude that symmetry is true, then given symmetry there is fairly strong evidence of a structural break during the second world war.

Parameter instability is a well-known cause of contaminated inference. Hence, it is prudent empirical practice in the above circumstances to re-estimate the model with the years 1941-46 excluded from the sample. When this is done, the results of the hypothesis tests for both the theoretical restrictions on the model parameters and model stability are encouraging. The symmetry hypothesis is not rejected at a 5 percent level of significance level, while global quasi-concavity is not rejected at the 10 percent level. Neither of these two restricted specifications show evidence of a further structural break at the 10 percent level of significance. While the unrestricted model suggests a marginal rejection of model stability, with the period since 1980 possibly having a different structure that previous years, this is tempered by three factors. First, neither of the restricted specifications are rejected in favor of the unrestricted model. Second, neither restricted specification presents strong evidence against a stable model structure, absent the war years. Third, the unrestricted model shows no evidence of model instability when the full sample is used. I conclude that the fully restricted, globally quasi-concave model over the period 1919-41 and 1947-94 is a reasonable model of aggregate U.S. food consumption.

Additional encouraging properties of the empirical results are evident in table 2. First, neither restricted specification shows evidence of autocorrelation in the error terms, either in the full or reduced sample periods. This is encouraging because the imposition of parameter restrictions such as symmetry usually tends to introduce serial correlation among the error terms. Second, there is no evidence of skewness in the residuals in either sample period, with the possible exception of the unrestricted specification when the war years are omitted. In addition, once the war years are removed from the sample, the two restricted models do not show statistical evidence of thicker tails in the error terms than occurs in the unrestricted model.[15] While there is evidence of thick tails in the joint distribution of the error terms, this is not surprising given the extensive time frame over which the model is estimated, and it does not negatively impact any of the asymptotic procedures used for estimation and inference in this model.

Two additional sets of empirical results are presented in table 2. First, results of testing for strict exogeneity of food expenditure strongly suggest that food expenditure is correlated with the conditional error terms. This conclusion is invariant to the degree of restrictiveness of the model specification and the sample period estimated. The common practice of using the price-weighted sum of (often a very small number of) quantities demanded as a right-hand-side regressor clearly is not a legitimate empirical practice. Second, I applied F-tests of the separability hypothesis to the unrestricted model in both

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[15] For example, the point estimate for the coefficient of excess kurtosis in the unrestricted model for the same sample period falls well within a 95 percent confidence interval of the corresponding estimate for the quasi-concave model. In other words, the parameter restrictions associated with symmetry and jointly with symmetry and quasi-concavity do not appear to create spurious outliers in the data.

sample periods.[16] Separability is marginally not rejected in the complete sample at the 5 percent significance level and marginally rejected at this level when the war years are excluded. Since the unrestricted model appears to be stable with the war years, but not stable when they are excluded, I tentatively conclude that separability of foods from other goods is a reasonable restriction for this data set. This issue clearly warrants further consideration, but I leave it to future work.

Table 3 reports the equation summary statistics for the fully restricted, globally quasi-concave and separable model specification for both sample periods. In this table, the average per capita expenditure levels for individual food items also are reported in constant 1967 dollars. For all commodities, eliminating the war years substantially reduces the equation standard error of the estimate, denoted by $\hat{\sigma}_{\varepsilon_i}$ in the table. This is consistent with the war years representing a separate structure in the market for food. This makes sense. For example, there were very high price supports for dairy products to encourage sufficient milk production to supply the Allied Armed Forces, as well as rationing and other quantity controls for pork and other foods. I conclude that including the war years is likely to contaminate estimation and inference in this data set.

Table 4 presents the estimated structural parameters associated with the constant terms, demographic variables, and lagged quantities consumed (i.e., myopic habit formation), with estimated asymptotic standard errors in parentheses below each parameter estimate. One notable feature in this table is that, with the exceptions of butter, cheese, poultry, and eggs, habit formation appears to be considerably weaker than previous studies of food demand suggest. This result is likely due to the inclusion of the variables associated with the age distribution and ethnic makeup of the U.S. population. These variables have changed substantially, although rather smoothly and quite nonlinearly, over time. Hence, they likely represent nonlinear trends in food consumption that previously have been tacitly proxied by lagged quantities demanded. An interesting question raised by this result is whether habit formation is present at all in these data, and if so, then is it myopic or rational?

For completeness, table 5 presents the estimated parameters associated with the negative of the inverse hessian for the food sector subutility function. Elements of this table form the lower triangle of the symmetric, positive definite matrix $\boldsymbol{C_{xx}} = \boldsymbol{LL'}$. The asymptotic standard errors in parentheses below the estimated coefficients are calculated with the delta method.

5. Food, Nutrition, and Welfare Effects of the U.S. Dairy Program

I believe that the econometric model developed and estimated above offers the potential to help us better understand U.S. food consumption in general. However, the primary

[16] Additional tests of symmetry and quasi-concavity were conducted without imposing separability, with results similar to those reported in table 2. Details of these results are available upon request.

purpose for its construction is as a tool for the analysis of the food, nutrition, and consumer welfare impacts of U.S. farm and food policy. The purpose of this section of the paper is to describe and discuss an application of this model to an analysis of the federal dairy program over the period 1949-94.

5.1 Brief History and Economic Structure of the U.S. Dairy Program

Although the federal dairy program has undergone some changes in its 50-year history, the basic structure remained essentially unchanged since 1949. There are two main components to the dairy program, a price support program and federally administered milk marketing orders. The current price support program was permanently established by the Agricultural Act of 1949. Classified pricing of fresh fluid milk in fluid milk markets under federal marketing orders was provided for by the Agricultural Marketing Agreement Act of 1937. In the Agricultural Act of 1949 and its subsequent amendments, Congress specified three guidelines for the operation of price supports: (1) Farm milk prices are to be supported at no less than 75 percent of parity.[17] (2) The Secretary of Agriculture of the U.S. Department of Agriculture (USDA) is authorized to determine the specific price support level within the range of 75 to 90 percent of parity. (3) The milk price would be supported through government purchases of milk products.

Raw milk is a bulky, perishable product. Therefore, the USDA cannot reasonably purchase fresh milk directly from farmers. Consequently, the Commodity Credit Corporation (CCC) of the USDA purchases butter, nonfat dry milk, and cheese from processors at pre-announced prices. The CCC-announced purchase prices are administered in an effort to attain the desired level of prices for manufacturing milk at the farm level. The CCC pays for the storage of the products purchased and disposes of these stocks over time through CCC sales to commercial handlers and processors at pre-announced release prices, through subsidized exports, and through donations to domestic and foreign food aid programs.

Federal milk marketing orders can be traced to the Agricultural Adjustment Acts of 1933 and 1935 and the Agricultural Marketing and Agreement Act of 1937. These orders set minimum prices that must be paid by processors to dairy farmers or their cooperatives for Grade A milk. Markets where federal orders are in place are those where producers of two-thirds of the milk marketed in the area or two-thirds of the number of producers marketing milk in the area have elected to come under a federal order. Only Grade A milk, that is, milk that meets the sanitary requirements to be legally sold as a fluid product, is regulated under federal milk marketing orders. Over 85 percent of all milk currently produced in the United States is Grade A milk, and federal marketing orders regulate over 80 percent of the Grade A milk produced in the country. State marketing orders that largely mimic the federal order program control virtually all of the remaining Grade A milk produced.

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[17] Parity is defined by an index of the cost of agricultural production based on the period 1910-14, a time when U.S. agriculture had done quite well.

Two major provisions of federal milk orders are classified pricing according to use and the pooling of all revenue from the sale of milk to obtain a single blend price to be paid to dairy farmers. Milk used for fluid products is designated Class 1. Milk used for soft manufactured products such as ice cream, cottage cheese, and yogurt is designated as Class 2. Milk used for hard manufactured products such as cheese, butter, and nonfat dry milk is classified as Class 3. All of the federal marketing orders set minimum prices on the basis of specified relationships to the price of manufacturing grade milk in the states of Minnesota and Wisconsin. The Class 3 price normally is equal to the Minnesota-Wisconsin base price. Class 1 and Class 2 prices are set at given differentials above the Class 3 price. Also, the Class 1 prices rise with increasing distance from Eau Claire, Wisconsin at a rate that generally reflects transportation costs.

The demand for fresh milk generally is not very responsive to price changes, due to the perishability and high transport costs of raw whole milk. However, the demand for manufactured milk products is more responsive to price changes because these products are less bulky and perishable, and hence more easily stored and transported. This difference in the responsiveness of demand to price is the basis for the classified pricing system in federal milk marketing orders. By raising the price of milk used for fresh products relative to the price of milk used for manufactured dairy products, producers' incomes can be increased with a blend price that is higher than the competitive, unregulated price.

The federal marketing order program and the price support program are closely related for two reasons. First, the blend price in Grade A marketing orders is significantly higher than the price for manufacturing milk, due to the classified pricing scheme. This leads to a surplus of Grade A milk production. There are no production controls, and surplus Grade A milk competes directly with Grade B milk, which can only be sold for manufactured products (Class 2 and 3 uses) due to the weaker sanitary restrictions on Grade B milk production. The competition of surplus Grade A milk with Grade B milk in the manufacturing market tends to depress the price for manufacturing milk. It also leads producers to move toward the higher cost of Grade A production, because the price difference between Grade A and Grade B milk is much greater than the slight difference in production costs to meet the tighter sanitary restrictions.[18] Second, federal milk marketing order minimum class prices are based on the Minnesota-Wisconsin price, which reflects the supply and demand situation for the entire industry. The CCC's purchases of butter, cheese, and nonfat dry milk generates a price floor for the Minnesota-Wisconsin price, which in turn produces a floor for all milk prices.

During the late 1970's and early 1980's, the U.S. dairy market grew seriously out of equilibrium, in large part due to ill-conceived policy changes. During the rapid rise in commodity prices that resulted from the OPEC oil embargo, the federal price supports

[18] It is estimated that virtually all remaining Grade B milk produced in the United States could qualify as Grade A if a market existed for the additional fluid grade milk (USDA 1984).

and marketing orders became less effective, and the support prices were no longer binding price floors by the end of 1976. As a result, the Food and Agriculture Act of 1977 raised the minimum support level to 80 percent of parity and introduced a midyear adjustment in the support price for farm milk to account for increases in the parity index during semiannual periods. As farmers responded to the higher milk prices with larger herds and consumers responded with lower purchases, the market moved quickly out of equilibrium. By 1980, the dairy market had large quantities of surplus production and the federal budgetary costs of the dairy programs began to increase rapidly. By 1980 and 1981, the situation had reached a point where the Class 3 support price was above a competitive market-equilibrium level (LaFrance and deGorter), and net USDA expenditures totaled more than $2.4 billion in 1981. Moreover, net government removals of butter accounted for nearly 75 percent of the total annual production of that product in 1981.

In response to the budgetary burdens of the large milk surpluses, the Agriculture and Food Act of 1981 departed from the traditional parity basis for supporting milk prices, and set the support price at the nominal level of $13.10 per hundred pounds (cwt) of milk sold as of September 1981. But the large surpluses continued and the Budget Reconciliation Act of 1982 froze support prices at $13.10 for two years and provided for a $0.50 per cwt assessment from milk producers' marketing receipts to partially offset the rapidly rising USDA outlays. The 1983 Dairy and Tobacco Adjustment Act lowered the support price to $12.60 in December of 1983, and provided for further reductions to $12.10 on April 1, 1985 and $11.60 on July 1, 1985 because CCC purchases continued to exceed 10 billion pounds, milk equivalent, per year (approximately 8 percent of total milk production). Since then, the five-year omnibus farm bills of 1985 and 1990 gradually lowered the support price first to $11.10 per cwt, then $10.60, and finally $10.10, where it remains at the present time. During this same period, however, the Class 1 price differentials were uniformly increased across the country by an average of just over $1.50 per cwt, so that the extent of Class 1 price discrimination in federal marketing orders is now around 50 percent of the common Class 2 and Class 3 support price.

The 1983 Dairy and Tobacco Adjustment Act also provided for a milk diversion program. For the period December 1, 1983 through March 31, 1985, a mandated assessment of $0.50 per cwt was made on all milk marketed for commercial use by U.S. producers in the 48 contiguous states. These funds were used to partially offset the USDA outlays for the dairy program. Also, producers who voluntarily elected to participate in the milk diversion program by reducing their milk marketings between 5 and 30 percent below their base period production, were paid $10 per cwt for these reductions. However, participation by dairy producers in the diversion program was very low and milk production in 1984 declined only 4 percent from the record level of 140 billion pounds produced in 1983.

As a result, one component of the 1985 omnibus farm bill was a milk production termination program, in which dairy farmers bid competitively to be paid on a per cwt basis to voluntarily cease milk production for 5 years, and slaughter or sell overseas their entire dairy herds. The goal of this "whole herd buyout" was to reduce the U.S. dairy herd by 10 percent, with an associated reduction in milk production. Participating farmers submitted bids for an amount to be paid per cwt of historical production for a period of 5 years in order to stop producing milk and sell their dairy cows and heifers either to a slaughterhouse or in other countries. Successful bids ranged from just over $10.00 to $22.50 per cwt and roughly 10 percent of the existing dairy herd was removed from the market over the period from April 1, 1986 to October 31, 1987. However, between 1980 and 1985, the number of replacement heifers in the aggregate U.S. dairy herd increased from just over 25 heifers per 100 producing milk cows to just under 50 heifers per 100 cows. The result was that total milk production actually increased by about 1.5 percent during the paid termination program, almost certainly the result of rational expectations on the part of dairy producers regarding the forthcoming dairy herd buyout program.

5.2 Modeling the Economic Impacts of the Dairy Program

To derive a counterfactual market equilibrium, I follow the modeling efforts of LaFrance and deGorter (1985) and LaFrance (1993a, 1993b). Basically, LaFrance and deGorter developed a farm-level partial equilibrium model of the U.S. dairy sector and simulated the competitive market equilibrium with no government intervention. For the competitive farm price of milk without federal programs, I simply extended their simulation of the years 1953-80 to the period 1949-1994.

In LaFrance (1993a, 1993b), I constructed a model of the farm-to-retail price linkages, including estimates of the mechanisms by which the federal government sets its wholesale purchase prices for butter, cheese, and powdered milk. I also estimated the relationship between the minimum Class 1 price support for fluid grade milk and the retail price of fresh milk products, as well as the wholesale-to-retail price relationships for the other major dairy products (butter, cheese, ice cream and frozen yogurt, and canned and nonfat dry milk). Table 6 reports the results of re-estimating these price equations for the period 1949-1994 in logarithmic form. In the table, the explanatory variables for the retail price of milk is the farm level minimum Class 1 support price, the U.S. average manufacturing wage rate, the producer price index (PPI) for materials and components, and the PPI for fuels and power. For the retail prices of butter, cheese, and canned and powdered milk the corresponding wholesale price variable is the pre-announced government purchase price of butter, cheese, and nonfat dry milk, respectively. The manufacturing wage rate and the PPIs for materials and components and for fuels and power are common to all price equations. Since ice cream is not directly supported by the government program, I used both the minimum Class 1 and Class 2 prices as well as the nonfood CPI in addition to the measures of producer costs at the wholesale/manufacturing level. Also in the table, ρ_1 and ρ_2 denote the coefficients for first- and second-order autocorrelation, R^2 is the coefficient of multiple determination, and DW is the Durbin-Watson test for remaining serial correlation in the error terms.

Once the farm-level competitive price and a set of farm-to-retail price linkage equations have been obtained, a set of counterfactual retail prices for each year are generated simply by replacing the farm-level minimum Class 1 and Class 2 support prices with the competitive price in the equations for the wholesale government purchase prices and in the retail milk and ice cream price equations. In each year, this generates a set of in the retail milk and ice cream price equations. In each year, this generates a set of prices for the government purchases that would be consistent with the goal of a stable market equilibrium but none of the attendant social costs of price discrimination and government purchases and storage of excess commodities. Obviously, though, it does not tell us how the market would actually function absent any government involvement. The primary insights we can obtain from this analysis relate to the consumer welfare costs and the food consumption and nutritional effects of price discrimination due to the U.S. dairy program.

Figure 2 depicts the simulated effects on retail quantities and prices of the dairy program over the period 1949-1994. The average percentage increase in the price of fresh milk at the retail level over this period appears to have been about 6.5 percent. Similarly, the average decreases in the prices of butter, cheese, ice cream, and canned and powdered milk were about 4.6 percent, 3.8 percent, 2.6 percent, and 1.9 percent, respectively. The resulting average decrease in the quantity consumed of fresh milk was about 2.5 percent, while the average increases in the consumption of butter, frozen dairy products, and canned and powdered milk were 3.3 percent, 7.7 percent, and 5.8 percent, respectively. Cheese consumption appears to be little effected by the federal controls.

By combining the nutrient content matrix in table 1 with a set of simulated changes in all foods consumed, it is straightforward to construct estimates of the impacts on nutritional consumption resulting from the federal dairy program. I generated these changes for each year over the period 1949-94. As we might expect, given the small percentage changes in food consumption, the nutritional implications of the dairy program appear to be relatively minor. With the dairy program in place, average percentage changes in nutrient consumption levels are as follows: calories and protein increase by less than 0.2%; fat decreases by 0.1%; carbohydrates increase by 0.8%; calcium decreases by 0.3%; phosphorous is unchanged; iron increases by 0.4%; magnesium increases by 0.2%; vitamin A decreases by 0.6%; thiamin decreases by 0.5%; riboflavin decreases by less than 0.1%; niacin increases by 0.5%; vitamin B_6 increases by less than 0.1%; vitamin B_{12} decreases by 1.1%; vitamin C decreases by 0.2%; zinc is unchanged; and cholesterol decreases by 1.0%. These orders of magnitude are sufficiently small that, for all practical purposes, the nutritional consequences of the dairy program may not be a primary concern in the ongoing policy debate.

It is also straightforward to estimate the consumer welfare effects of the relative price distortions created by the dairy program. The compensating variation measure is based on equation (3.3) above, or alternatively we could use the equivalent variation measure, which in this case is defined by

$$(5.1) \qquad ev(\boldsymbol{p}_x^0, \boldsymbol{p}_x^1, m, \boldsymbol{s}) = -\left(\boldsymbol{m} - \alpha_x(\boldsymbol{s})'\boldsymbol{p}_x^0 - \alpha_y(\boldsymbol{s})\pi(\boldsymbol{p}_y)\right)$$

$$+\left(\boldsymbol{m} - \alpha_x(\boldsymbol{s})'\boldsymbol{p}_x^1 - \alpha_y(\boldsymbol{s})\pi(\boldsymbol{p}_y)\right) \cdot \sqrt{\frac{(\boldsymbol{p}_x^0)'\boldsymbol{C}_{xx}\boldsymbol{p}_x^0 + 2(\boldsymbol{p}_x^0)'\gamma_{xy}\pi(\boldsymbol{p}_y) + \gamma_{yy}\pi(\boldsymbol{p}_y)^2}{(\boldsymbol{p}_x^1)'\boldsymbol{C}_{xx}\boldsymbol{p}_x^1 + 2(\boldsymbol{p}_x^1)'\gamma_{xy}\pi(\boldsymbol{p}_y) + \gamma_{yy}\pi(\boldsymbol{p}_y)^2}}$$

$$= \sqrt{\frac{(\boldsymbol{p}_x^0)'\boldsymbol{C}_{xx}\boldsymbol{p}_x^0 + 2(\boldsymbol{p}_x^0)'\gamma_{xy}\pi(\boldsymbol{p}_y) + \gamma_{yy}\pi(\boldsymbol{p}_y)^2}{(\boldsymbol{p}_x^1)'\boldsymbol{C}_{xx}\boldsymbol{p}_x^1 + 2(\boldsymbol{p}_x^1)'\gamma_{xy}\pi(\boldsymbol{p}_y) + \gamma_{yy}\pi(\boldsymbol{p}_y)^2}} \cdot cv(\boldsymbol{p}_x^0, \boldsymbol{p}_x^1, m, \boldsymbol{s}).$$

In each year, the two welfare measures are within 0.1% of each another in the present problem, so I concentrate on compensating variation. Figure 3 depicts the total U.S. consumer welfare costs of the dairy market price distortions in millions of 1995 dollars per year, obtained by multiplying the per capita measure by the U.S. population in each year. The compensating variation measure of consumer losses is quite volatile, as is the corresponding equivalent variation measure. The notable exceptions to positive losses to consumers are 1977 and 1978, when the price supports were not binding during the period of rapid commodity price inflation following the OPEC embargo, and the period from 1989 through 1994. In this latter period, the price support level for manufacturing grade milk has fallen in real terms to a very low level and consumers may now be net beneficiaries of the excess supply problems that plagued the dairy industry throughout most of the 1980's. Nevertheless, the average annual compensating variation for the relative price distortions over the 46-year period from 1949-94 is just over one billion dollars (in 1995 dollars, using the nonfood CPI to adjust for inflation). In 1980, this economic cost approached 5 billion 1995 dollars per year.

6. Conclusions

This paper presents results on an econometric model of per capita food consumption and nutritional intake for the United States. The model is fully consistent with economic theory, the consumption of foods for both nutrition and taste, and strict aggregation across income, demographic factors, and varying micro-parameters. Explicit parametric solutions for the global imposition of the necessary and sufficient conditions for weak integrability were derived and implemented. The empirical application estimates a system of demands for twenty-one food items using annual U.S. per capita time series data for 1918-1994. The empirical properties of the model overall are quite encouraging. Results of the hypothesis tests of the restrictions required for economic theory suggest that these conditions are readily accommodated by this data set and this model structure. This is an encouraging result given the restrictive nature of the model structure that is required for strict aggregation. This model offers a reasonable, coherent approach to studying the first-order consumer effects of changes in farm and food policies in the United States. Finally, the paper reports an application of the model to an analysis of the consumption, nutrition, and consumer welfare effects of the U.S. dairy program over the period 1949-94. Due to the relatively inelastic nature of the demand functions for dairy products, the quantity effects, and hence the nutritional effects, of the federal dairy program are small. But the consumer costs of this program are not.

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Figure 1. Model Stability Tests: U.S. Food Demand.

(Separable Model, 1919-1994)

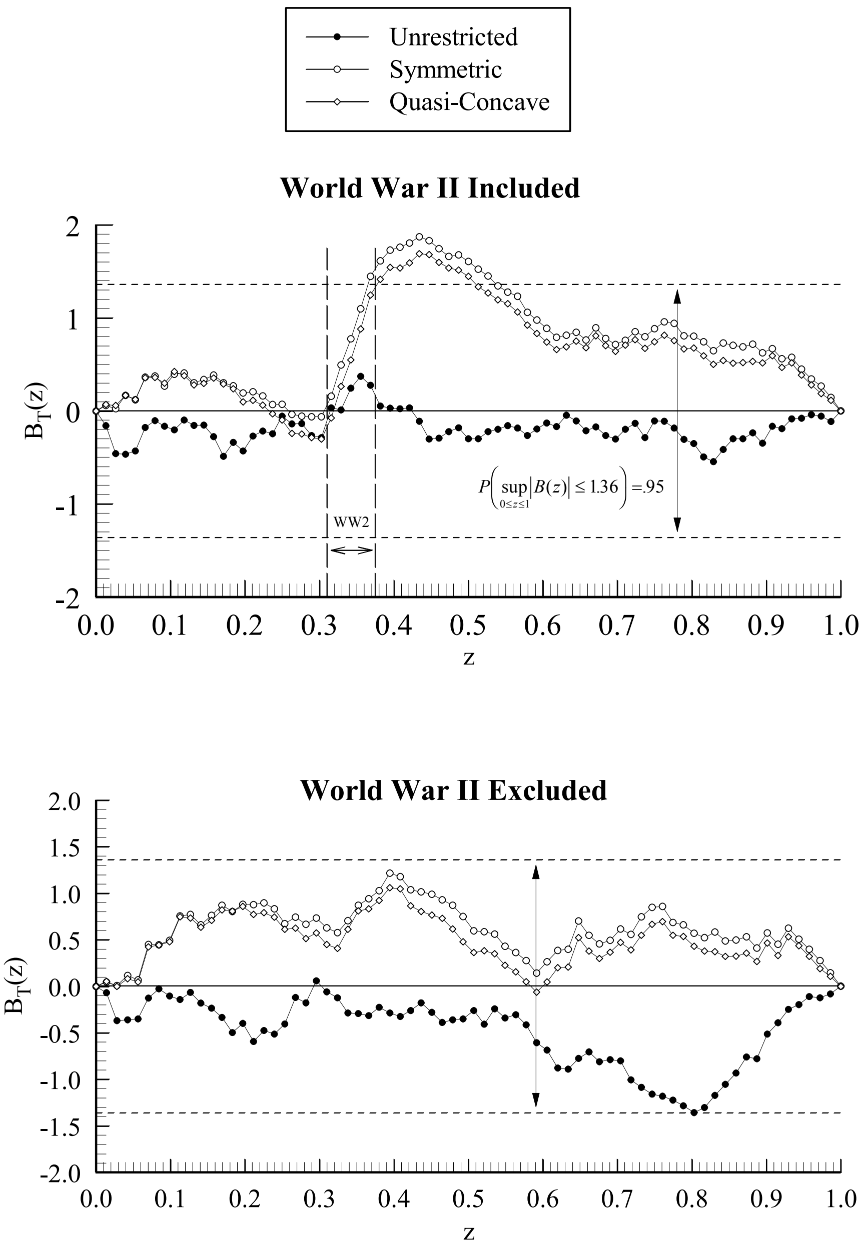

Figure 2. Changes in Dairy Product Demands due to the Dairy Program, 1949-1994.

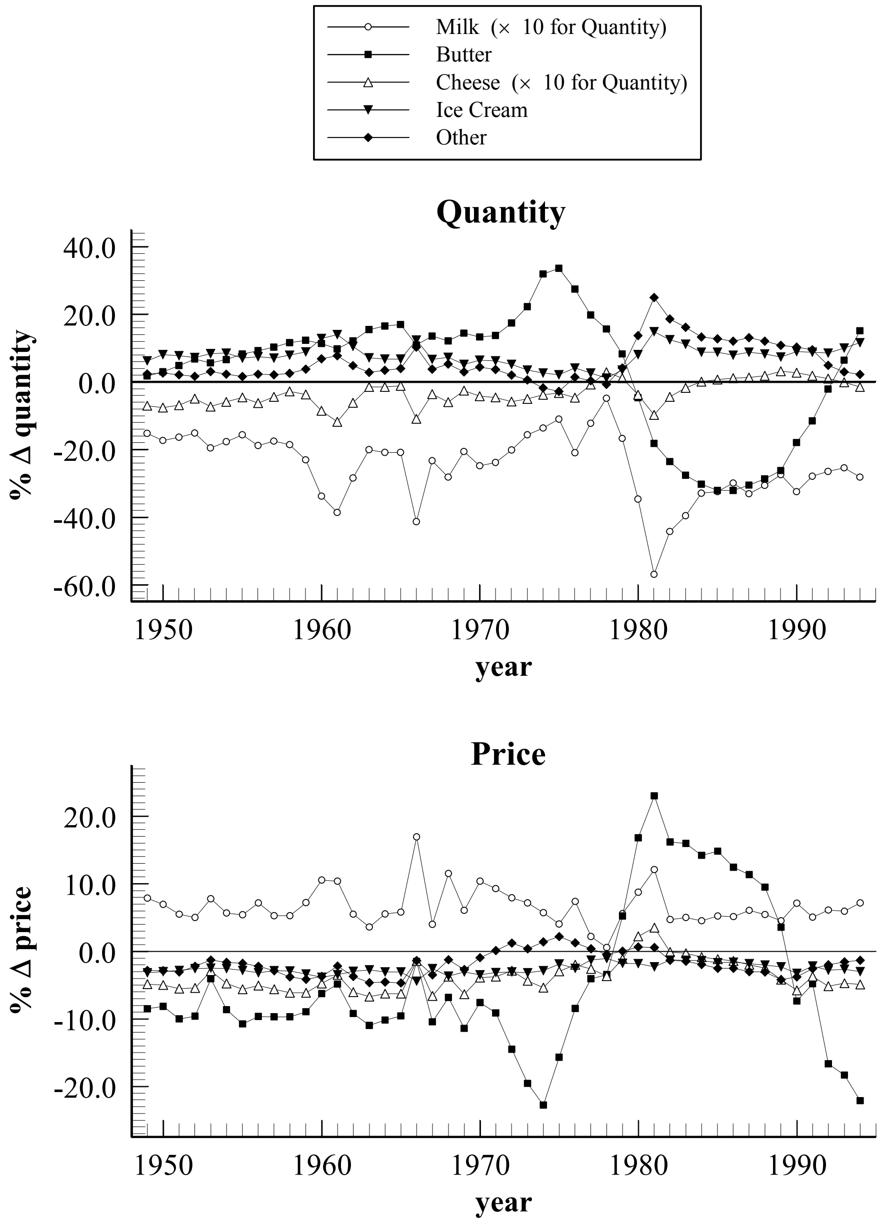

Figure 3. Compensating Variation: 1949-1994.

(millions of 1995 $ per year)

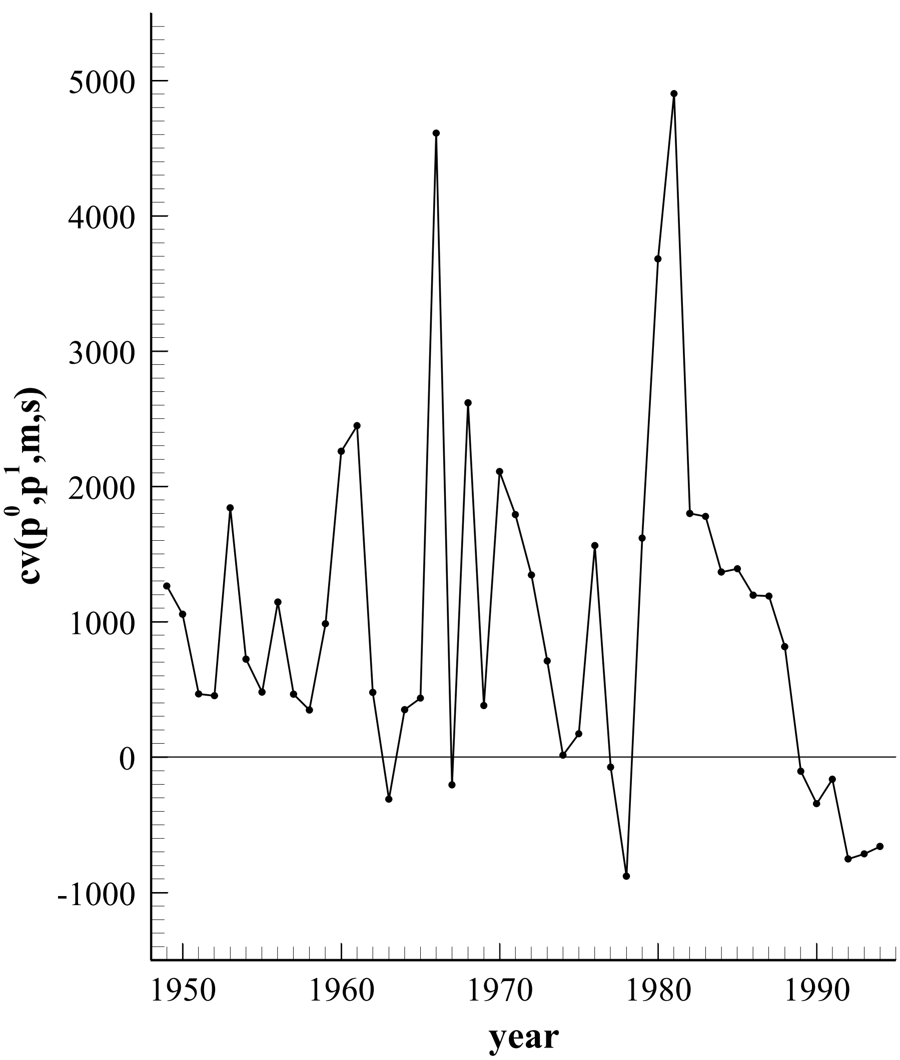

Table 1. Nutrient Content of U.S. Foods (Nutritional Units per Pound of Food, 1948-83 averages).

	Calories	Protein	Fat	Carbos	Calcium	Phosph	Iron	Magnes	A	Thiamin	Riboflav	Niacin	B_6	B_{12}	C	Zinc	Choles
	kilo-cals	grams	grams	grams	mil grm	mil grm	mil grm	mil grm	ret	mil grm	mil grm	mil grm	mil grm	mil grm	mil grm	mil grm	mil grm
Milk	280.2	14.92	14.92	22.20	522.6	426.6	0.26	60.9	100.2	0.179	0.753	0.38	0.195	1.72	4.4	1.79	55.9
Butter	3259.4	4.44	369.15	6.10	110.5	110.3	1.03	9.7	2529.7	0.034	0.162	0.37	0.031	0.60	0.0	0.24	930.0
Cheese	1348.0	102.86	98.69	9.65	2311.3	1765.3	2.27	95.7	589.2	0.117	1.395	0.54	0.355	4.25	0.0	10.66	294.4
Ice Cream	450.9	15.20	33.72	21.25	511.5	408.0	0.22	47.4	222.3	0.187	0.671	0.31	0.155	1.85	2.9	1.76	80.7
Canned & Dry Milk	899.1	63.11	19.01	119.88	2306.0	2013.0	1.36	262.3	144.2	1.535	3.634	2.07	0.769	6.13	14.0	7.08	84.0
Beef & Veal	1053.3	75.72	80.79	0.00	43.1	700.6	11.27	73.5	27.0	0.351	0.695	18.17	1.378	6.24	0.0	14.05	232.6
Pork	1972.6	59.47	189.12	2.49	42.3	674.4	8.54	72.1	0.0	2.919	0.720	14.81	1.260	5.05	0.0	6.34	264.1
Other Red Meat	840.8	93.78	47.34	8.78	120.4	966.8	13.10	82.9	6757.2	0.892	4.530	30.12	1.828	90.36	31.7	16.34	808.2
Fish	863.3	112.74	39.40	3.02	310.7	1307.4	6.92	164.0	121.7	0.331	0.685	30.68	1.839	27.72	3.5	12.69	363.2
Poultry	648.8	61.48	42.52	0.88	37.5	520.1	4.52	63.3	465.0	0.237	0.745	19.11	1.199	4.51	7.4	5.52	272.4
Fresh Citrus Fruit	108.9	1.84	0.47	27.08	53.8	44.1	0.82	31.2	75.7	0.194	0.064	0.71	0.179	0.00	112.2	0.38	0.0
Other Fresh Fruit	251.5	2.89	1.81	64.16	39.3	74.4	2.03	62.3	459.8	0.171	0.181	1.93	0.751	0.00	42.6	0.49	0.0
Fresh Vegetables	177.3	8.53	1.20	38.94	155.2	212.5	3.50	111.5	1344.1	0.390	0.339	3.06	0.680	0.00	121.0	1.30	0.0
Potatoes	331.8	8.25	0.66	73.45	34.7	196.4	2.41	83.1	328.7	0.385	0.138	5.44	0.835	0.00	62.6	1.54	0.0
Processed Fruit	227.7	8.42	0.99	52.63	89.7	170.3	4.84	74.7	956.2	0.325	0.243	3.92	0.556	0.00	64.4	1.48	0.0
Proc. Vegetables	713.4	34.31	29.60	90.93	210.8	620.0	9.98	257.4	825.3	0.891	0.435	11.11	0.998	0.00	54.6	4.20	0.0
Fats & Oils	3834.0	0.94	429.36	0.78	0.0	0.0	0.00	0.0	543.9	0.000	0.000	0.00	0.000	0.00	0.0	0.13	101.4
Eggs	634.4	49.57	44.38	3.67	211.4	760.6	7.97	42.5	631.3	0.401	1.298	0.33	0.475	7.30	0.0	4.69	1964.1
Cereal	1705.3	47.09	5.57	361.73	81.1	495.8	12.45	151.4	14.3	1.987	1.083	14.95	0.471	0.04	1.9	3.96	0.0
Sugar	1684.0	0.08	0.00	441.08	9.7	3.8	0.58	3.2	0.0	0.003	0.010	0.03	0.004	0.00	0.1	0.06	0.0
Coffee, Tea & Cocoa	497.4	10.31	9.67	29.25	101.6	383.5	6.58	307.1	68.74	0.03	0.28	3.84	0.0	0.00	0.0	0.10	0.0

Table 2. Model Diagnostics: U.S. Food Demands, 1919-1994.

	With World War II			Without World War II		
	UNR	SYM	Q-C	UNR	SYM	Q-C
s(S)	1515.9	1361.9	1321.2	1415.7	1228.1	1250.0
ρ	-.124	-.044	-.0053	-.135	-.039	-.028
σ_ρ	.026	.027	.028	.027	.029	.029
t_ρ	4.78	1.61	0.19	5.02	1.35	0.97
η_3	.070	.0083	.012	.147	.045	.067
σ_{η_3}	.061	.061	.061	.063	.063	.063
t_{η_3}	1.14	0.14	0.19	2.31	0.71	1.05
η_4	.200	.658	.637	.451	.675	.609
σ_{η_4}	.123	.123	.123	.127	.127	.127
t_{η_4}	1.63	5.37	5.19	3.55	5.32	4.80
J-B $\chi^2(2)$	3.94	28.84	26.98	17.96	28.80	24.15
P-value	0.14	5.5×10^{-7}	1.4×10^{-6}	1.3×10^{-4}	5.6×10^{-7}	5.7×10^{-6}
Expenditure Exogeneity Tests						
$\bar{\hat{u}}$	1.696	3.398	3.506	1.624	5.150	5.061
$\sigma_{\bar{\hat{u}}}$	.340	1.016	1.001	.343	1.364	1.332
$t_{\bar{\hat{u}}}$	4.986	3.344	3.503	4.739	3.776	3.800
P-value	3.1×10^{-6}	4.1×10^{-4}	2.3×10^{-4}	1.1×10^{-5}	8.0×10^{-5}	7.2×10^{-5}
F-Tests						
Separability	1.55			1.60		
P-value	.053			.043		
Theory		1.09	1.84		1.18	1.12
P-value		.202	9.7×10^{-11}		.058	.122
Model Stability Tests						
$\max\lvert B_T(z)\rvert$	.55	1.87	1.69	1.39	1.22	1.01
P-value	.92	.0018	.0066	.042	.10	.26

Notes: UNR, SYM, and Q-C are the unrestricted, symmetric, and quasi-concave specifications, respectively; s(S) is the residual sum of squares criterion at the second iteration; ρ is the common first order autocorrelation coefficient; η_3 is the coefficient of skewness; η_4 is the coefficient of excess kurtosis; and J-B $\chi^2(2)$ is the Jarque-Bera test for normality of the standardized residuals.

Table 3. Equation Summary Statistics: U.S. Food Demands, 1919-1994.
(Separable, Globally Quasi-concave Model)

	With World War II			Without World War II		
	Avg. Expd.	R^2	$\hat{\sigma}_{\varepsilon_i}$	Avg. Expd.	R^2	$\hat{\sigma}_{\varepsilon_i}$
Fresh Milk and Cream	36.25	.9953	.5524	35.44	.9973	.4002
Butter	8.36	.9914	.4789	8.16	.9965	.3118
Cheese	11.39	.9952	.5102	11.80	.9983	.3056
Frozen Dairy Products	4.26	.9581	.2534	4.21	.9879	.1348
Other Dairy Products	3.38	.9141	.3035	3.28	.9866	.1097
Beef and Veal	66.62	.9885	2.681	68.14	.9950	1.759
Pork	35.34	.9520	1.543	35.09	.9749	1.119
Other Red Meat	10.20	.9566	.4523	10.02	.9588	.4331
Fish	7.90	.9883	.3794	8.07	.9949	.2552
Poultry	15.94	.9746	.6933	15.73	.9893	.4547
Fresh Citrus Fruit	4.75	.8256	.4688	4.52	.6712	.3851
Fresh Non-citrus Fruit	11.81	.9034	1.303	11.75	.9486	.9652
Fresh Vegetables	16.30	.9868	.4237	16.22	.9817	.4141
Potatoes	8.24	.9367	.4232	8.13	.9631	.3219
Processed Fruit	23.24	.9824	1.564	23.82	.9883	1.293
Processed Vegetables	11.34	.9717	.4536	11.26	.9890	.2886
Fats and Oils	13.50	.9603	.4284	13.49	.9738	.3582
Eggs	16.49	.9951	.5282	15.81	.9989	.2428
Flour and Cereals	19.63	.9668	.5454	19.73	.9889	.3229
Sugar	25.45	.9780	.9061	25.91	.9877	.6720
Coffee, Tea, and Cocoa	12.35	.9694	.6043	12.54	.9803	.4819
Total Food Expenditure	362.7	.9902	5.324	363.2	.9925	4.802

Table 4. Estimated Intercepts, Demographics, and Habit Coefficients.

(Separable, Globally Quasi-concave Model: War II Excluded)

		Age Distribution Variables			Ethnicity Variables		
	Constant	**Mean**	**Variance**	**Skewness**	**Black**	**Other**	x_{t-1}
Fresh Milk and Cream	371.60 (79.04)	-2.252 (2.425)	3.305 (0.670)	-.7544 (.7276)	-20.10 (13.46)	-3.754 (9.115)	.3676 (.0576)
Butter	4.33 (13.60)	.0070 (.2593)	-.2890 (.0917)	-.0308 (.0790)	1.340 (1.927)	-2.249 (1.163)	.7446 (.0841)
Cheese	-16.37 (11.58)	.6105 (.3335)	-.1206 (.0796)	.0807 (.0848)	.271 (1.881)	3.062 (1.323)	.5028 (.1089)
Ice Cream and Frozen Yogurt	-40.47 (27.81)	.0594 (.7583)	.8219 (.2795)	.0363 (.1979)	1.014 (4.249)	.7419 (2.685)	.3905 (.1215)
Canned and Powdered Milk	33.82 (24.15)	-.2162 (.7751)	1.081 (.2870)	-.4834 (.1816)	-3.800 (4.511)	.7849 (2.492)	.3202 (.1345)
Beef and Veal	-378.30 (29.39)	1.801 (.8713)	1.848 (.2137)	-.0269 (.2435)	31.87 (5.107)	-21.43 (3.398)	.0221 (.0471)
Pork	149.79 (27.03)	1.029 (.8692)	.9235 (.2282)	.1426 (.2409)	-16.47 (4.598)	5.214 (3.284)	.0751 (.0396)
Other Red Meat	27.51 (13.09)	.1148 (.4227)	-.0170 (.1134)	.0954 (.1125)	-1.889 (2.430)	.0164 (1.565)	.0740 (.1248)
Fish	43.48 (12.23)	.3116 (.3377)	-.2024 (.0817)	.1578 (.0913)	-4.403 (2.002)	5.746 (1.354)	.2595 (.0855)
Poultry	31.11 (20.58)	.0801 (.5246)	.2411 (.1637)	.0719 (.1448)	-3.901 (3.328)	12.93 (2.865)	.5049 (.0751)
Fresh Citrus Fruit	69.47 (40.46)	6.619 (1.458)	-.3149 (.3077)	.1313 (.3374)	-22.87 (7.547)	6.362 (4.777)	-.0489 (.0936)
Fresh Non-Citrus Fruit	1061.11 (97.25)	-4.292 (2.489)	-4.102 (.6853)	.6095 (.6761)	-68.18 (15.16)	60.00 (10.55)	-.4822 (.0757)

Numbers in parentheses are asymptotic standard errors.

Table 4. Estimated Intercepts, Demographics, and Habit Coefficients, Continued.

(Separable, Globally Quasi-concave Model: War II Excluded)

	Constant	Age Distribution Variables Mean	Variance	Skewness	Ethnicity Variables Black	Other	x_{t-1}
Fresh Vegetables	299.99 (50.40)	7.032 (1.555)	.3073 (.3879)	1.507 (.3864)	-45.64 (8.970)	33.91 (5.939)	.1833 (.0928)
Potatoes	579.74 (100.17	-9.816 (2.854)	-2.187 (.6830)	.0082 (.7347)	-6.228 (16.05)	18.81 (10.16)	-.0178 (.0954)
Processed Fruit	-210.61 (64.66)	3.131 (1.743)	1.251 (.4555)	.2825 (.3431)	7.189 (7.102)	2.235 (5.876)	.2771 (.0732)
Processed Vegetables	39.44 (44.08)	7.199 (1.459)	-.3394 (.3400)	1.847 (.3610)	-29.33 (7.343)	21.13 (4.820)	.3110 (.0680)
Fats and Oils	22.44 (24.38)	3.356 (.7099)	-.2798 (.1852)	.9162 (.1957)	-13.02 (4.090)	15.80 (3.010)	.2167 (.0793)
Eggs	54.41 (41.36)	-.7551 (1.335)	.3849 (.3513)	-.1594 (.2345)	-2.559 (4.416)	-.38687 (3.773)	.7194 (.0718)
Flour and Cereals	1085.14 (125.96)	-9.148 (2.637)	-4.559 (.7171)	.2412 (.6415)	-49.30 (13.78)	52.05 (9.453)	.2762 (.0884)
Sugar	185.19 (54.41)	6.726 (1.748)	-2.415 (.3695)	1.812 (.5018)	-26.31 (10.24)	24.14 (6.912)	.0412 (.0593)
Coffee, Tea, and Cocoa	22.13 (9.034)	.7609 (.3029)	.2149 (.0715)	-.0033 (.0786)	-4.142 (1.671)	1.588 (1.111)	.2128 (.0599)
Nonfood Expenditure	-4069.95 (1229.15)	322.42 (38.09)	13.38 (11.37)	89.91 (9.802)	-916.74 (185.04)	1274.47 (139.32)	

Numbers in parentheses are asymptotic standard errors.

Table 5. Negative Inverse Hessian, Food Subutility Function: Separable, Globally Quasi-concave Model, World War II Excluded.

	Milk and Cream	Butter	Cheese	Frozen Dairy	Other Dairy	Beef and Veal	Pork	Other Meat	Fish	Poultry
Milk and Cream	.722 (.128)									
Butter	.00689 (.00644)	.00509 (.00103)								
Cheese	.00100 (.0101)	-.00090 (.00086)	.00432 (.00158)							
Frozen Dairy	-.0641 (.0390)	-.00253 (.00281)	-.00653 (.00471)	.145 (.0303)						
Other Dairy	-.118 (.0400)	-.00456 (.00288)	-.00676 (.00412)	.00998 (.0188)	.0718 (.0241)					
Beef and Veal	-.0283 (.0104)	.00252 (.00096)	-.00361 (.00135)	.00678 (.00403)	.0109 (.00369)	.0620 (.00566)				
Pork	.0102 (.0127)	-.00182 (.00140)	-.00500 (.00185)	.00728 (.00508)	.00779 (.00470)	-.0195 (.00313)	.0908 (.00837)			
Other Meat	.0313 (.0162)	-.00121 (.00109)	-.00080 (.00186)	-.00985 (.00705)	-.0142 (.00652)	-.0148 (.00238)	-.0107 (.00253)	.0379 (.00513)		
Fish	.0217 (.0087)	-.00277 (.00081)	.00417 (.00112)	-.00182 (.00430)	-.00538 (.00384)	-.00325 (.00125)	-.00691 (.00177)	-.00043 (.00176)	.00634 (.00142)	
Poultry	-.0674 (.0138)	.00030 (.00139)	.00260 (.00188)	-.01330 (.00545)	.00188 (.00563)	-.00546 (.00211)	-.00441 (.00311)	.00205 (.00268)	.00202 (.00365)	.0210 (.00346)

Table 5. Negative Inverse Hessian, Food Subutility Function: Separable, Globally Quasi-concave Model, World War II Excluded, Cont.

	Milk and Cream	Butter	Cheese	Frozen Dairy	Other Dairy	Beef and Veal	Pork	Other Meat	Fish	Poultry
Fresh Citrus Fruit	.00699 (.0213)	-.00035 (.00252)	.00076 (.00257)	-.00111 (.00756)	-.00046 (.00694)	-.00169 (.00470)	-.00866 (.00609)	.00108 (.00330)	.00100 (.00209)	-.00980 (.00563)
Fresh Non-Citrus Fruit	0.0978 (.0502)	-.0111 (.00492)	.00373 (.00694)	-.0714 (.0241)	.0207 (.0190)	-.0194 (.00992)	-.0248 (.0131)	-.0235 (.00969)	.00803 (.00656)	.0210 (.00954)
Fresh Vegetables	.0436 (.0367)	-.0107 (.00340)	.00017 (.00433)	.0267 (.0164)	.0153 (.0152)	-.00154 (.00579)	.0127 (.00787)	-.00906 (.00800)	.00786 (.00415)	-.00419 (.00651)
Potatoes	-.0235 (.0488)	.00868 (.00534)	-.00264 (.00705)	-.0194 (.0163)	.0138 (.0168)	-.00659 (.00960)	-.00285 (.0126)	.00126 (.00799)	-.00771 (.00648)	-.0120 (.0119)
Processed Fruit	-.00823 (.0160)	-.00416 (.00162)	.00032 (.00222)	-.0205 (.00705)	-.00309 (.00606)	.00132 (.00345)	.00682 (.00436)	-.00558 (.00278)	-.00045 (.00202)	-.00557 (.00335)
Processed Vegetables	.0160 (.0406)	.00409 (.00300)	.00203 (.00462)	.0447 (.0187)	-.00455 (.0176)	-.0241 (.00511)	-.0216 (.00653)	.0121 (.00752)	.0101 (.00421)	.0206 (.00643)
Fats and Oils	.0174 (.0170)	-.00322 (.00157)	.00908 (.00212)	-.0210 (.00893)	-.0173 (.00730)	-.0147 (.00284)	-.0110 (.00380)	.00766 (.00330)	.00886 (.00206)	.00770 (.00303)
Eggs	.0297 (.0146)	.00050 (.00129)	-.00169 (.00188)	-.00736 (.00602)	.00442 (.00607)	-.00125 (.00175)	.00478 (.00279)	-.00677 (.00271)	-.00318 (.00165)	-.00962 (.00260)
Flour and Cereals	-.256 (.0993)	.00138 (.00663)	-.00298 (.00913)	.0284 (.0407)	.0176 (.0361)	-.0297 (.0108)	-.0222 (.0136)	.00458 (.0172)	-.00732 (.00854)	.0361 (.0134)
Sugar	-.0332 (.0271)	.00751 (.00270)	.00349 (.00354)	.00426 (.00784)	.0187 (.00817)	-.0180 (.00545)	-.0100 (.00666)	.00536 (.00482)	-.00164 (.00351)	.0154 (.00538)
Coffee, Tea and Cocoa	-.00407 (.00289)	-.00024 (.00035)	-.00076 (.00039)	.00227 (.00113)	.00181 (.00102)	.00016 (.00075)	.00182 (.00092)	.00030 (.000486)	-.00052 (.00036)	.00021 (.00096)

Table 5. Negative Inverse Hessian, Food Subutility Function: Separable, Globally Quasi-concave Model, World War II Excluded, Cont.

	Fresh Citrus Fruits	Fresh Noncitrus Fruits	Fresh Vegetables	Potatoes	Processsed Fruit	Processed Vegetables	Fats and Oils	Eggs	Flour and Cereals	Sugar	Coffee, Tea, and Cocoa
Fresh Citrus Fruit	.0454 (.0107)										
Fresh Non-Citrus Fruit	.00353 (.0211)	.262 (.0587)									
Fresh Vegetables	-.0172 (.0109)	.0201 (.0289)	.0625 (.0209)								
Potatoes	.0303 (.0221)	.0503 (.0498)	-.0517 (.0326)	.368 (.0757)							
Processed Fruit	.0200 (.00789)	.0207 (.0157)	-.0168 (.00953)	.0156 (.0173)	.0419 (.00810)						
Processed Vegetables	-.0117 (.00955)	-.00482 (.0242)	.00818 (.0170)	-.0137 (.0233)	-.0286 (.00902)	.124 (.0271)					
Fats and Oils	.00429 (.00576)	.0179 (.0125)	-.00270 (.00836)	.00886 (.0126)	.00631 (.00390)	.00501 (.00865)	.0244 (.00567)				
Eggs	-.000677 (.00481)	.0225 (.00900)	.0103 (.00689)	-.0296 (.00955)	.00139 (.00304)	-.0212 (.00658)	-.00552 (.00305)	.0246 (.00399)			
Flour and Cereals	-.0173 (.0188)	.1607 (.0504)	-.00140 (.0334)	-.00245 (.0503)	-.0221 (.0170)	.0419 (.0380)	.00408 (.0167)	.0204 (.0148)	.282 (.112)		
Sugar	-.0154 (.0110)	-.0549 (.0220)	-.00508 (.0161)	-.0530 (.0267)	-.0128 (.00891)	.0395 (.0122)	.00684 (.00542)	.00809 (.00422)	.0237 (.0276)	.120 (.0196)	
Coffee, Tea and Cocoa	.00536 (.00190)	-.00451 (.00364)	.00366 (.00195)	.00524 (.00356)	-.00074 (.00141)	-.00294 (.00134)	-.00089 (.00085)	-.00130 (.000641)	-.00420 (.00283)	.00047 (.00181)	.00418 (.00052)

Table 6. Logarithmic Retail and Government Purchase Price Equations, 1949-1994.

	Const.	Min. Class 1 Price	Min. Class 2 Price	Govt. Butter Price	Govt. Cheese Price	Govt. Powder Milk Price	Manuf. Wage	Mtls. PPI	Fuel PPI	Non-food CPI	ρ_1	ρ_2	R^2	DW
<u>**Retail Prices:**</u>														
Milk	-2.938	.4432					.1115	.2166	.0934		1.41	-.48	.999	1.88
	(.110)	(.0413)					(.0771)	(.1195)	(.0450)		(.017)	(.017)		
Butter	-2.567			.5393			.1126	.4632	-.0813		.75		.996	2.10
	(.228)			(.0505)			(.0905)	(.1836)	(.0753)		(.098)			
Cheese	-1.374				.2258		.3432	.4098	.0491		.86		.998	1.73
	(.359)				(.1030)		(.1230)	(.2147)	(.0883)		(.075)			
Ice Cream	-0.8161	-.1205	.0529				-.6961	.1786	.1048	1.287	.62		.999	1.71
	(.0977)	(.0680)	(.0588)				(.0905)	(.1836)	(.0753)	(.092)	(.115)			
Canned Milk	-2.261					.1301	.0688	.7412	.0363		.96	-.33	.998	2.13
	(.138)					(.0520)	(.0946)	(.1791)	(.0774)		(.019)	(.019)		
<u>**Government Wholesale Purchase Prices:**</u>														
Butter	3.294		.7318				-.1054	.0612	-.1442		1.49	-.59	.981	2.36
	(.264)		(.1493)				(.2061)	(.3188)	(.1221)		(.014)	(.014)		
Cheese	2.190		.9555				.2784	-.3303	.0343		.63	-.31	.999	1.81
	(.045)		(.0327)				(.0383)	(.0882)	(.0444)		(.020)	(.020)		
Powdered Milk	1.454		.7923				.4495	-.1030	.1557		1.21	-.33	.997	2.20
	(.242)		(.1442)				(.1868)	(.3194)	(.1261)		(.019)	(.019)		