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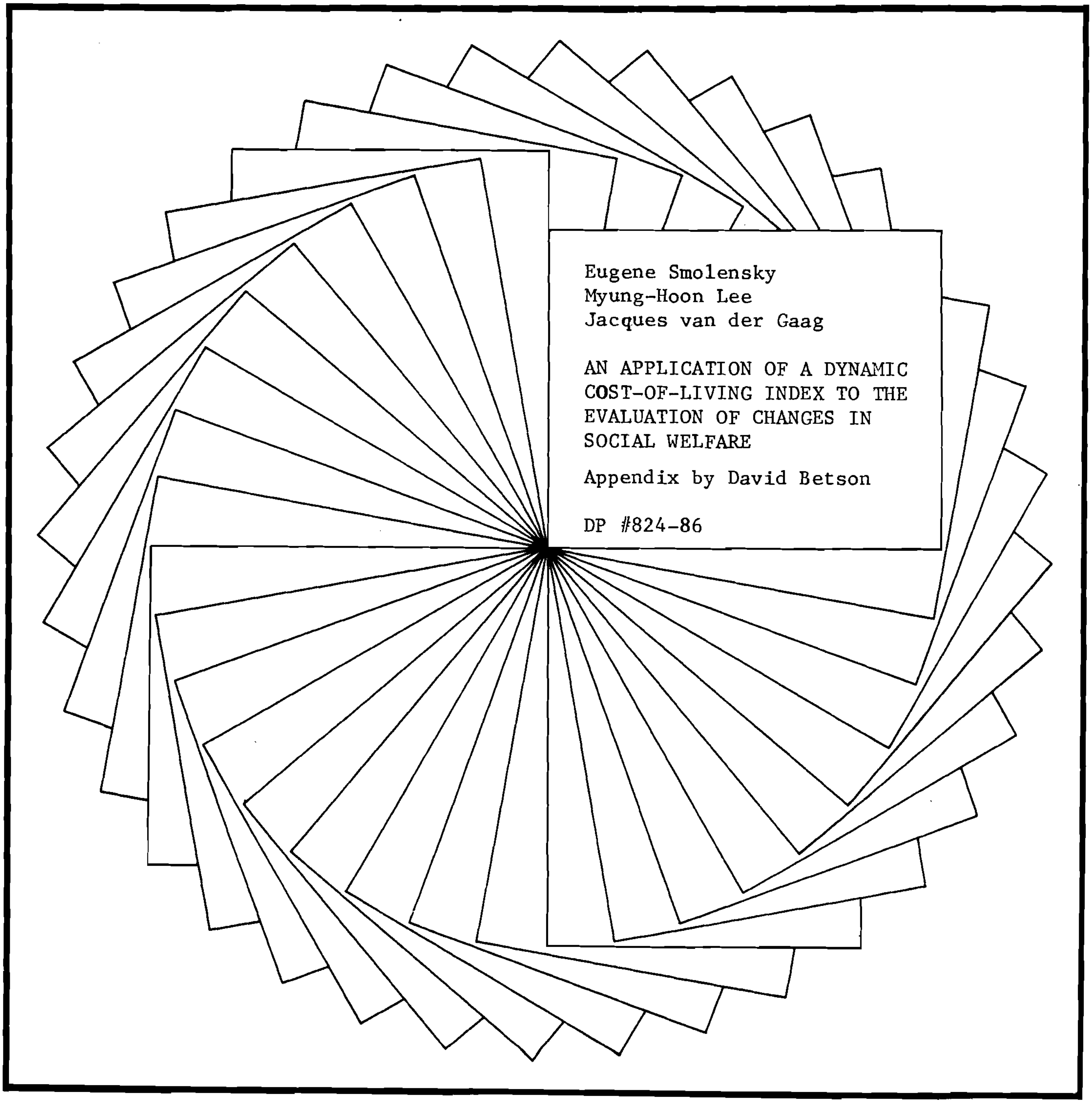

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AN APPLICATION OF A DYNAMIC COST-OF-LIVING INDEX TO THE EVALUATION OF CHANGES IN SOCIAL WELFARE

Appendix by David Betson

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Appendix 2: A Nonparametric Test of Taste Changes during the 1960s

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Abstract

We start with two observations that constitute an apparent paradox. Increasing real income (i.e., CPI-corrected income) in the U.S. has not been accompanied by increasing public satisfaction since at least 1960. We then add a third observation--large unexplained shifts in demand functions. These shifts in demand, when translated into a real welfare index, support the widely perceived malaise of the recent past. Since we have not explained the shifts in demand patterns, we do not have an explanation, in any sense, of the malaise. We do offer some conjectures which could be tested were the models by which we measure real income changes made more comprehensive.

Our empirical results were obtained by estimating a complete demand system after adding price data to a pooled sample of households drawn from two surveys of consumer expenditures taken about a decade apart--early 1960s and early 1970s. By adding the price data, it became possible for us to do what earlier studies could not do: derive true cost-of-living indices from cross-section data.

Since our results imply substantial taste changes between 1960-61 and 1972-73, we derive and implement a nonparametric test for taste changes as an appendix (Appendix 2). The test provides some support for the view that tastes did, in fact, change during the 1960s.

An Application of a Dynamic Cost-of-Living Index to the Evaluation of Changes in Social Welfare

After the oil embargo of 1973, a small body of literature developed around the paradoxical question, "Why do Americans seem to think that things are so bad when the data indicate that real income is continuing to rise?" That literature has recently been summarized by Gottschalk and Maloney (1983). They find that, although by some measures real economic well-being has declined, more careful measures indicate (as Thurow (1980) and Adam Smith (1981) maintained) that contrary to public opinion, real per capita disposable income continued to rise after the embargo, albeit more slowly than in preceding years. Gottschalk and Maloney (henceforth G-M) conclude, therefore, that there is indeed a paradox to be explained. The explanation they offer is that averages mask diversity, and that since 1973 there has been a very substantial increase in the proportion of families whose real incomes have declined. Between 1973 and 1979, 42 percent of families in the G-M panel (the Michigan Panel Study of Income Dynamics) experienced declining real income, up from 25 percent in the period 1967 to 1973. That so substantial a minority of households experienced an income decline seems sufficient to explain the sense of dissatisfaction that emerged after the 1973 oil embargo. It is not, however, a complete explanation. G-M admit of two other factors of some significance. They find some, but very modest, support for Levy's (1982) argument that real money income of families was sustained by the entry of wives into the workplace, and that therefore our usual measures overstate the rate of growth by excluding the concomitant decline in leisure and home production. G-M also grant that money illusion associated with inflation, Thurow's explanation, may have some validity as well.

It is difficult to remember now that the U.S. rate of inflation was high enough to become a matter of considerable public concern even before the 1973 oil embargo. Public dismay pushed a very conservative though pragmatic president into a wage/price freeze in August of 1971. An assessment in the Economic Report of the President for 1973 captures the sentiment at the time: "...the big increase in production in the year just ended was accompanied by a reduced rate of inflation. Consumer prices increased a little more than 3 percent from 1971 to 1972--a far cry from the runaway inflation rate of 6 percent that confronted us in 1969" (U.S. Council of Economic Advisers, 1973, p. 3; emphasis added). As the University of Michigan Survey Research Center data pictured in Figure 1 show, the proportion of households expecting to be worse off in the coming year hit a local peak in 1970. The percentage of families who believed they would be better off in the coming year reached a postwar peak as early as 1960. A decade and a half of strong growth was not translated into public perceptions of rising well-being. In this regard, the paradox chronicled by G-M for the period after the oil embargo merely extended a long-standing trend.[1]

In this paper we show that the subjective feeling of dissatisfaction has a behavioral counterpart: during the period 1960-73, household demand patterns experienced a remarkable shift that cannot be explained by the changes in income, relative prices, and demographic factors that took place during the same period. More important, we show how this shift in demand, suitably translated into a consumption-based dynamic welfare index, helps to explain the apparent paradox. We show that real income was in fact declining for the vast majority of households between 1960 and 1973.

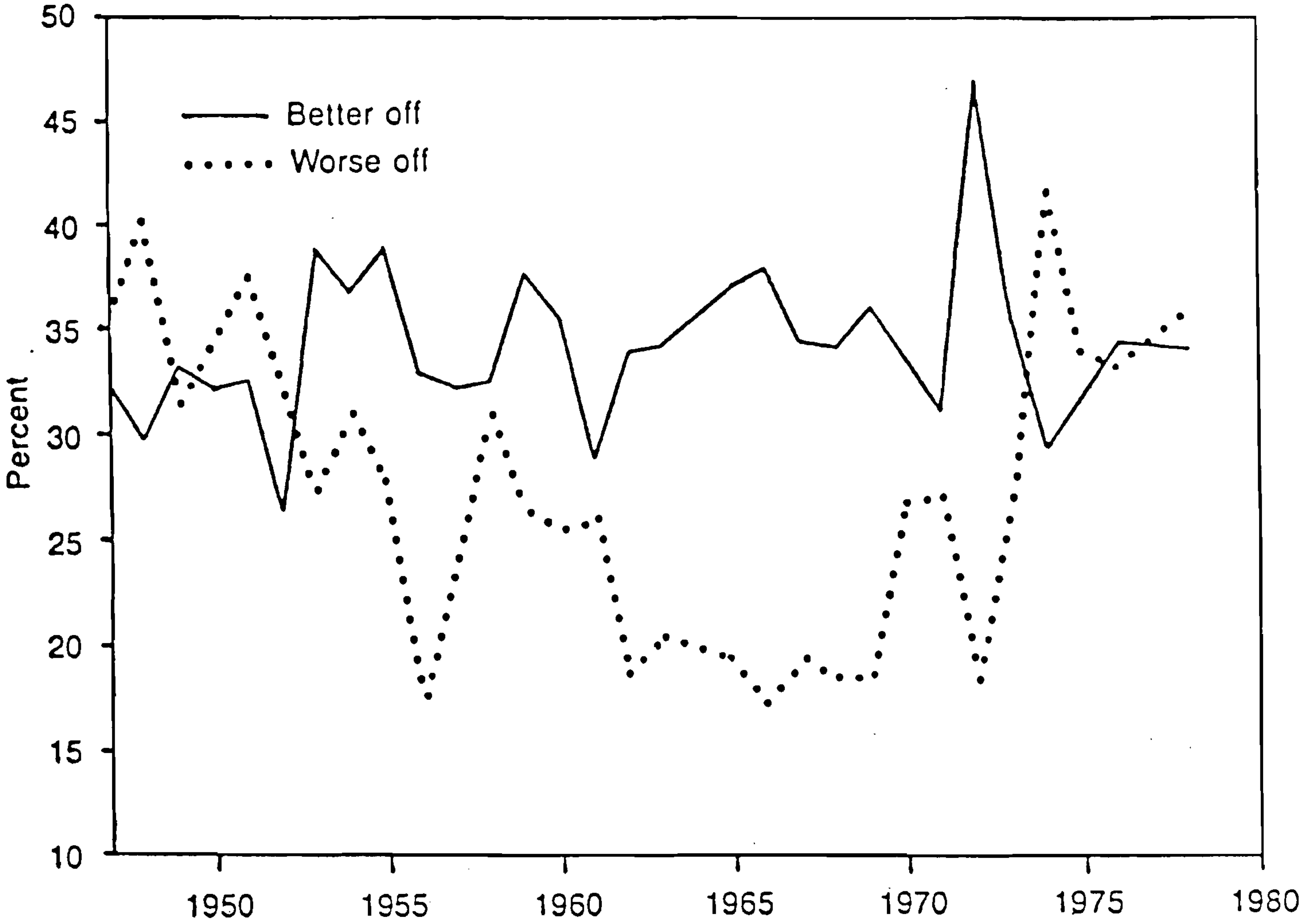
50
45
40
35
30
25
20
15
10
Percent
1950
1955
1960
1965
1970
1975
1980
Better off
Worse off

Figure 1

Prior studies of the course of real income over the period rely on the CPI to convert nominal to real income changes. These years place a heavy burden on this index. The inflation of this period was not the classical textbook case during which prices of all commodities rise proportionately in response to a helicopter money drop. From 1960-61 through 1972-73, the cost of food for our sample rose by 145 percent while transportation costs rose by 130 percent. It is also true, of course, that consumption patterns vary across household types and over time. The highest income decile in 1960-61 spent 21 percent of income on food and 12 percent on shelter; these shares in the lowest decile were 34 percent and 27 percent, respectively. In 1972-73 these shares were 17 and 19, and 29 and 27 percent, respectively. The food share for single women under 35 in 1972-73 was 14 percent! Obviously the price rises even in the decade before the 1973 oil embargo could have had substantially different consequences for the economic well-being of households of different income, size, age, and sex compositions.

There is a small literature on the effects of rising prices during the later sixties and early seventies on the level and distribution of real income. Some of these studies take fixed budget shares which vary across different classes of households as weights and obtain the weighted average of price indices of individual commodities (e.g., Duffy et al., 1980). Others have calculated true cost-of-living indices, that is, the ratio of the minimum cost of attaining a reference utility index with base period prices and comparison period prices (Braithwait, 1980; Palmer and Barth, 1977). In almost every instance these studies found that the distributional effects of rising prices were minimal and that real income rose in the same pattern as the simple deflated aggregate data indicate.

When we confine ourselves to standard static techniques we obtain the same result. A dynamic welfare index in which time is permitted to affect the estimates of the underlying demand equations yields quite different results. However, we do not find that inflation per se is directly responsible for the perceived decline in economic well-being.

In what follows we first describe the complete demand system which we estimated after adding price data to a pooled sample of households drawn from two surveys of consumer expenditures taken about a decade apart--early 1960s and early 1970s. By adding the price data, it became possible for us to do what earlier studies could not do: derive true cost-of-living indices from cross-section data. We then describe the cost-of-living indices we implemented and briefly present our empirical results.

Our key finding is that there was a large shift in the demand functions between 1960-61 and 1972-73. The fourteen time dummies in the model are statistically significant. In addition, as is evident in Appendix 1, there was a substantial increase in the marginal propensity to spend on shelter. The implication of these shifts, when taken through a particular cardinal index, is that real welfare for the average household declines in the vast majority of household types recognized by the model. Since until very recently cardinality has been explicitly rejected by most economists, we then briefly allude to the newly emerging literature on the cardinality vs. ordinality debate. We do not, however, attempt to solve the severe technical and conceptual problems associated with cardinal indexes.

We began by merging price information from the CPI and the worker budgets developed by the Bureau of Labor Statistics, with expenditure data from the Bureau of Labor Statistics Consumer Expenditure Surveys of 1960-61 and 1972-73.[2] This new data set was used to estimate the following Linear Expenditure System (LES):[3]

$$q_{it} = \gamma_{it} + (\beta_{it} / P_{it}) (Y_t - \sum_{k=1}^{N} P_{kt} \gamma_{kt}) \qquad i = 1, N; \; t = 0, 1$$

$$\beta_{it} = B_i + \mu_i T$$

$$\gamma_{it} = A_i + \delta_i' h + \pi_i T + \varepsilon_{it}$$

where q_{it} is the quantity of good i consumed in year t, i = 1 . . . 7;

P_{it} is the price of good i in year t;

Y_t is income (= total consumption expenditures) in year t;

h is a vector of household characteristics;

T is a dummy variable (T = 0 in 1960-61; T = 1 in 1972-73);

β_{it}, γ_{it}, B_i, A_i, δ_i, μ_i, and π_i are parameters; and

ε_{it} is a disturbance term.

The unit of observation is the household; household subscripts have been suppressed.

Assuming normality of the disturbance term ε_{it}, we obtain the FIML-estimates of the 131 parameters, using observations on 2432 households in 1960-61 and 2543 in 1972-73. (The households in the two periods are not the same households.) Total consumption is divided into seven consumption goods. The vector h includes fourteen household characteristics.[4] The dummy variable T tests whether preferences have remained stable over the period considered. From the estimation results, we calculate true cost-of-living indices for each household type over the period and compare them with the standard Laspeyres index.

To calculate the true cost-of-living indices we proceed as follows. The fitted demand equations specify the optimal demand for each good as a function of income, all prices, and household characteristics. Substituting these optimal quantities back into the utility function and inverting, we find the minimum expenditure or cost function. The implied utility function is of the standard Stone-Geary form:

$$U_{it} = \Sigma \ \beta_{it} \ \mathrm{Log}(q_{it} - \gamma_{it}).$$

The true cost-of-living index for household type h (J^h) is:

$$J^h = \frac{C^h(\bar{U}_o^h, \ P73^h)}{C^h(\bar{U}_o^h, \ P60^h)}$$

where $\bar{U}^h$ is a prespecified welfare level for household type h;[5]

C^h is the cost function of household type h; and

$P60^h$, $P73^h$ are vectors of prices in 1960-61 and 1972-73, respectively, faced by household type h.

Thus, J^h is the ratio of expenditures needed in 1972-73 to expenditures needed in 1960-61 for household type h to attain the same welfare level $\bar{U}_o^h$.

We assume that all households of the same type have the same utility function, but we allow these functions to shift over time. Indeed, we find the time dummies to be statistically significant and large (see Appendix 1); hence we make use of two alternative dynamic cost-of-living indices--the Fisher-Shell index (FS^h) (1969) and the cardinal index (CI^h). They are defined as follows for household type h, where R60 and R73 represent the preference orderings before and after they change, and $\bar{U}f^h$ and $\bar{U}c^h$ are prespecified welfare levels:

$$FS^h = \frac{C^h(\bar{Uf}^h, P73^h, R73)}{C^h(\bar{Uf}^h, P60^h, R73)}$$

$$CI^h = \frac{C^h(\bar{Uc}^h, P73^h, R73)}{C^h(\bar{Uc}^h, P60^h, R60)}.$$

For the mean income of household type h, $Y60^h$, $\bar{Uf}^h$ is determined at the welfare level attained at $Y60^h$ given $P60^h$ and R73, while $\bar{Uc}^h$ is determined as the welfare level attainable at $Y60^h$ given $P60^h$ and R60.

FINDINGS

As Table 1 indicates, total consumption expenditures grew, in nominal terms, by no less than 17 percent and by as much as 118 percent for the different demographic groups over the period. In Table 2 this growth is deflated by a Laspeyres index which is specific to each household type and calculated from our data. The variation across demographic groups of the Laspeyres index (not shown) is small (137.1 to 138.8) and leaves the rank ordering of groups in Table 1 virtually unchanged. Not all groups, however, now experience a gain in real income. For five groups (roughly 6 percent of the sample) average incomes decline (a ratio less than 1), and the gain for most groups becomes quite modest. The Fisher-Shell index, as a comparison of Tables 2 and 3 indicates, yields just about the same outcome as the Laspeyres index. The cardinal index (Table 4), on the other hand, entirely changes our conception of what happened over the period. All but two groups suffer declining real income, and for some (especially single men and women over 55) the decline is truly enormous.[6]

Table 1

Ratios of Nominal Total Consumption Levels, 1972-73/1960-61

Household Composition	Age of Household Head Below 35	35-55	56-65	Over 65
One Person				
Male	2.18	1.79	1.30	1.43
Female	2.11	1.51	1.57	1.44
Two Persons				
Couple	1.83	1.59	1.47	1.55
Three Persons				
Couple, 1 child				
< 6	1.54	1.71		
6-11	1.30	1.77		
12-17		1.35	1.17	
18+		1.45	1.46	1.50
Four Persons				
Couple, 2 children				
< 6	1.37	1.30		
6-11, < 6	1.50	1.65		
6-11	1.49	1.66		
12-17, 6-11		1.44		
12-17		1.47		
18+, 6-17		1.39	1.43	
18		1.60	1.69	
Five Persons				
Couple, 3 children				
6-11, < 6	1.63	1.56		
12-17, 6-11		1.43		
18+, 6-17		1.41		

Source: Calculated by the authors from samples drawn from the Bureau of Labor Statistics Consumer Survey Expenditures of 1960-61 and 1972-73. No ratio is imputed for those cells in which sample size was less than 19.

Table 2

Ratio of Real Total Consumption Levels, 1972-73/1960-61, Deflated by Laspeyres Index

Household Composition	Age of Household Head Below 35	35-55	56-65	Over 65
One Person				
Male	1.59	1.31	.94	1.04
Female	1.54	1.11	1.15	1.05
Two Persons				
Couple	1.33	1.16	1.07	1.13
Three Persons				
Couple, 1 child				
< 6	1.13	1.25		
6-11	.94	1.29		
12-17		.98	.85	
18+		1.06	1.06	1.08
Four Persons				
Couple, 2 children				
< 6	1.00	.95		
6-11, < 6	1.09	1.20		
6-11	1.08	1.20		
12-17, 6-11		1.04		
12-17		1.06		
18+, 6-17		1.00	1.03	
18+		1.16	1.22	
Five Persons				
Couple, 3 children				
6-11, < 6	1.18	1.13		
12-17, 6-11		1.03		
18+, 6-17		1.01		

Source: See Table 1.

Table 3

Ratios of Real Total Consumption Levels, 1972-73/1960-61, Deflated by Fisher-Shell Index

Household Composition	Age of Household Head Below 35	35-55	56-65	Over 65
One Person				
Male	1.61	1.32	.95	1.05
Female	1.57	1.12	1.16	1.07
Two Persons				
Couple	1.35	1.18	1.08	1.14
Three Persons				
Couple, child				
< 6	1.14	1.26		
6-11	.96	1.30		
12-17		.99	.86	
18+		1.07	1.08	1.09
Four Persons				
Couple, 2 children				
< 6	1.01	.96		
6-11, < 6	1.10	1.21		
6-11	1.09	1.21		
12-17, 6-11		1.05		
12-17		1.07		
18+,. 6-17		1.02	1.04	
18+		1.18	1.24	
Five Persons				
Couple, 3 children				
6-11, < 6	1.20	1.15		
12-17, 6-11		1.04		
18+, 6-17		1.03		

Source: See Table 1.

Table 4

Ratios of Real Total Consumption Levels, 1972-73/1960-61 Deflated by Cardinal Index

Household Composition	Age of Household Head			
	Below 35	35-55	56-65	Over 65
One Person				
Male	107.0	80.1	54.7	49.3
Female	95.6	66.5	61.9	52.2
Two Persons				
Couple	99.4	85.0	75.1	70.2
Three Persons				
Couple, 1 child	81.7	91.3		
< 6	73.8	98.1		
6-11		76.9	66.8	
12-17		84.2	83.8	76.8
18+				
Four Persons				
Couple, 2 children				
< 6	74.0	69.8		
6-11, < 6	82.8	90.0		
6-11	82.0	90.5		
12-17, 6-11		82.3		
12-17		83.6		
18+, 6-17		83.7	83.7	
18+		100.1	97.1	
Five Persons				
Couple, 3 children				
6-11, < 6	88.2	86.9		
12-17, 6-11		83.7		
18+, 6-17		86.3		

Source: See Table 1.

The source of the different results from the two indices is perhaps best explained by an example. At the 1960-61 group average income of $3902, our estimates yield a utility level for a single male aged 35-55 of 1.08 in 1972-73 but 2.15 in 1960-61. Fisher and Shell (1969) argue that the difference between 2.15 and 1.08 is without "operational content." Their index, which evaluates the change in relative prices, is essentially concerned with the curvature of the 1.08 indifference curve. The cardinal index, in contrast, takes as entirely pertinent the measured decline in utility for two different single men, aged 35-55, separated in time by twelve years. In this instance, the effect of the drop in cardinal-measured utility far outweighs the welfare loss due to changes in relative prices over the period.

DISCUSSION

It has recently, and suddenly, become quite acceptable for economists to discuss taste changes, even though Marschak's (1978) judgment that "To enter the field of taste changes one ought to find danger exhilarating" still holds. Some of this literature is undoubtedly a reaction to the effort by Stigler and Becker (1977) to rid economics of taste changes altogether. (Pollak, 1978, appears to fall into this category.) A part of the literature stems from Scitovsky's (1976) innovative book. Much of it, however, rests on the repeated finding that when time is permitted to affect estimated demand functions, the effect, as in our case, is large (e.g., Darrough, Pollak, and Wales, 1983). Most important, it is obvious that tastes do change and it is important for predictive and for policy purposes to theorize about those changes (Hirschman, 1982). Even the fact that individuals hold to different preferences at different points in time is important for policy (Schelling, 1984). Despite all this activity on the dynamics of demand, only Phlips (1974) has contested the blanket condemnation by Fisher and Shell of the cardinal index calculated here. However, Fisher and Shell's basic point, "the comparison between a man's utility now and his utility yesterday stands on precisely the same lack of footing as the comparisons of the utilities of two different men" (p. 99), is by necessity being ignored or circumvented all across the subspecialities of economics.[7] We need equivalence scales to hand out benefits and to assess taxes. We need true cost-of-living indexes so as to index those same benefits and taxes. We have attacked the problem directly; others have resorted to appeal to a social welfare function.

At issue here is just how we ought to model an individual's perception of his own economic well-being over time. When a respondent to a Gallup poll says that he expects to be just as well off next year as this, does he mean that he will be able to buy the same bundle of goods he buys today? A bundle of goods that yields equivalent utility to today's bundle today? Does he care that, on average, his peers will be able to spend 2 percent more next year? We need to formally model and test the perception process. All we have done here is to suggest, by example, that if we proceed down that road, we may better understand the economic environment and better evaluate economic policies.

It also needs to be said, however, that the index implemented here has little to commend it other than that it is a natural and simple extension of everyday practice which sheds some light on a small puzzle.[8] It suffers from the technical problem that the utility function is not unique to the demand equations we estimated. We did not try to solve the technical and conceptual problems inherent in the use of a cardinal welfare index. We did present an observation on consumption behavior that may help to explain an apparent paradox. This observation, an unexplained shift in demand patterns (which is quite common in consumption analysis), should facilitate the search for underlying causes behind the feeling of malaise in the seventies.

To recapitulate, we started with two observations that constitute an apparent paradox. Increasing real income (i.e., CPI-corrected income) has not been accompanied by increasing public satisfaction. We then added a third observation--large unexplained shifts in demand functions. These shifts in demand, when translated into a real welfare index, support the subjectively measured but widely perceived malaise of the recent past. Since we have not explained the shifts in demand patterns, we do not have an explanation, in any sense, of the malaise. The arguments assembled by Gottschalk and Maloney to explain malaise in the post-embargo period are all candidates for explaining the earlier period as well. If panel data were available for the earlier period, it seems plausible that we would find an increasing tendency for the proportion of individuals with year-to-year declines in income to have grown once the pattern of brief cycles set in, during the late fifties. (The interruption of that pattern during the Vietnam war raises some doubts about our ability to generalize from external events, however.) The role of the entry of wives into the labor force, emphasized by Levy, as well as more general changes in labor supply, is supported by the work of Barnett (1981). He finds that by jointly treating labor supply and commodity demands in a complete demand system (the Rotterdam model), "unexplained exogenous time trends in preferences are found not to exist" (p. 7). Since our own results are so substantially affected by an increase in the marginal propensity to spend on housing (see Appendix 1), it may be desirable to add the interdependence of consumption and portfolio choices to Barnett's consumption-leisure interaction in the complete demand system. It is plausible to believe that our dynamic linear expenditure system picked up the shift to landholding and to durables such as housing which is characteristic of inflationary periods. The message of this paper, then, is this: Our capacity to measure real income changes is surprisingly limited. We need to substantially enlarge our models.

Notes

[1] Easterlin's (1974) well-known investigations into the time-series relationship between economic growth and happiness end in 1970, but one of his conclusions is nevertheless pertinent: "Between 1966 and 1970, however, all income classes show a noticeable decline [in the percentage very happy], and in 1970 there is no class which is higher [happier] than it was in 1963" (p. 111).

[2] See Lee (1982) for more detail on the data set and methodology.

[3] We chose the LES for its convenience. First, it economizes on the number of parameters to be estimated. Second, despite its limitations, the LES has been widely estimated on both cross-section and time-series data. This facilitated comparison of our estimation results with those in the literature. The income and price elasticities of our demand equations are well within the range obtained by others. For direct comparisons, see Lee (1982). Estimating the Almost Ideal Demand System (Deaton and Muellbauer, 1980) on the same data set yields demand equations qualitatively similar to those of the LES system.

[4] In accordance with the Barten (1964) model for incorporating household characteristics, we do not allow the β's to vary across households. Though demographic variables and time could have been interacted, the simplifying assumption of constant demographics over the period was made to keep the number of parameters manageable. With seven goods, fourteen household characteristics (see Appendix 1), and seven error terms, our model already contains 131 parameters. Allowing for the interaction between demographic attributes and time would add 98 parameters.

[5] The prespecified welfare level is a function of income, and hence the true cost-of-living index also depends on income. The index is less dispersed across household types as the prespecified level of welfare rises, but the qualitative results are not affected.

[6] Tables 2 and 3 are highly correlated with Table 1: Spearman's Rho equals .998 and .999 respectively. The correlation of Table 1 with Table 4, however, is only .678.

[7] The origins of this tension between spinning an elegant theory of markets and serving the goal of improving welfare is described by Cooter and Rappoport (1984) and by Roy (1984). We find our own position, or positions, to have been perfectly captured by Roy when he says, "The battle lines were drawn: ... you were either against or in favour of such weighty matters of State as progressive income taxation and other 'egalitarian' measures in general. Roughly speaking, demand theorists aligned themselves on one side and welfare economists on the other; while if you were a demand theorist sometimes and a welfare economist at other times, you might throw consistency to the winds and happily be schizophrenic" (p. 361).

[8] Actually, as indicated earlier in the text, our index undoubtedly explains too much, since almost everyone's real income declines. Examination of individual histories with panel data, as in the G-M study, would probably reveal that a larger proportion of the population experienced real and substantial measured gains over the period.

Appendix 1

Table A.1 shows 124 of the 131 parameters we estimated. In addition to the seven estimates of the disturbance term (ε_i) that are not reported here, we estimated six B_i, six μ_i, seven A_i, seven π_i, and 98 δ (7 categories × 14 demographic variables). B_i and μ_i for the category "Others" were derived by using the budget constraint.

Estimates of both time dummies, μ_i (on the marginal propensities to spend) and π_i (on the intercepts), are statistically highly significant as implied by the coefficient/standard error ratios in parentheses, and all π_i are positive except for shelter. During the same period, β_i for shelter increased by 0.36, from 0.11 to 0.47. To compensate for this large increase, the β_i for all other categories decreased during the period ($\Sigma\ \beta_i = 1$). The largest drop in β_i was for transportation. The direction and magnitude of changes in β_i have important implications for the implied income elasticities. For instance, shelter was a necessity in 1960-61 but becomes a luxury good in 1972-73, while the reverse was the case for transportation.

Table A.1

Parameter Estimates of the Demand System

Coefficients	Food	Shelter	Energy	Home Furnishings	Clothing	Transp.	Others
1. B_i (Constant)	.14	.11	.02	.13	.11	.26	.23
	(113.93)	(86.48)	(21.77)	(85.91)	(120.70)	(103.20)	NA
2. μ_i (Time dummy)	-.08	.36	-.01	-.04	-.04	-.11	-.08
	(-55.41)	(269.63)	(-10.80)	(-21.92)	(-39.91)	(-37.49)	NA
3. A_i (Constant)	1.48	1.62	.03*	-2.86	-2.07	-6.78	-1.72
	(2.67)	(3.61)	(.20)	(-5.11)	(-5.10)	(-6.38)	(-2.14)
4. π_i (Time dummy)	6.02	-4.38	2.32	5.43	3.30	18.00	7.38
	(11.09)	(-5.98)	(20.70)	(9.77)	(8.34)	(17.81)	(9.28)
5. Log Household Size	6.10	.10*	1.82	1.16	1.81	2.57	2.65
	(27.85)	(.36)	(19.86)	(5.23)	(12.62)	(4.79)	(8.72)
6. Sex of Household Head	-2.99	-.96	-.33	-1.34	-.77	-6.29	-4.24
	(-15.26)	(-3.58)	(-4.90)	(-7.66)	(-5.91)	(-12.64)	(-16.06)
<u>Children's Ages</u>							
7. Under 6	-1.85	2.00	.01*	2.73	-.48	-2.67	-1.12
	(-6.99)	(5.53)	(.08)	(11.48)	(-2.64)	(-4.09)	(-3.01)
8. 6-11, under 6	-.08*	1.95	.02*	1.43	.06*	-3.34	-1.53
	(-.23)	(4.71)	(.15)	(4.37)	(.26)	(-4.03)	(-2.99)
9. 6-11	.67	1.11	-.30	.82	.54	-2.24	.93
	(1.74)	(2.54)	(-1.94)	(2.86)	(2.43)	(-2.58)	(2.07)
10. 12-17, under 6	-.18*	1.55	-.57	.58*	1.03	-4.54	-1.03
	(.36)	(3.01)	(-2.97)	(1.28)	(3.75)	(-3.56)	(-1.72)
11. 12-17, 6-11	.69	.02*	.43	.25*	1.15	-6.71	-1.54
	(1.90)	(.06)	(-2.90)	(.74)	(5.18)	(-7.60)	(-3.07)
12. 12-17	1.83	.16*	.02*	.01*	1.70	-.12*	.27*
	(5.88)	(.39)	(.17)	(.03)	(8.54)	(-1.59)	(.62)
13. Over 18, under 6	1.04*	1.13*	-.70	-.81*	1.99	-1.68*	-.29*
	(1.13)	(1.45)	(-2.21)	(-.91)	(4.54)	(-.85)	(-.28)

—table continues—

Table A.1, continued

Coefficients	Food	Shelter	Energy	Home Furnishings	Clothing	Transp.	Others
14. Over 18, 6-17	4.19	1.04	-.18*	.08*	2.95	3.15	4.72
	(12.14)	(2.50)	(-1.26)	(.24)	(14.01)	(4.08)	(12.19)
15. Over 18	2.36	.17*	-.06*	.64	2.09	5.38	4.12
	(8.77)	(.56)	(-.56)	(2.70)	(13.05)	(9.34)	(13.20)
Household Head's Age							
16. Under 35	-3.01	-1.03	-1.11	-1.21	-.72	-.96	-2.34
	(-17.04)	(-4.33)	(-15.44)	(-8.44)	(-5.86)	(-2.28)	(-10.29)
17. 55-65	.30*	-.21*	.21	.09*	-.08*	-2.10	-.23*
	(1.60)	(-.91)	(2.52)	(.56)	(-.66)	(-4.69)	(-1.01)
18. Over 65	-1.19	-.49	-.01	-1.02	-1.52	-5.59	-2.04
	(-5.37)	(-1.84)	(-.11)	(-4.98)	(-10.83)	(-10.56)	(-7.14)

Notes: Estimate/Standard Deviation Ratio in parentheses.

"Others" column is derived by using the "adding-up" condition ($\Sigma \beta_{it} = 1$). Total number of estimated parameters is 18 × 7 - 2 = 124, or (number of rows) × (number of commodity categories) - (derived coefficients) = 124. Additional estimates of disturbance term (ε_i) are not reported here.

NA = Not available.

*Not significant at 5% level.

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Appendix 2: A Nonparametric Test of Taste Changes during the 1960s

David Betson

The purpose of this short note is to report upon the results of a direct nonparametric test of the hypothesis that the tastes of U.S. consumers changed during the 1960s.

STATISTICAL MOTIVATION FOR THE HYPOTHESIS TEST

Assume that we have longitudinal data available on an individual's expenditures over 1000 observation periods. We wish to test the hypothesis that the individual's preferences in the first 500 periods differed from those of the second 500 periods. How should this conjecture be tested?

One possible strategy would be to adopt a specific parametric formulation of the individual's preferences. For example, they could be represented by the Linear Expenditure System (LES) as was done in this paper. Once a specific parameterization of preferences is chosen, the preference ordering can be characterized by a set of coefficients estimated from the available data, as in Appendix 1. The question, "Do the individuals' preferences differ in the two subperiods?" would thereby be transformed into asking whether the coefficients estimated from the two subsamples are different. In the body of the paper, this purpose is served by explicitly making use of time dummies in pooled cross-section data. The statistically significant coefficients in the two time dummies constitute evidence that preferences changed.[1]

This approach could be called appropriately "the parametric strategy." By adopting it, however, we also implicitly adopt an additional maintained hypothesis: a specific parameterization of preferences. Specifying a particular demand system increases the probability of accepting the hypothesis that preferences changed even when they did not. In our example, the LES parameterization of preferences may not represent the individual's true preferences. The possibility would remain then that the differences in coefficients result from errors in the approximation of the individual's true preferences due to the functional form chosen. Experimenting with alternative parameterizations of the individual's preferences could increase confidence in the finding that preferences did change, but the possibility that preferences may not have changed could never be ruled out.

A test that does not rely upon a specific parameterization of preferences would reduce the chance of committing the above Type I error. Samuelson (1948) demonstrates that if a finite set of price/expenditure data does not violate a set of binary relationships known as the Strong Axiom of Revealed Preference (SARP), then there exists a set of preferences that rationalize the data. That is, there would exist a utility function which when maximized subject to a given set of prices and income would yield the expenditure pattern found in the data. Hence if the price/expenditure data do not violate SARP in either the first or second sample of 500 observations, then there must exist a utility function which would rationalize the data in each subsample. In this case, the test of the hypothesis of a change in preferences becomes, "Does the whole sample of 1000 observations obey SARP?" If it does, the hypothesis that preferences have remained constant over the period cannot be rejected. However, if the complete sample violates SARP, then it follows that preferences must have changed between the two periods.

This extended example serves to motivate the strategy that underlies the results we present in this note. Since the test we will utilize does not require a specific parameterization of preferences, we will refer to it as a nonparametric test of the changing taste hypothesis.

Some further comments on this strategy are in order. It has already been noted that the primary advantage of the nonparametric approach is that it reduces the likelihood of committing a Type I error--that is, rejecting the true hypothesis that tastes have not changed. However this strategy is open to a Type II error; accepting the constant tastes hypothesis when tastes have changed. SARP violations will indicate that a given set of data cannot be rationalized by any utility function. In the absence of violations of SARP, however, there is no guarantee that the preference structures which rationalize each of the three samples of our example are the same. Finally, testing a sample for violations of SARP and finding none is not conclusive proof that the individual was maximizing a given set of well-behaved preferences. All that can be learned from this test is that the observations of the individual's behavior were not inconsistent with utility maximization. It could be the case that with more observations, we would find some inconsistencies.

Landsburg (1981), using this nonparametric research strategy with aggregate British data, found no evidence for taste changes during the period 1900 to 1955. To our knowledge, no one has employed this strategy to test for taste changes using individual household data.

TWO OTHER AXIOMS FROM REVEALED PREFERENCE

Reference has been made to a set of conditions known as the Strong Axiom of Revealed Preference (SARP). The purpose of this section is to define what is contained in these conditions and in two other sets of conditions denoted as the Weak Axiom of Revealed Preference (WARP) and the Generalized Axiom of Revealed Preference (GARP).

Recall that the principal goal of these conditions is to state the necessary and sufficient conditions for a set of price/expenditure data to be consistent with utility maximization. Denote any of the observations on prices and expenditures as the pair $(\underline{p},\underline{x})$, where $\underline{p}$ is a list of prices and $\underline{x}$ is the associated chosen consumption bundle. Assume that a utility function $U(\underline{x})$ rationalizes the data. Hence if $\underline{x}_i$ was chosen when the individual faced $\underline{p}_i$ prices and another bundle $\underline{x}_j$ could have been purchased at $\underline{p}_i$, then one must conclude that $\underline{x}_i$ was at least preferred to $\underline{x}_j$, i.e., $U(\underline{x}i) \geq U(\underline{x}_j)$. In this case where $\underline{p}_i\underline{x}_i \geq \underline{p}_i\underline{x}_j$, we say that $\underline{x}_i$ was directly revealed preferred to $\underline{x}_j$ and write $\underline{x}_i \; R^0 \; \underline{x}_j$. Clearly for the data to be consistent with a strictly concave utility function, $\underline{x}_j$ could not be directly revealed preferred to $\underline{x}_i$ if $\underline{x}_i$ was directly revealed preferred to $\underline{x}_j$ unless the two bundles were equal. This observation leads us to what is known as the Weak Axiom of Revealed Preference (WARP):

WARP: if $\underline{x}_i \; R^0 \; \underline{x}_j$ and $\underline{x}_i \neq \underline{x}_j$ implies it is not the case that $\underline{x}_j \; R^0 \; \underline{x}_i$.

To check whether these conditions are fulfilled by a set of price/expenditure data of N observations, one constructs a NxN matrix, C_j

where

$$c_{ij} = \begin{cases} 1 \text{ if } \underline{p}_i\underline{x}_i \geq \underline{p}_i\underline{x}_j, \text{ i.e., } \underline{x}_i\ R^0\ \underline{x}_j \\ 0 \text{ otherwise.} \end{cases}$$

WARP will be violated if the product $c_{ij}c_{ji}$ is equal to one for any i and j and $\underline{x}_i$ is not equal to $\underline{x}_j$.

Given the method by which these conditions were developed, it can easily be seen that WARP provides the necessary conditions for utility maximization. It can be shown that only in the case of two goods does WARP provide both the necessary and sufficient conditions for utility maximization. When there are more than two goods, it is possible to construct examples in which every binary comparison between bundles of goods is rational in the sense that it obeys WARP, yet the binary comparisons are not transitive. That is, it might be the case that $\underline{x}_1\ R^0\ \underline{x}_2\ R^0\ \underline{x}_3$ but $\underline{x}_3\ R^0\ \underline{x}_1$! This intransitive behavior would not be consistent with utility maximization subject to a strictly concave utility function. The sequences of binary comparisons are denoted as the "transitive closure" of R^0 and written as R. The Strong Axiom of Revealed Preference checks the data to see if there exist any intransitivities of the binary relationships expressed in R^0. In particular,

SARP: if $\underline{x}_i\ R\ \underline{x}_j$ and $\underline{x}_i \neq \underline{x}_j$ then it is not the case $\underline{x}_j\ R\ \underline{x}_i$.

Obviously constructing all possible sequences of binary relationships in order to check for the transitivity of the relationship R^0 can be quite difficult. However, Varian (1982) has shown, using the Warshall algorithm, which utilizes principles of graph theory, that the transitive closure of R^0 can be computed using the matrix $\underline{C}$ with less than N^3 calculations and comparisons. The output of this algorithm is a matrix $\underline{T}$ where

$$t_{ij} = \begin{cases} 1 \text{ if } \underline{x}_i \; R \; \underline{x}_j \\ 0 \text{ otherwise.} \end{cases}$$

The SARP conditions can be checked by computing the product $t_{ij}t_{ji}$. If the product is equal to 1 and $\underline{x}_i$ is not equal to $\underline{x}_j$, then there exists a SARP violation.

As noted above, SARP contains the necessary and sufficient conditions for there to exist a strictly concave utility function which rationalizes the data. However if the utility function is strictly concave, then the demand functions must be single valued. That is, for a given set of prices there will exist a unique bundle of goods that maximizes the individual's preferences. A weaker restriction on preferences would allow for multiple-valued demand functions. The necessary and sufficient conditions for utility maximization which allow for demand correspondences are known as the Generalized Axiom of Revealed Preference (GARP). Using the notation defined above,

GARP: if $\underline{x}_i$ R $\underline{x}_j$ implies that it is not the case that $\underline{p}_j\underline{x}_j > \underline{p}_j\underline{x}_i$.

EMPIRICAL CONSIDERATIONS

We have discussed testing the taste change hypothesis with data of a single individual's expenditures over a long period of time. Unfortunately, we do not have data of this kind. Instead, we have price/expenditure data for different individuals from two time periods. No individual is in the sample more than once.

One consequence of utilizing the 1960 and 1972 Consumer Expenditure Surveys (CEX) is that we must maintain another untestable hypothesis, namely, that within a time period all individuals have the same set of preferences. Obviously this is a rather strict and unrealistic assumption. To make this assumption a little more palatable, we formed 23 subsamples of the data, where each subsample represents a relatively homogeneous demographic group. See Table A2.1 for a description of the groupings, which closely resemble those in the body of the paper.

The data contain many instances in which exceedingly large shares of income are devoted to a single commodity group. For example, several households are reported to have devoted more than 90 percent of their total expenditures to transportation. To examine the potential effects of these "measurement" errors, a second sample was constructed by deleting any observation if, for any commodity, the share of spending out of income was in excess of two standard deviations of the average share for that commodity in that year. In Table A2.1, we report the number of observations by demographic group for the FULL sample and this TRUNCATED sample. On average, this rule reduced the 1960 sample by 29 percent and the 1972 sample by 33 percent.

RESULTS

Tables A2.2 and A2.3 present the results of testing the data for violations of WARP, SARP, and GARP in the FULL and TRUNCATED samples, respectively. For each demographic grouping within each sample, the

Table A2.1

Demographic Grouping of Data

Description	Number of Observations in			
	Full Sample		Truncated Sample	
	1960	1972	1960	1972
Head Under 35 Years Old				
1. Singles	152	346	89	182
2. Childless couples	189	290	143	193
3. Male head with all children under 6	564	420	487	301
4. Male head with children under and over 6	240	210	220	159
5. Male head with all children over 6	68	124	55	96
6. Female head with all children under 6	27	54	15	16
7. Female head with children under and over 6	24	36	18	12
8. Female head with all children over 6	26	42	20	17
Head 35-55 Years Old				
9. Singles	215	214	137	129
10. Childless couples	387	199	310	141
11. Male head with all children under 12	492	241	435	202
12. Male head with all children between 12 and 17	228	171	187	139

Table A2.1 (Continued)

Table A2.1 (Continued)

Description	Number of Observations in Full Sample 1960	Full Sample 1972	Truncated Sample 1960	Truncated Sample 1972
13. Male head with children under 12 and children between 12 and 17	416	282	350	235
14. Male head with children over and under 18	274	309	224	243
15. Male head with all children over 18	150	143	119	109
16. Female head with children	165	162	105	100
Head 56-65 Years Old				
17. Singles	172	208	104	117
18. Childless couples	383	319	295	221
19. Male head with all children under 18	128	84	107	67
20. Male head with all children over 18	113	130	91	96
21. Female head with children	48	33	31	27
Head 66 Years Old and Older				
22. Singles	303	499	138	213
23. Childless couples	407	385	291	249
Total Sample Size	5170	4902	3665	3264

Table A2.2

Violations Expressed as a Percentage of the Maximum Number of Possible Violations--Full Sample

	1960 Sample			1972 Sample			1960-72 Combined Sample		
Group	WARP	SARP	GARP	WARP	SARP	GARP	WARP	SARP	GARP
1	0.3	1.1	0.9	0.4	7.0	6.3	0.4	18.5	17.2
2	0.3	1.0	0.9	0.5	9.3	8.3	0.4	29.4	28.0
3	0.3	15.9	14.9	0.5	20.2	18.8	0.4	57.3	56.6
4	0.4	6.3	5.8	0.5	7.8	7.0	0.4	38.0	36.5
5	0.5	1.4	1.2	0.7	7.0	6.3	0.6	7.8	7.0
6	0.6	0.6	0.6	0.6	1.1	1.0	0.5	0.9	0.9
7	1.1	1.1	1.1	0.5	1.3	1.3	0.6	1.8	1.8
8	0.6	0.6	0.6	1.7	3.4	3.3	1.1	3.6	2.9
9	0.4	3.5	3.1	0.6	5.1	4.6	0.4	14.6	13.4
10	0.4	10.8	10.2	0.5	7.4	6.6	0.4	39.5	38.3
11	0.3	15.4	14.5	0.4	4.3	3.8	0.3	62.4	61.0
12	0.3	1.9	1.5	0.6	6.4	5.7	0.4	11.5	10.3
13	0.3	4.0	3.4	0.4	6.6	6.0	0.3	22.3	20.9
14	0.3	2.4	2.1	0.4	4.8	4.2	0.3	22.5	21.2
15	0.3	1.9	1.7	0.5	2.8	2.5	0.3	3.7	3.0
16	0.4	5.0	4.7	0.7	5.5	4.8	0.5	15.0	14.1
17	0.4	2.0	1.8	0.8	19.8	19.2	0.6	24.4	23.3
18	0.3	8.6	8.1	0.4	12.1	11.3	0.3	45.2	43.9
19	0.3	0.8	0.8	0.4	0.9	0.9	0.3	2.2	1.9
20	0.6	2.3	2.3	0.3	0.6	0.5	0.4	3.1	2.7
21	0.5	1.5	1.5	0.6	0.8	0.8	0.4	1.5	1.4
22	0.7	26.7	25.4	0.5	9.8	8.7	0.8	77.4	75.5
23	0.5	16.0	15.3	0.9	78.0	76.3	0.4	78.6	77.1

Table A2.3

Violations Expressed as a Percentage of the Maximum Number of Violations-- Truncated Sample

	1960 Sample			1972 Sample			1960-72 Combined Sample		
Group	WARP	SARP	GARP	WARP	SARP	GARP	WARP	SARP	GARP
1	0.3	0.4	0.4	0.3	1.5	1.2	0.2	4.0	3.3
2	0.3	0.6	0.6	0.5	6.8	6.1	0.4	10.3	9.1
3	0.3	13.1	12.1	0.4	11.2	10.1	0.3	43.4	41.6
4	0.4	3.6	3.2	0.5	3.5	2.9	0.4	6.6	5.7
5	0.6	1.1	1.1	0.6	4.0	3.6	0.6	4.7	4.2
6	--	--	--	--	--	--	0.4	0.9	0.9
7	0.7	0.7	0.7	--	--	--	0.7	0.9	0.9
8	--	--	--	--	--	--	0.3	0.6	0.6
9	0.3	1.8	1.7	0.3	1.5	1.4	0.2	2.5	2.0
10	0.4	6.0	5.6	0.4	2.0	1.7	0.3	23.4	22.5
11	0.3	8.1	7.4	0.4	3.3	2.9	0.2	24.8	23.5
12	0.3	1.6	1.4	0.5	2.4	2.2	0.3	4.9	4.1
13	0.3	3.4	2.9	0.3	4.7	4.3	0.3	17.6	16.4
14	0.3	2.1	1.9	0.3	3.8	3.3	0.3	11.9	10.8
15	0.3	1.8	1.5	0.6	3.5	3.1	0.4	3.0	2.5
16	0.5	2.6	2.5	0.5	2.0	1.8	0.5	5.1	4.7
17	0.4	1.3	1.2	0.6	3.9	3.5	0.5	10.3	9.5
18	0.3	4.8	4.3	0.4	2.8	2.4	0.3	14.6	13.4
19	0.3	0.5	0.4	0.4	0.8	0.8	0.3	1.9	1.7
20	0.5	2.4	2.3	0.3	0.5	0.4	0.4	2.8	2.4
21	0.4	1.3	1.3	0.6	0.6	0.6	0.4	1.0	1.0
22	0.4	1.6	1.3	0.5	7.0	6.3	0.5	21.5	20.2
23	0.5	15.2	14.4	0.4	4.0	3.3	0.4	37.7	36.7

Note: -- implies there were no violations in this sample.

three revealed preference conditions were checked in the 1960 and 1972 time periods separately and then in the combined 60 and 72 samples. Since the various demographic groups differ in size quite dramatically, the number of violations are reported in proportion to the maximum possible violations that occur in each sample.[2] We will denote this proportion as π.

These results indicate that the assumption that the individual behavior is consistent with utility maximization is indeed questionable in this expenditure data. In the 1960 and 1972 data, there are numerous violations of the necessary and sufficient conditions for utility maximization (namely SARP and GARP). In the 1960 FULL sample, percentages of maximum violations range from a low of less than 1 percent to over 13 percent. In the 1972 FULL sample, the range was much greater.

On the surface this result is surprising, given the nature of the data and the SARP and GARP conditions. Recall that within any given year, the only price variation within the data was interregional price variation. Hence a priori one would not expect there to be significant degree of price variation. For the time being, assume there isn't any price variation in the data. Given this assumption, it would be possible for a single individual to violate the WARP and SARP conditions only if there were two observations with the same amount of total expenditures but different allocations. This violation would result from the fact the data were not consistent with single-valued demand functions (i.e., strictly concave utility functions). Now given that we have observations from different individuals, one might expect a certain number of violations for this reason. However, this type of violation would not be a violation of GARP, which allows for multiple-valued demands (i.e., concavity of the utility function). Hence if there was no price variation, one would not expect any GARP violations.

However, as the tables indicate, there are a significant number of GARP violations. Apart from indicating that the expenditure data are inconsistent with the utility maximization, these violations could be consistent with the utility maximization hypothesis if the expenditure data contained "measurement error." Hence the "true" expenditure data, which are not observed, could be rationalized by a single set of preferences.

The TRUNCATED sample was created to reduce the amount of "measurement" error in the data. A quick comparison between Tables A2.2 and A2.3 leads one to conclude that errors in variables may be responsible for the numerous violations in the data. Even after controlling for the effect of sample size, the truncation of the data does significantly reduce the percentage violations.[3] However, even after throwing out the "bad" observations, there are still significant numbers of violations in the single-period samples.

Given these results, how would one now propose a test of preference change? One possible route (the one that will be explored here) is to ask whether combining the 1960 and 1972 sample for each demographic group leads to many more violations. If it does, then one might conclude that a change in preferences did indeed occur.

The revised test was performed in the following manner. First, the percentage violations (π) from the 1960 TRUNCATED sample were used to predict how the percentage of violations varied with sample size, by regressing log $\pi/(1-\pi)$ on the subsample size (N) and N^2.[4] The results of this regression are given below.

Dependent Variable: $\log(\pi/(1-\pi))$

Sample: the 23 subsamples of the 1960 TRUNCATED sample

	Parameter Estimate	t Ratio
Intercept	-12.883	-5.56
N	7.441	2.93
N^2	-1.179	2.21

The next step was to take the above estimates of how π varies with subsample size and use them to predict log $(\pi/(1-\pi))$ for the 1960, 1972 and combined subsamples in the TRUNCATED sample. The final step was to regress the errors in the predictions (actual minus the predicted values of log $(\pi/(1-\pi))$) on a dummy representing whether the observation was from the 1972 subsample (D72) or the combined sample (DCOM). The results of this regression are given below:

Dependent Variable: Prediction error of log $(\pi/(1-\pi))$

Sample: 69 subsamples from the TRUNCATED sample

	Parameter Estimate	t Ratio
Intercept	0.000	0.00
D72	-.765	-.50
DCOM	3.154	2.04

Given the construction of the dependent variable, the intercept should be and is equal to zero. These results indicate that in the

TRUNCATED sample, there exist no differences between the 1960 and 1972 samples other than those that are explained by size of the sample. The final coefficient DCOM finally provides the information we have been looking for. Based upon this sample (the TRUNCATED sample), one can conclude that in the combined sample there exist significantly more violations than would be expected. Hence there exists some evidence for the hypothesis that tastes have changed over the period of the 1960s.[5]

Notes to Appendix 2

[1] Obviously this is just one possible explanation. Another does not involve a taste change. Assume that tastes did not change and that the sole reason for violations was measurement error.

[2] For a sample with N observations, the maximum number of violations is equal to $N(N-1)/2$.

[3] In order to control for the effect that increasing the number of observations may have on the percentage of violations, the log $\pi/(1-\pi)$ was regressed upon N, N^2, and a dummy variable for observations from the TRUNCATED sample.

[4] Throughout this paper, we have assumed that the percentage of violations (π) is logistically distributed. We maintained this assumption in order to easily account for the bounded nature of the error distribution in the regression models.

[5] Throughout, we have had to deal with the stochastic nature of the data in a rather unsatisfactory manner. In the absence of panel expenditure data, one must maintain that individual preferences are the same. Hence any violations of revealed preference axioms in a single period must be attributed either to measurement error or errors in maximizing behavior. To formulate a stochastic test that would explicitly model measurement error, Varian has proposed the following. Assume that any violation results from measurement error. Hence the actual expenditure data can be thought of as the "true" data plus an error. Varian proposes to construct a set of "true" expenditure data that would obey GARP by minimizing the squared sum of differences between the observed expenditure data and the "true" data. For example, with expenditures on seven

commodities and 500 observations, this would entail constructing 3500 expenditure observations which not only minimize the square of the differences but also obey the GARP conditions. Having constructed this new data set, one would test, using a Mean Squared Error chi-squared test, whether one had to "alter" the data too much to get it to obey GARP. Obviously computer costs make it infeasible to implement such a test at this time.

References to Appendix 2

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