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Advertising, Structural Change, and U.S. Non-Alcoholic Drink Demand

by

Hui Xiao
Henry W. Kinnucan
and
Harry M. Kaiser

Department of Agricultural, Resource, and Managerial Economics
College of Agriculture and Life Sciences
Cornell University, Ithaca, New York 14853

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Abstract

The dominant pattern in U.S. non-alcoholic drink consumption over the past 25 years has been a steady increase in per capita soft-drink consumption, largely at the expense of coffee (and to a lesser extent) milk consumption. Our findings suggest that the major factor governing this pattern is structural change. Specifically, trend was found to be statistically significant in three of the four equations estimated in the Rotterdam system. Moreover, the estimated trend-related changes in per capita consumption (-1.0 percent per year for milk, 2.1 percent for soft drinks, and -3.7 percent for coffee and tea) leave at most 28 percent of the observed quantity variation for 1990-1994 to be accounted for by changes in relative prices, income, and advertising. Advertising effects are statistically significant, but modest. The question of whether milk advertising is profitable when demand interrelationships are taken into account must await additional research.

Key words: advertising, beverage demand, milk consumption, structural change.

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Introduction

Galbraith's hypothesis "If advertising affects the distribution of demand between sellers of a particular product, it must also be supposed that it affects the distribution between products" (Galbraith, p. 205; see also pp. 214-215) assumes added significance in the context of non-alcoholic beverage advertising. At $1.1 billion in 1994 alone (Appendix Table 2), this group is one of the most heavily advertised in the U.S. economy. Moreover, two items in the group -- milk and fruit juices -- are the target of significant levels of generic advertising (over $100 million in 1994) funded by the dairy industry and citrus growers. Another $114 million now exists for the milk moustache print campaign by milk processors (USDA, AMS). Although substantial research has been done to determine whether generic advertising of milk and fruit juices pays (e.g., Ward and Dixon; Blisard *et al.*; Kaiser *et al.*; Wohlgenant and Clary; Lee and Brown), no study has investigated the beverages in an integrated framework that takes into account the full array of substitution effects.[2] For example, a successful fluid milk advertising campaign might erode the demand and price for citrus products. In addition, the decrease in citrus price could lower the milk price through second-round or feedback effects. These spillover and feedback effects have not been addressed in the milk and citrus advertising literatures, which could cause the estimated returns to be overstated (Kinnucan, 1996).

In this paper, we determine whether advertising of non-alcoholic beverages has any detectable effect on aggregate demand for the individual beverages. Owing to the importance of demand interrelationships, special attention is given to spillover effects, i.e., whether one beverage's advertising affects the demand for related beverages. A secondary objective is to test Theil's theory that advertising elasticities are proportional to price elasticities. Theil's proportionality hypothesis has been maintained in synthetic models of advertising effectiveness (Wohlgenant) and in econometric estimation (e.g., Duffy 1987, 1990; Selvanathan 1989a; Green, Carman and McManus). As a by-product of our analysis, we test whether structural change plays a role in the observed consumption pattern, particularly the rise in soft-drink consumption between 1970 and 1994 from 24.3 gallons per person to 52.2 gallons and the decline in milk consumption from 31.3 gallons per person to 24.7 gallons (Appendix Table 1).

We begin by discussing the model, data, and estimation procedures. Hypothesis tests, parameter estimates, and elasticities are then presented and discussed. The paper concludes with a summary of the major findings, including implications for the recent expansion in milk advertising financed by fluid milk processors.

Model The Rotterdam model was selected because it is consistent with demand theory (Theil 1965; Barnett); it is as flexible as any other local approximating form (Mountain); it lends itself to advertising applications (e.g., Brown and Lee 1993; Duffy 1987, 1990); and prior testing indicated that the estimated advertising effects from the Rotterdam model were similar to those obtained from its major rival, the (linear approxinmate) Almost Ideal Demand System, and from a double-log specification (Xiao).

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[1]Xiao is an econometrician for American Express in Phoenix, Arizona; Kinnucan is a professor in the Department of Agricultural Economics and Rural Sociology at Auburn University; and Kaiser is an associate professor in the Department of Agricultural, Resource, and Managerial Economics, Cornell University, and is Director of the Cornell Commodity Promotion Research Program. Appreciation is expressed to Noel Blisard, Chanjin Chung, Jong-Ying Lee, Phil Vande Kamp, and Robert Nelson for helpful comments on an earlier draft and to Leen Boon for assistance with data collection. The authors also thank Una Moneypenny for editorial assistance on this manuscript. Funds supporting this research were provided in part by a grant from the National Institute for Commodity Promotion Research and Evaluation. Responsibility for final content, however, rests with the authors.

[2]An exception to this statement is Goddard and Tielu's study of milk advertising in the Ontario market.

Several approaches can be used to augment the Rotterdam specification to include advertising effects. The most common approach, suggested by Theil (1980), is to view advertising as a "taste shifter" that affects marginal utility. In this formulation, advertising enters the model as a price deflator (e.g., Duffy 1987; Brown and Lee 1993). An alternative approach, advocated by Stigler and Becker, is to view advertising (or other information sources) as an input in the household production function. In this formulation, advertising enters the (derived) demand function for market goods as a separate shift variable along with prices and income (e.g., Kinnucan, Xiao, Hsia, and Jackson).

Testing the simple-shift specification against the taste-shift specification using citrus data, Brown and Lee (1993) found them to be statistically equivalent. Accordingly, in this study both forms of the Rotterdam model are estimated to determine the sensitivity of parameter estimates to model specification. The four-equation system consists of demand equations for fluid milk, fruit juices (chiefly orange and apple), soft drinks, and coffee and tea. Thus, weak separability of the non-alcoholic drink group is treated as a maintained hypothesis and total group expenditure is used in place of income in the absolute-price form of the Rotterdam model.[3]

Demographics have been shown to be a significant determinant of milk consumption (Ward and Dixon; Blisard *et al.*). To incorporate demographics into demand systems in a way that preserves the adding-up condition, Pollak and Wales advocate the technique of scaling and translating. With translating, the demographic variables are assumed to affect demand through the income term; with scaling, demographic variables influence demand through the price terms. In a preliminary analysis, we tested the scaling and translating approaches against an alternative model in which the demographic variables (age and food-away-from-home expenditures) enter as simple shift variables. Results indicated a superior fit for the simple-shift specification, the approach taken here.[4] As noted by Piggott *et al.* (1996, p. 270, note 6), adding-up is preserved in the simple-shift specification by requiring the coefficients of the demographic variables to sum to zero across equations.

Because quarterly data were not available on a national basis for any of the beverages except milk, the model is specified in annual form. Although an annual model may produce upward--biased estimates of advertising responses (Clarke, 1976), it has the advantage that lag structures need not be specified. Clarke (1976) finds that the advertising carryover period for mature, frequently-purchased, low-priced items is generally nine months or less. Tomek and Cochrane suggested that long-run demand equations for food items encompass a period of one year or less, an hypothesis that is consistent with the beverage demand literature (e.g., Ward and Dixon; Brown and Lee, 1993).

The basic specification is:

<u>Model A:</u>

(1) $w_{it}\, dln\, q_{it} = a_i + b_i\, dln\, Q_t + \sum_j^4 c_{ij}\, dln\, p_{jt} + \sum_j^4 d_{ij}\, dln\, A_{jt} + e_i\, dln\, AGE_t + f_i\, dln\, FAFH_t + v_{it}$

<u>Model B:</u>

(2) $w_{it}\, dln\, q_{it} = a_i{}' + b_i{}'\, dln\, Q_t + \sum_j^4 c_{ij}{}'\,(dln\, p_{jt} - \gamma_j\, dln\, A_{jt}) + e_i{}'\, dln\, AGE_t + f_i{}'\, dln\, FAFH_t + v_{it}{}'$

where Model A corresponds to the simple-shift specification suggested by Stigler and Becker's analysis and Model B corresponds to Theil's taste-shift specification. In Theil's (1980) original specification, the γ_j parameters in Model B are identical for all j. Following Duffy (1987, p.

[3]For a clear discussion of the distinction between the absolute- and relative-price versions of the Rotterdam model in an advertising context, see Selvanathan (1989b).

[4]Owing to degrees-of-freedom problems, the scaling and translating tests were restricted to the taste-shift version of the Rotterdam model. For details, see Xiao.

1053), we have left γ free to vary between goods to test for differing degrees of effectiveness among the campaigns.

In these models, i indexes the equation (i = 1, 2, 3, 4 for milk, juices, soft drinks, and coffee and tea, respectively) and t indexes the time period (t = 2,3,...,25 for 1971 to 1994). The term $dlnQ_t = \sum_i w_{it}\, dln\, q_{it}$ is the Divisia volume index, which can be interpreted as a third-order approximation to real expenditure on the beverage group (Goldberger, p. 95). The coefficient, w_{it}, corresponds to the expenditure share of beverage item i in year t, q_{it} denotes per capita consumption of beverage item i in year t, p_{jt} is the nominal price of beverage item j in year t, A_{jt} is the real per capita advertising expenditure on beverage item j in year t, AGE_t is the proportion of the U.S. population less than age five in year t, $FAFH_t$ is the ratio of food-away-from-home expenditures to food-at-home expenditures in year t, and v_{it} and v_{it}' are random error terms. An intercept is included in equations (1) and (2) to test for non-specific structural change.

An implicit assumption underlying equations (1) and (2) is that brand and generic advertising have identical effects on aggregate demand. This assumption does not affect soft drinks or coffee and tea, as advertising for these beverages is strictly brand. Nor does it affect milk, since the milk advertising data used in this study are strictly generic. For juices, the data contain significant amounts of both types of advertising as there are both strong brands (e.g., SunKist; Citrus Hill) and active support of generic advertising by citrus growers. To the extent that brand advertising merely shifts market share with no effect on aggregate demand, combining brand advertising with generic would tend to bias the own-advertising coefficient for juices toward zero. However, empirical results for a wide range of products suggest that brand advertising does more than shift market share (Duffy 1987, 1990; Brester and Schroeder; Kaiser and Liu).

Price symmetry and price homogeneity are tested in equation (1) by imposing, respectively, the restrictions $c_{ij} = c_{ji}$ for all i and j and $\sum_j c_{ij} = 0$ for all i. Similar restrictions apply to equation (2), i.e., price symmetry implies $c_{ij}' = c_{ji}'$ for all i and j and price homogeneity implies $\sum_j c_{ij}' = 0$ for all i.

Advertising symmetry and advertising homogeneity (e.g., see Selvanathan 1989b) are tested, respectively, by imposing $d_{ij} = d_{ji}$ for all i and j and $\sum_j d_{ij} = 0$ for all i in Model A.[5] For Theil's specification (Model B), advertising symmetry is tested jointly with price symmetry, as advertising effects are assumed to be proportional to price effects.

Engel aggregation requires that $\sum_i b_i = 1$. Based on the proposition that an advertising-induced increase in the demand for one commodity must be offset by a decrease in the demand for at least one other commodity if the budget constraint is to be satisfied, Basmann (p. 53) developed an adding-up restriction for advertising responses, namely $\sum_i w_i E_{ij}^A = 0$ for all j where E_{ij}^A is the advertising elasticity (defined below). In terms of equation (1), the Basmann aggregation condition implies that $\sum_i d_{ij} = 0$ for all j.

In estimation, one equation is dropped from the system to avoid singularity in the regressors. Because the adding-up conditions are used to obtain coefficients for the deleted equation, adding up is treated as a maintained hypothesis in the Rotterdam model. In addition, the differentials in equations (1) and (2) are approximated by first differences; thus, the intercepts must sum to zero, i.e., $\sum_i a_i = \sum_i a_i' = 0$. Likewise, the coefficients for AGE_t and $FAFH_t$ must sum to zero, i.e., $\sum_i e_i = \sum_i e_i' = 0$ and $\sum_i f_i = \sum_i f_i' = 0$. Further, the price coefficients across equations must sum to zero, i.e., $\sum_i c_{ij} = \sum_i c_{ij}' = 0$. Finally, the coefficients are regarded as fixed constants even though they embed budget shares, which generally change over time. Although these empirical details compromise the generality of the Rotterdam specification, the model is still regarded as a flexible approximation to an unknown true demand system (Barnett; Mountain).

Elasticities from Model A are calculated using the expressions:

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[5]Selvanathan's analysis (1989b, p. 218) identifies a weaker form of symmetry, namely $d_{ik} \leq d_{ki}$. The difference arises from a less restrictive assumption (than Theil's) about how advertising affects marginal utilities.

(expenditure elasticities)

$$E_i^Y = b_i / w_i$$

(Hicksian price elasticities)

$$E_{ij}^* = c_{ij} / w_i$$

(advertising elasticities)

$$E_{ij}^A = d_{ij} / w_i$$

(age elasticities)

$$E_i^{AGE} = e_i / w_i$$

(eating-away-from-home elasticities)

$$E_i^{FAFH} = f_i / w_i$$

Elasticities from Model B employ the same expressions with the following substitutions: $b_i = b_i'$, $c_{ij} = c_{ij}'$, $d_{ij} = -\gamma_j c_{ij}'$, $e_i = e_i'$, and $f_i = f_i'$. Expenditure elasticities are expected to be positive, own-price elasticities negative, and the Hicksian cross-price elasticities are expected to be positive, since beverage products are generally considered to be normal goods and substitutes for each other.

Advertising elasticities in general are *a priori* indeterminate (Basmann, p. 53; Green, Carman, and McManus, p. 65). However, intuitively one would expect own-advertising effects to be positive and cross-advertising effects to be negative for substitute goods. The age and eating-away-from-home elasticities are expected to be positive and negative, respectively, for milk. No *a priori* expectations are placed on the age and eating-away-from-home coefficients for the remaining beverages other than, when combined with the estimated coefficients for fluid milk, they add up to zero across equations.

Data

The models were estimated with annual time-series data covering the period 1970-94.[6] Consumption data for fluid milk, fruit juices, soft drinks, and coffee and tea were obtained from Putman and Allshouse, Table 37. Because tea consumption is modest (about seven gallons per person per year) and has changed little (from a low of 6.8 gallons per person in 1970 and 1990 to a high of 7.7 gallons in 1976), data for tea and coffee were combined. Bottled water consumption, which increased from 1.2 gallons per person per year in 1976 (the first available figure) to 9.2 gallons in 1993, is not considered in this study because the series is incomplete. The included beverages account for 92.5 percent of total non-alcoholic beverage consumption in 1993.

Price data were obtained primarily from the U.S. Department of Labor's *CPI Detailed Report*. To facilitate the computation of budget shares, the CPIs for each beverage were converted to per-gallon prices using standard unit conversions. A composite price series for coffee and tea was obtained by taking the quantity-share weighted average of the tea and coffee prices. As a proxy for the price of juices, the price of frozen orange-juice concentrate was used because orange juice represents the major component of the juice category. A complete description of the price series, along with data sources, is provided in the data appendix.

The advertising data were obtained from annual issues of *AD $ SUMMARY* published by Leading National Advertisers, Inc. LNA is a tracking service agency that estimates the advertising expenditures for all brands (including industry organizations such as the National Dairy Board) that spend at least $25,000 per year in a particular medium. The media tracked by LNA include network, spot, syndicated, and cable television; network and national spot radio; magazines (including Sunday supplements); newspapers; and outdoor. A complete description of the LNA data used in this study is provided in the data appendix. The advertising data were divided by the CPI for all items (1982-84 = 100) to remove the effects of inflation. Sources and definitions for the CPI, population, age, and food-away-from-home variables are provided in the data appendix.

[6]The sample covers a period of substantial changes in the level of soft drink and milk advertising. For example, milk advertising in the early 1980s (prior to the implementation of federal legislation authorizing the nationwide mandatory check-off) was about $23 million per year; by 1994 it had more than tripled to $79 million. Soft drink advertising, over the same period, increased from $250 million per year to $462 million (see Appendix Table 2). No attempt was made in this study to determine whether the large increases in expenditures affected response coefficients.

Estimation Procedure

The models were estimated using seemingly unrelated regressions (SUR) to accommodate the imposition of the parametric restrictions. Simultaneous-equation procedures are not used because previous research suggests that price endogeneity is relatively unimportant in demand-system estimation when the commodities in question constitute a small portion of the consumer budget (Bronsard and Salvas-Bronsard), as is the case for non-alcoholic beverages. Theil's theory of rational random behavior suggests that group expenditure is independent of the error term in the Rotterdam model. This was confirmed in Brown, Behr, and Lee's analysis.

The adding-up constraint implies that only three equations in the system are independent. The usual procedure, followed in this study, is to drop one equation, estimate the remaining system, and calculate the parameters from the omitted equation using the classical restrictions. The estimates provided in this paper were obtained using the systems estimator in Eviews, the Windows version of Micro-TSP.

Theoretical restrictions were successively imposed and tested using the Wald criterion. Based on these tests, an appropriately restricted model was used to test for structural change, i.e., whether the nonprice variables in equations (1) and (2) (including trend) are significant. All tests, unless indicated otherwise, use a significance level of 5 percent. Elasticities are evaluated at mean budget shares for 1990-94, the last five years in the sample. Due to the first-difference form of the Rotterdam model, the first observation is lost and the parameter estimates are based on 24 annual observations.

Results

Preliminary tests based on the Durbin-Watson statistic showed no evidence of serial correlation in the unrestricted equations. Wald tests indicated that price and advertising homogeneity and advertising symmetry are compatible with the data, but that price symmetry is not (Table 1). Similar results were obtained by Goddard and Tielu in their study of non-alcoholic beverage advertising in the Ontario market, although in their study both advertising symmetry and price symmetry were rejected. To conserve degrees of freedom and to provide a basis for assessing the effects of the restriction on the estimated parameters, price symmetry is imposed on Model A. The imposition of price symmetry, as noted by Goddard and Tielu (p. 270), has the further advantage in assuring that parameter estimates are consistent with consumer theory.

Further testing indicated that trend and advertising contribute significantly to the explanatory power of both models ($p < 0.0002$, see Table 1). *AGE* and *FAFH*'s contribution, however, is marginal. Specifically, *AGE*, when considered separately, is not significant in either model, and *FAFH* is significant in Model B but not Model A. However, *AGE* and *FAFH* are jointly significant at the 6.5 percent level in Model A and at the 0.44 percent level in Model B and for this reason are retained. Theil's assumption that the proportionality constant γ in Model B is the same for all goods is rejected at the 0.02 percent level. Based on these tests, the restricted forms selected for coefficient estimation are Model A4 and Model B1 in Table 1. Model B1 is less restrictive than Model A4 in that Model B1 does not impose price symmetry. Thus, a comparison of the parameter estimates from A4 and B1 permits an evaluation of the extent to which the classical restrictions affect the parameter estimates.

Most of the estimated coefficients have the expected signs and are significant (Table 2). The Durbin-Watson statistics for the restricted models (A4 and B1) are similar to the Durbin-Watson statistics for their unrestricted counterparts (Models A and B), which suggests that the restrictions do not induce specification error. (Recall that price symmetry, which is imposed in Model A4, was rejected by the Wald test.) The R^2s range from 0.47 for milk to 0.71 for juices in Model A4 and from 0.57 for soft drinks to 0.75 for juices in Model B1. Thus, between 47 percent and 75 percent of the observed year-to-year changes in beverage consumption can be "explained" by the models, with

Table 1. Wald Tests of Model Restrictions

Model	Restrictions	Computed χ^2	p-value
A	Maintained hypothesis (Equation 1)	--	--
A1	Price homogeneity (PH)	6.706	0.0819
A2	Price homogeneity and price symmetry (PS)	21.980	0.0012
A3	PH, PS and advertising homogeneity (AH)	4.556	0.2073
A4	PH, PS, AH and advertising symmetry (AS)	7.4091	0.2847
A5	PH, PS, AH, AS and $a_i = 0$, all i	19.944	0.0002
A6	PH, PS, AH, AS and $e_i = 0$, all i	5.262	0.1536
A7	PH, PS, AH, AS and $f_i = 0$, all i	5.472	0.1403
A8	PH, PS, AH, AS and $e_i = f_i = 0$, all i	11.867	0.0650
A9	PH, PS, AH, AS and $a_i = e_i = f_i = 0$, all i	66.094	0.0000
A10	PH, PS, AH, AS and $d_{ij} = 0$, all i, j	85.802	0.0000
B	Maintained hypothesis (Equation 2)	--	--
B1	PH	0.7075	0.8714
B2	PH and PS	30.424	0.0000
B3	PH and $a_i = 0$, all i	26.351	0.0000
B4	PH and $e_i = 0$, all i	6.106	0.1066
B5	PH and $f_i = 0$, all i	10.751	0.0132
B6	PH and $e_i = f_i = 0$, all i	18.879	0.0044
B7	PH and $a_i = e_i = f_i = 0$, all i	100.58	0.0000
B8	PH and $\gamma_1 = \gamma_2 = \gamma_3 = \gamma_4 = \gamma \neq 0$	19.487	0.0002
B9	PH and $\gamma_1 = \gamma_2 = \gamma_3 = \gamma_4 = 0$	21.917	0.0002

Table 2. SUR Coefficient Estimates of Non-Alcohol Beverage Demand for Alternative Forms of the Rotterdam Model, 1971-94 Annual data

Equation	Price Coefficients				Advertising Coefficients				Expend.	Intercept	*AGE*	*FAFH*	R^2	*D.W.*
	c_{i1}	c_{i2}	c_{i3}	c_{i4}	d_{i1}	d_{i2}	d_{i3}	d_{i4}	b_i	a_i	e_i	f_i		
Model A4[a]														
Milk	-0.0453				0.0009				0.0850	-0.0028	0.0560	-0.0632	0.474	2.05
	(-4.87)[b]				(1.05)				(2.73)	(-3.46)	(1.23)	(-2.56)		
Juices	0.0310	-0.0670			0.0092	0.0219			0.1909	-0.0016	0.1197	0.0772	0.709	2.58
	(3.78)	(-2.80)			(4.60)	(2.36)			(2.52)	(-0.78)	(1.04)	(1.30)		
Soft Drinks	0.0080	0.0287	-0.0551		-0.0047	0.0068	-0.0377		0.4608	0.0091	-0.0688	-0.0535	0.564	2.11
	(1.01)	(1.63)	(-2.76)		(-1.76)	(0.66)	(-2.02)		(4.78)	(3.42)	(-0.50)	(-0.70)		
Coffee & Tea	0.0063	0.0073	0.0185	-0.0321	-0.0054	-0.0380	0.0355	0.0078	0.2633	-0.0047	-0.1077	0.0395	0.481	2.53
	(1.39)	(0.66)	(1.44)	(-2.42)	(-2.16)	(-4.56)	(2.67)	(0.59)	(2.78)	(-1.88)	(-0.81)	(0.56)		
Model B1[a]														
Milk	-0.0145	0.0229	-0.0112	0.0028	0.1077				0.0464	-0.0030	0.0613	-0.0470	0.586	1.58
	(-2.28)	(3.54)	(-1.97)	(1.40)	(3.63)				(1.72)	(-3.88)	(1.67)	(-2.02)		
Juices	-0.0600	-0.0147	0.0503	0.0243		-0.6125			0.3556	-0.0027	0.1068	0.0620	0.746	1.99
	(-3.60)	(-1.08)	(2.93)	(2.99)		(-3.39)			(5.06)	(-1.35)	(1.08)	(1.02)		
Soft Drinks	0.0263	0.0378	-0.0603	-0.0037			-0.7757		0.3693	0.0102	-0.2422	-0.0682	0.566	1.96
	(1.29)	(2.08)	(-2.82)	(-0.53)			(-2.60)		(4.03)	(3.81)	(-1.93)	(-0.82)		
Coffee & Tea	0.0481	-0.0460	0.0212	-0.0234				1.4950	0.2286	-0.0046	0.0742	0.5329	0.658	1.89
	(2.86)	(-2.87)	(1.41)	(-3.13)				(2.59)	(2.99)	(-2.22)	(0.72)	(0.83)		

[a] Model A4 imposes price homogeneity and symmetry and advertising homogeneity and symmetry on text equation (1); Model B1 imposes price homogeneity on text equation (2).

[b] Numbers in parentheses are asymptotic *t*-ratios.

a slight edge given to Model B1. However, sufficient differences exist between the models to suggest that inferences are sensitive to model choice, especially with respect to estimated advertising effects.

Price Effects

The major difference in the estimated price effects between the models pertains to juices. In Model A4 the estimated own-price effect for juices is significant; in Model B1 it is not. In addition, Model B1 indicates that juices and milk are net complements, at least when viewed from the perspective of juice consumption, whereas Model A4 indicates that milk and juices are net substitutes.

Similar results are obtained with respect to juices and coffee and tea. In particular, Model A4 indicates that juices and coffee and tea are independent, whereas Model B1 indicates that juices and coffee and tea are not independent. The conflicting results can be traced to the imposition of the price symmetry restriction in Model A4. If price symmetry is not imposed (Model B1), the estimated cross-price effects between juices and milk and between juices and coffee and tea are decidedly asymmetric (compare the estimated coefficients c_{12} and c_{21} and c_{24} and c_{42} for Model B1 in Table 2). Imposing the symmetry restriction in Model A4 produces a net effect that in the case of juices and milk is positive (implying substitutes) and in the case of juices and coffee and tea is zero (implying independence).

Turning to elasticities, all own-price elasticities are similar and plausible (Table 3). Model A4's own-price elasticities for milk (-0.161) and for juices (-0.426) compare favorably with estimates in the literature. Ward and Dixon's (p. 735) own-price elasticity estimate for milk is -0.153; Brown, Behr and Lee's (p. 137) estimates for individual juice products range from -0.892 for grape juice to -1.606 for grapefruit juice.[7] The models are consistent in suggesting that the demand for soft drinks and coffee and tea is price inelastic. Model A4's elasticities for soft drinks and coffee and tea are -0.127 and -0.253, respectively; Model B1's corresponding estimates are -0.139 and -0.184. The major differences in the models pertain to cross-price elasticities. In addition to showing a number of complementary relationships, Model B1's cross-effects tend to be more pronounced than Model A4's.

Expenditure Effects

Estimated expenditure effects are consistent in the two models. That is, total beverage expenditure is a significant determinant of the demand for milk, juices, soft drinks and coffee and tea. Elasticity estimates indicate that milk is the least responsive to changes in beverage expenditure (0.165 to 0.302), followed by soft drinks (0.85 to 1.06). Juices (1.21 to 2.62) and coffee and tea (1.80 to 2.07) vie for the most expenditure-responsive members of the group. These results are consistent with previous findings. Ward and Dixon's estimate of the income elasticity for milk is 0.293; Brown and Lee's (1993, p. 431) estimates of the expenditure elasticities for citrus products range from 0.94 to 1.03

Age, FAFH, and Trend Effects

Among the variables indicating structural change, *AGE* and *FAFH* have the least influence and trend the most. Both models are consistent in showing that *FAFH's* effect is limited to milk. The estimated *FAFH* effect is inelastic (-0.167 to -0.225), which suggests that further increases in the food-away-from-home/food-at-home expenditure ratio will have only a modest depressing effect on milk consumption.

The age effect, which is more significant in Model B1 than in Model A4, appears to be limited to milk and soft drinks. For soft drinks, Model B1 provides an *AGE* elasticity estimate of -0.56. Thus,

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[7]That Brown, Behr and Lee's own-price elasticities are larger in absolute value than ours is expected: narrowly defined products have more substitutes than the corresponding aggregate. Also, Brown, Behr and Lee's estimates are Marshallion elasticities, which are more elastic than their Hicksian counterparts when expenditure effects are positive, as they are in Brown, Behr and Lee's study.

Table 3. Hicksian Price Elasticities and Expenditure Elasticities for Non-Alcoholic Beverages, United States, Evaluated at 1990-94 Mean Data Points

Quantity of:	Price of:[a]				Expenditure
	Milk	Juices	Soft Drinks	Coffee & Tea	
MODEL A4:					
Milk	-0.1608*	0.1102*	0.0284	0.0224	0.3022*
Juices	0.1971*	-0.4260*	0.1827	0.0465	1.2140*
Soft Drinks	0.0183	0.0660	-0.1268*	0.0426	1.0600*
Coffee & Tea	0.0498	0.0578	0.1455	-0.2530*	2.0730*
MODEL B1:					
Milk	-0.0515*	0.0813*	-0.0398*	0.0010	0.1650
Juices	-0.3813*	-0.0934	0.3200*	0.1548*	2.2623*
Soft Drinks	0.0605	0.0869*	-0.1389*	-0.0086	0.8501*
Coffee & Tea	0.3791*	-0.3619*	0.1672	-0.1843*	1.8001*

[a] Asterisk indicates the estimated coefficient is significant at the 5 percent level according to a two-tail *t*-test. Elasticities are evaluated at 1990-94 mean conditional budget shares as follows: $w_1 = 0.2813$, $w_2 = 0.1571$, $w_3 = 0.4344$, $w_4 = 0.1270$.

a one-percent increase in the proportion of the U.S. population under age five, *ceteris paribus*, is associated with a decline in per-capita soft drink consumption of 0.56 percent. The corresponding elasticity for milk obtained from Model B1 is 0.22, which suggests that milk consumption is less sensitive to changes in the age structure than soft-drink consumption. It also suggests that recent increases in the under-age- five population proportion (see Appendix Table 2) will provide a modest boost to milk consumption, *ceteris paribus*. Caution, however, is required in interpreting these elasticities in that the estimated age effects are model sensitive.

Trend effects (the Rotterdam model's intercept) are significant in all equations except juices and are robust across the models (Table 2). According to Deaton and Muellbauer (p. 70), the intercepts can be interpreted as the *per annum* change in the budget share w_i that would take place in the absence of changes in real total expenditure and relative prices. Applying this interpretation to the estimated values in Table 2, there appears to have been a trend increase in the share for soft drinks, largely offset by trend decreases in the budget shares going to coffee and tea and milk. These changes are perhaps the most important and obvious shifts in the pattern of U.S. non-alcoholic beverage consumption over the past 25 years (see Appendix Table 1). The fact that they apparently cannot be explained in terms of changes in real income, price structure, advertising, and the demographic variables suggests that structural change is at work. In particular, it appears that changes in consumers' tastes are an important contributing factor to the observed consumption pattern.

To gauge the relative importance of taste change, we computed the following trend coefficients:

$$\Sigma_i^{TREND} = (a_i/w_i)*100 \qquad i = 1, ...,4$$

where Σ_i^{TREND} is the *per annum* percent change in quantity where prices, expenditure, advertising, and demographics are held constant. The numerical values for these trend coefficients based on Model A1 using budget shares for 1990-94 are as follows: Σ_1^{TREND} = -1.00, Σ_2^{TREND} = -1.02 (insignificant), Σ_3^{TREND} = 2.10, and Σ_4^{TREND} = -3.70. According to these estimates, taste changes alone would be associated with a decline in per-capita milk consumption of 1 percent per year, an increase in per capita soft-drink consumption of 2.1 percent per year, and a decline in per-capita coffee and tea consumption of 3.7 percent per year. A comparison of the actual and predicted changes based on taste change for 1990-94 is as follows:

	Predicted	Actual	Ratio
milk	-4.98%	-3.89%	1.28
soft drinks	10.47%	12.74%	0.82
coffee & tea	-18.50%	-15.61%	1.18

As can be seen, for the commodities with a significant trend, all but 18-28 percent of the observed consumption change can be explained by changes in taste. Stated another way, economic variables (including advertising) appear to account for at most 28 percent of the observed consumption pattern between 1990 and 1994. The finding that taste change accounts for a large portion of the observed consumption pattern reinforces inferences based on the statistical tests (see Table 1, restrictions A5 and B3), which suggest that trend cannot be deleted from the demand system without causing serious specification error.

Advertising Effects

The robustness issue is most pronounced in the advertising effects. For example, using a *t*-ratio of 1.65 to indicate significance -- the cut-off for a two-tail test at the 10 percent level -- two of the estimated own-advertising effects that are significant in Model B1 (milk and coffee and tea) are not significant in Model A4. (Significance in Model B1 is determined by testing whether the compound term $\gamma_j\, c_{ij}'$ is zero -- see Table 4, note a.) Moreover, the own-effect for juice advertising, which is significant in Model A4, is insignificant in Model B1.

The only consistency between the two models in the estimation of own-advertising effects is for soft drinks. In this case, however, the own-advertising effect is negative. One interpretation of this result is that satiation effects are at work. Soft-drink advertising at $462 million in 1994 was five times milk advertising and nearly double the level of juice advertising and coffee and tea advertising (Appendix Table 2). A negative own-advertising effect for soft-drinks was also obtained by Goddard and Tielu[8]

Turning to the cross-advertising effects, both models are consistent in showing that milk advertising has no effect on soft-drink demand and negatively affects coffee and tea demand (Table 4, column). Also, both models are consistent in showing that soft-drink advertising has no effect on milk demand (Table 4, row). Similar results were obtained by Goddard and Tielu with respect to the Ontario market. Specifically, soft drink and juice advertising were found to have little effect on milk demand, while milk advertising had a relatively large effect on juice demand (Goddard and Tielu, p. 273). However, in our results the models are inconsistent in indicating how milk advertising affects juice demand. In particular, Model A4 indicates that milk advertising has a positive effect on juice demand whereas Model B1 indicates that the effect is negative.

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[8] The Almost Ideal Demand System and double- log models estimated in preliminary analysis also produced a negative and statistically significant own-advertising effect for soft drinks. Thus, the result is robust to functional form. That satiation effects may be at work receives support from a study by Clarke (1973) in which he found that advertising competition had forced a number of brands to increase advertising expenditures "...passed the point of diminishing returns" (p. 259). In fact, by the end of the study period, nine of the 18 brands were overspending, including five of the six largest brands. The intensity of advertising competition that encourages the overspending was demonstrated by Clarke (1973, p. 258) with the following example: "A11 is the largest selling brand in the industry, but if A11 increased advertising by 1percent, its two major competitors could cut A11's expected sales *increase* (sic) from an 8.4 percent gain to only a 4.94 percent gain by increasing their own advertising 1 percent! (sic)" This may explain the "cola wars" and the consequent overspending implied by our estimates.

The Proportionality Hypothesis

Given the conflicting results produced by the two models, especially with respect to advertising effects, the question arises whether they are statistically equivalent. To test this, we formed the hypothesis:

(3a) $H_N : d_{ij} = - \gamma_j c_{ij}$

For all i and j

(3b) $H_A : H_N$ not true

where c_{ij} and d_{ij} are the price and advertising coefficients, respectively, in equation (1), and γ_j are the proportionality coefficients in equation (2). Hypothesis (3a) is Theil's proportionality hypothesis. When the restriction is true, equation (1) reduces to equation (2). Thus, to determine whether models A4 and B1 are equivalent, it is sufficient to test hypothesis (3).

The tests were conducted using a Wald statistic as indicated in Table 5. Because Model B1 does not impose price symmetry and Model A4 does, we also tested less restrictive forms of the two models. Specifically, to remove the effect of price symmetry, we tested Model B1 against Model A1. In yet a third test, we contrasted Model B against Model A, perhaps the purest test in that none of the classical restrictions is imposed on either model.

Results from all three tests indicate decisive rejection of the proportionality hypothesis ($p < 0.0000$). That is, Theil's hypothesis that advertising elasticities are proportional to price elasticities is not supported by our data. This is true notwithstanding the latitude given the hypothesis in our model; namely, that the proportionality factor γ be permitted to vary across goods. (Recall that Theil posited that the γ_j s are the same for all goods.) Thus, the two models are not statistically equivalent, which implies that the parameter estimates from the Theil specification (Model B1) should be treated with caution.

Spillover Effects

A critical issue from the standpoint of advertising benefit-cost analysis is spillover, i.e., whether one commodity's advertising affects the demand for related goods. Returning to the cross-advertising elasticities in Table 4, and focusing on Model A4, fully two-thirds of the estimated cross-effects are significant at the 5 percent level or lower. Morever, among the significant cross-elasticities, most are larger in absolute value than the corresponding own-advertising elasticities. For example, the cross-elasticities of milk advertising with respect to juice demand (0.059) and coffee and tea demand (-0.043) are at least 13 times larger in absolute value than milk's own-advertising elasticity (0.003), which is not significant at usual probability levels.

The foregoing elasticity estimates suggest that milk advertising may be more effective at altering the demand for related beverages than at increasing its own demand, a result consistent with Goddard and Teilu's findings. Specifically, Goddard and Tielu's (p. 273) cross-elasticities of milk advertising with respect to tomato (0.086),orange (-0.100), and apple (-0.037) juice demand are at least nine times larger in absolute value than the own-advertising elasticity for milk (0.004). Similar results obtain both in our study and in Goddard and Tielu's for juice and soft drink advertising, although the ratios of cross-effects to own-effects are not as pronounced as for milk.

Concluding Comments

Results presented in this paper support Galbraith's hypothesis. Specifically, the hypothesis that advertising has no effect on the aggregate demand for specific items within the non-alcoholic beverage groups is rejected decisively. However, the estimated direct effects of advertising are modest and, with the exception of soft drinks, fragile. For example, the estimated own-advertising elasticity for milk ranges from a statistically insignificant 0.0032 in the simple-shift version of the Rotterdam model to 0.0055 in the taste-shift specification. Although the latter estimate is statistically significant, it is so tiny as to suggest

Table 4. Advertising and Demographic Elasticities for Non-Alcoholic Beverages, United States, Evaluated at 1990-94 Mean Data Points

Quantity of:	Advertising of:				*AGE*	*FAFH*
	Milk	Juices	Soft Drinks	Coffee & Tea		
MODEL A4:						
Milk	0.0032	0.0327*	-0.0167	-0.0192*	0.1989	-0.2245*
Juices	0.0585*	0.1394*	0.0433	-0.2419*	0.7612	0.4913
Soft Drinks	-0.0108	0.0158	-0.0868*	0.0817*	-0.1564	-0.1232
Coffee & Tea	-0.0426*	-0.2991*	0.2798*	0.0618	-0.8479	0.3109
MODEL B1:						
Milk	0.0055**	-0.0088**	0.0043	-0.0011	0.2179	-0.1673*
Juices	-0.2336**	-0.0572	0.1960**	0.0948**	0.6790	0.3943
Soft Drinks	0.0470	0.0674	-0.1077**	-0.0067	-0.5575	-0.1570
Coffee & Tea	-0.5668**	0.5411	-0.2499	0.2756**	0.5841	0.4196

[a] Single asterisk indicates that the estimated coefficient is significant at the 5 percent level according to a two-tail *t*-test. Double asterisk indicates the estimated coefficient is significant at the 5 percent level according to a Wald test of the non-linear restriction $\gamma_j c_{ij}' = 0$. Elasticities are evaluated at 1990-94 mean conditional budget shares as follows: $w_1 = 0.2813$, $w_2 = 0.1571$, $w_3 = 0.4344$, $w_4 = 0.1270$.

Table 5. Wald Tests of the Proportionality Hypothesis

Model Comparison	Computed χ^2	Probability	Result
A vs. B	93.143	0.000000	Reject Model B
A1 vs. B1	80.759	0.000000	Reject Model B1
A4 vs. B1	42.614	0.000000	Reject Model B1

Note: Model B and its variants contain the proportionality hypothesis. Tests are conducted for $\gamma_1 = 0.1077$, $\gamma_2 = -0.6125$, $\gamma_3 = -0.7757$, and $\gamma_4 = 1.4950$, the point estimates given in Table 2.

that even with the large spending increases associated with the recent (post-1994) fluid milk processor initiative, there is little to expect in the way of changes in per-capita milk consumption.[9] Similar inferences apply to juice advertising, although the own-advertising elasticity estimate from the statistically superior simple-shift specification (0.1394) is large enough in a relative sense to suggest that changes in juice advertising might have important effects on the consumption pattern.

Theil's hypothesis that advertising effects are proportional to price effects is rejected by our data. Given the importance of the hypothesis in simulation work (Wohlgenant), model specification (Green, Carman and McManus), and estimation (Duffy 1987, 1990; Selvanthanan 1989a; Brown and Lee 1993), further testing is warranted. Specifically, it would be useful to test the hypothesis on a wider array of goods and with other data sets and models to establish robustness. Clearly, given its elegance and usefulness, it would be imprudent to abandon the proportionality hypothesis on the basis of a single test.

The dominant pattern in U.S. non-alcoholic beverage consumption over the past 25 years has been a steady increase in per-capita soft-drink consumption, largely at the expense of coffee consumption and, to a lesser extent, milk consumption. Although changes in relative prices, real beverage expenditures, and advertising have influenced this pattern, our results suggest that the major factor responsible for the observed consumption pattern is structural change. The basis for this claim is that the trend coeffecient in each of the estimated equations except juice is significant and numerically large. Specifically, *per annum* changes in per-capita consumption related strictly to trend are estimated at -1.0 percent for milk, 2.1 percent for soft drinks, and -3.7 percent for coffee and tea. Applying these coefficients to the observed consumption pattern for the most recent five years of our sample (1990-94) we find that fully 80 percent of the observed change can be explained by trend (taste change), leaving only 20 percent to be accounted for by economic variables, including advertising. The effects of demographic variables, namely the aging of the U.S. population and the increased incidence of meals taken away from home, appear to be confined to milk, and to be less important than taste change in explaining the observed consumption pattern.

[9] This is not to say that the fluid milk processor initiative is necessarily unprofitable. Given the size of the U.S. fluid milk market ($15.2 billion at retail in 1994) relative to the processor advertising investment ($114 million per year), it does not take much of a demand increase to yield a favorable benefit-cost ratio, especially if fluid milk supplies are relatively price inelastic. And in light of the positive spillover effect of milk advertising onto the juice market indicated in Table 4 (Model A4), it is possible that the total elasticity for milk advertising is positive, even if the partial elasticity is zero. A total elasticity for milk advertising in principle could be calculated using procedures described by Piggott, Piggott, and Wright or by Kinnucan (1997) (see also Kinnucan and Belleza). However, that would entail specifying a complete structural model of the U.S. non-alcoholic beverage market, including linkages between market levels, which is beyond the scope of this paper.

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Data Appendix

Retail price series were developed in a three step-procedure. First, the U.S. city average price of each beverage in December 1995 was obtained from the *CPI Detailed Report*. These prices were then divided by each beverage's CPI for December 1995 (1982-84 = 100) to obtain an average price for the 1982-84 base period. The base-period prices were then multiplied by each beverage's annual CPI (1982-84 = 100) to form the price series. Because the *CPI Detailed Report* does not list a price for tea, a modified version of the foregoing procedure had to be used for tea. In addition, unit conversions and other complications arose with the other beverages. Details are provided below.

Fluid milk price: The fluid milk price was proxied as a simple average of whole and low-fat milk prices. The December 1995 U.S. city average price for fresh, whole, fortified milk is \$2.518 per gallon; the corresponding price for fresh, low-fat milk is \$2.310 per gallon. Applying the foregoing procedure to the simple average of these two prices yields a base-period price of $1.806 per gallon.

Fruit juice price: The price of frozen orange-juice concentrate was taken as a proxy for fruit-juice price. The December 1995 U.S. city average price of frozen orange-juice concentrate was $1.573 per 16 oz. (473 ml). Since one gallon equals 3,800 ml, a gallon of concentrate cost $12.637. Assuming that concentrate is mixed with water in a 3:1 ratio (one part concentrate to 3 parts water), this implies a December 1995 price of $4.212 per gallon drinking juice. Applying the foregoing procedure to this price yields a base-period price of $3.080 per gallon.

Soft drink price: The price of regular cola in two liter containers was taken as a proxy for the price of soft drinks. The December 1995 U.S. city average price of regular cola was $0.996. Using the conversion 3.8 liters equals one gallon, this translates into a December 1995 cola price of $1.892 per gallon. Applying the foregoing procedure to this price yields a base-period price for soft drinks of $1.597 per gallon.

Coffee price: The price of coffee was measured as the simple average of *instant* and *ground roast* coffee price. The December 1995 U.S. city average price of *instant* coffee is $10.299 per pound. Each pound of instant makes approximately 186.8 cups of 6.0 oz. fluid coffee or 8.759 gallons. So the December 1995 price of *instant* coffee is $1.176 per gallon drinking coffee. The December 1995 U.S city average price of *ground roast* coffee is $3.507 per pound. Ground roast coffee makes approximately 59.8 cups of 6.0 oz. fluid coffee, or 2.803 gallons. So the December 1995 price of *ground roast* coffee is $1.251 per gallon drinking coffee. Applying the foregoing procedure to the simple average of the instant and ground-roast prices yields a base-period price for coffee of $0.777 per gallon.

Tea price: The tea price series was complicated by the fact that the *CPI Detail Report* does not list a price for tea and ceased publishing a price index for tea in 1977. The latter problem was solved by constructing a price index (1982 = 100) for the period 1975-95 based on data in *Tropical Products: World Markets and Trade* (pp. 36-37) provided by the International Tea Committee (ITC). This index was spliced to the USDL's tea index to obtain an index for the entire sample period 1970-94. To convert the index to actual prices, the price of tea in 1978 was obtained from ITC data published in *Estimated United States Average Retail Price of Food*, which lists an average price for tea in 1978 of $1.235 for tea bags, 40-bag package. Assuming that each tea bag produces approximately 7.2 oz. of tea, this translates into 2.242 gallons of liquid tea per package, or a 1978 price of $0.551 per gallon. Dividing this price by the CPI for tea in 1978 (1982 = 100) provides an estimate of the base-period price. Multiplying the base-period price by the annual CPI for tea (1982 = 100) provided the tea price series.

The composite price series for coffee and tea was constructed as a quantity-share weighted average of the foregoing tea and coffee prices. The price and quantity series used in this study are given in Appendix Table 1. The advertising series and related data are given in Appendix Table 2. Basic data sources for the nonprice series and special notes are as follows:

q_1 to q_4: The source for the quantity data is Putman and Allshouse, Table 37.

p_1 to p_4: The basic data source for the price series was the U.S. Department of Labor's *CPI Detailed Report*, Table P4, pp. 234-35, which reports average retail food prices for U.S. cities and four regions. This source, however, does not list a price for tea. The sources and methods used to obtain a tea price series are provided in the appendix narrative.

a_1 to a_4: The basic source for the advertising data is *AD $ SUMMARY* published by Leading National Advertisers, Inc., New York City. The relevant LNA categories are as follows: F131 (milk, butter, eggs), F171(coffee, tea, and cocoa), F172 (fruit drinks), F221 (regular soft drinks), F222 (diet soft drinks), F223 (non-carbonated soft drinks). Because of definitional changes and aggregation, several adjustments had to be made before these data could be used for analysis. First, in 1984 LNA broadened the juice category (F172) to include powdered drinks, which was formerly in the F223 category. At the same time, LNA added a new category (F224), bottled water, which was formerly in F223. Since it was not possible to isolate the proportion of F172 expenditures that is strictly juice advertising in the redefined series, it was decided that the best approach was simply to add the three categories. That is, in our study, fruit-juice advertising is measured as F172+F223+F224.

The second adjustment has to do with the F131 category. This category includes expenditures for butter and eggs as well as fluid milk. To isolate the milk expenditures, we collected data for F131 "brands" as follows: National Dairy Board, California Milk Advisory Board, American Dairy Association, United Dairy Industry Association, Mid-Atlantic Farmers' Milk, Dairymans' Dairy Products, and Cow Dairyman Association. Thus, the data for fluid milk advertising used in this study refer strictly to generic advertising expenditures as reported by LNA. (The series excludes expenditures by the newly-formed Fluid Milk Processors' Board as that campaign commenced in 1995, a year later than our sample period.)

The third adjustment has to do with missing values. Data prior to 1974 for juices, soft drinks and coffee and tea were unavailable. For milk, no data were available for 1974 and 1975. The latter two data points were obtained by interpolation. For the other beverages, the missing values were "backcast" from the regression equation $AD_{it} = \alpha + \beta t + \gamma t^2 + \epsilon_t$ where AD_{it} is the total advertising expenditure for good i in period t as reported by LNA, and t is a trend variable that assumes the values 5,6,...,10 for 1974-83. The regressions were run on the combined juice series F172+F223+F224, the combined soft-drink series F221+F222, and the single series F171 for coffee and tea. The missing values for 1970-73 were computed from the estimated regressions by setting t = 1, 2, 3, and 4, respectively, and computing AD_{it} when the residuals are zero.

AGE: The proportion of the U.S. population less than age five was obtained from Table B-30 in *Economic Report to the President*, p. 315.

FAFH: This is expenditures on food-away-from-home divided by expenditures on food-at-home. Data source is Putman and Allshouse, Table 98, p. 136.

POP: Resident U.S. population on July 1. Source: Putman and Allshouse, Table 115.

CPI: Consumer Price Index for all items for all urban consumers. Source: Putman and Allshouse, Table 101.

Appendix Table 1. Quantity and Retail Price Data for Non-alcoholic Beverages, United States, 1970-94

YEAR	q_1	q_2	q_3	q_4	p_1	p_2	p_3	p_4
	(------------ Gallons/person			----------------)	(--------------- Dollars/gallon			----------------)
1970	31.3	5.7	24.3	40.2	0.904	1.291	0.609	0.250
1971	31.3	5.7	25.5	40.4	0.928	1.341	0.644	0.258
1972	31.0	6.2	26.2	40.9	0.942	1.434	0.656	0.255
1973	30.5	6.0	27.6	40.7	1.031	1.445	0.674	0.279
1974	29.5	6.0	27.6	40.7	1.235	1.497	0.834	0.327
1975	29.5	6.6	28.2	38.9	1.236	1.616	1.026	0.358
1976	29.3	6.9	30.8	40.2	1.301	1.655	0.994	0.475
1977	29.0	7.0	33.0	32.0	1.314	1.993	1.041	0.830
1978	28.6	6.4	34.2	34.5	1.390	2.114	1.131	0.774
1979	28.2	6.8	34.7	36.2	1.551	2.311	1.234	0.744
1980	27.6	7.2	35.1	34.0	1.688	2.473	1.383	0.801
1981	27.1	7.4	35.4	33.2	1.783	2.833	1.522	0.696
1982	26.4	6.8	35.3	32.8	1.793	2.965	1.562	0.707
1983	26.3	8.4	35.2	33.3	1.805	2.973	1.602	0.729
1984	26.4	7.3	35.9	33.9	1.819	3.248	1.626	0.813
1985	26.7	7.7	35.7	34.5	1.847	3.412	1.642	0.754
1986	26.5	7.9	35.8	34.6	1.836	3.239	1.654	0.919
1987	26.3	8.2	41.9	33.6	1.871	3.377	1.688	0.806
1988	25.8	8.2	44.7	32.6	1.914	3.788	1.688	0.799
1989	26.0	7.7	45.4	33.0	2.064	3.905	1.731	0.847
1990	25.7	6.9	46.3	33.6	2.288	4.313	1.790	0.833
1991	25.7	7.9	47.9	33.6	2.210	4.080	1.805	0.809
1992	25.4	7.3	48.5	32.9	2.282	4.270	1.835	0.785
1993	24.9	8.4	50.2	30.5	2.309	4.040	1.851	0.765
1994	24.7	8.6	52.2	28.1	2.369	4.059	1.848	0.934

Note: The subscripts are defined as follows: 1 = fluid milk, 2 = juices, 3 = soft drinks, and 4 = coffee and tea. See appendix narrative for sources and explanatory notes.

Appendix Table 2. Advertising and Remaining Data Used to Estimate the Non-alcoholic Beverage Demand System, United States, 1970-94

YEAR	A_1	A_2	A_3	A_4	*AGE*	*FAFH*	*POP*	*CPI*
	(------------ Million dollars			----------------)	(%)	(Ratio)	(Thous.)	
1970	1.903	9.308	24.173	19.711	8.372	0.356	203984	0.388
1971	4.246	21.117	48.346	39.422	8.304	0.360	206827	0.405
1972	11.346	32.924	72.519	59.133	8.147	0.371	209284	0.418
1973	12.101	44.732	96.692	78.844	7.952	0.375	211357	0.444
1974	12.853	31.332	97.004	73.857	7.709	0.365	213342	0.493
1975	13.691	49.711	108.654	89.378	7.464	0.398	215465	0.538
1976	14.529	77.690	135.078	116.005	7.163	0.427	217563	0.569
1977	16.239	83.426	134.631	109.883	7.067	0.444	219760	0.606
1978	15.948	110.466	179.964	167.732	7.069	0.465	222095	0.652
1979	19.144	122.060	237.990	211.525	7.137	0.474	224567	0.726
1980	22.256	120.370	252.695	226.367	7.224	0.476	227225	0.824
1981	22.747	139.172	238.063	226.148	7.346	0.502	229466	0.909
1982	25.643	115.962	257.707	237.892	7.420	0.527	231664	0.965
1983	27.302	148.523	321.234	214.831	7.489	0.546	233792	0.996
1984	4.956	195.280	362.288	244.390	7.487	0.555	235825	1.039
1985	23.056	187.830	384.472	239.620	7.482	0.561	237924	1.076
1986	55.795	186.050	392.375	231.370	7.464	0.578	240133	1.096
1987	54.969	211.660	389.182	215.830	7.435	0.595	242289	1.136
1988	54.844	229.300	457.548	282.690	7.426	0.608	244499	1.183
1989	59.867	246.159	428.224	317.255	7.483	0.599	246819	1.240
1990	28.369	262.460	497.875	340.450	7.542	0.591	249402	1.307
1991	31.653	239.874	477.846	264.986	7.599	0.589	252131	1.362
1992	28.882	228.528	470.847	253.787	7.637	0.615	255028	1.403
1993	72.954	224.990	434.422	260.885	7.627	0.649	257783	1.445
1994	78.969	266.681	462.122	280.454	7.571	0.666	260341	1.482

Note: The subscripts are defined as follows: 1 = fluid milk, 2 = juices, 3 = soft drinks, and 4 = coffee and tea. See appendix narrative for sources and explanatory notes.

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