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THE RELATIVITY OF UTILITY:

EVIDENCE FROM PANEL DATA

by

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CONTENTS

1. Introduction

The idea that utility is a relative concept is an old one and has found its way in various social sciences under headings such as relative deprivation theory (e.g., Davis (1959), Runciman (1966)), adaptation level theory (e.g., Helson (1964, 1971)), reference group theory (Hyman and Singer (1968)) etc. Economics is largely an exception in this respect. Utility (or welfare), is usually modelled by economists as being constant and independent of the situation of others. There are some exceptions, e.g., Duesenberry (1949), Leibenstein (1950), Easterlin (1974) and Pollak (1976, 1978).

Still it is clear that dependence of individual utility functions on the behavior of others has strong implications for both positive economics and (applied) welfare economics. Recent papers by, for example, Layard (1980), Frank (1982) and Rader (1980) attest to this.

In this paper we provide new evidence on a theory which implies that utility is an entirely relative concept. As the ideas tested here have been motivated and explained at various places, (e.g., Kapteyn et al. (1980), Kapteyn and Wansbeek (1982)), we concentrate on some methodological issues. The main improvement over earlier tests is that, for the first time, panel data are available. Various strong assumptions which, in the earlier tests, were necessitated by deficiencies in the data can now be avoided. To bring out methodological issues as clearly as possible, a fair amount of space is devoted to an introduction of the utility concept being used in the test and the way it is measured. After that we specify the relativistic model which explains differences in utility functions between different individuals. Next, the test results are presented. A discussion of the results concludes the paper.

2. The Utility Concept

Consider an indirect utility function defined on prices and (after tax) income. Within a community where individuals can be assumed to face the same prices, the indirect utility function can be taken to be exclusively a function of income. Suppose, we are able to observe this indirect utility function for each individual in the community. Partly due to the lack of price variation across individuals, it will generally be impossible to retrieve the corresponding direct utility functions solely on the basis of this information. However, for tests of a relativistic theory of utility we do not need to know the complete direct utility function per individual. Implications of the theory for differences in direct utility functions between individuals carry over to implications regarding indirect utility functions. If we are thus able to measure indirect utility functions per individual, we may expect to be able to carry out at least some tests of a relativistic utility theory.

In this study we use individually measured utility functions of income, whose theoretical basis is similar, though not identical, to that of an indirect utility function. The concept used is the individual welfare function of income (WFI), introduced by Van Praag (1968, 1971). Since its theoretical foundation has been described in various papers (e.g., Van Praag (1968, 1975)), we only give a brief introduction, and concentrate on its measurement, because of its importance for an appraisal of the empirical tests.

Van Praag assumes that individuals are able to rate income levels on a bounded ratio scale. The scale has zero as its lowest point (the worst possible income, i.e. zero) and one as its highest point (the best possible income, i.e. infinity). More specifically, his theory implies that an individual n will evaluate any income y according to his WFI $U_n(y)$, which has approximately the following functional form:

$$(2.1) \qquad U_n(y) \approx \Lambda(y; \mu_n, \sigma_n) \equiv N(\ln y; \mu_n, \sigma_n),$$

where $\Lambda(\cdot; \mu_n, \sigma_n)$ is the lognormal distribution function with median $\exp(\mu_n)$ and log-variance σ_n^2, and $N(\cdot; \mu_n, \sigma_n)$ is the normal distribution function with mean μ_n and variance σ_n^2. The lognormal distribution function serves here as a purely mathematical description of $U_n(y)$. It does not entail any probabilistic connotation.

For various tests of Van Praag's hypothesis we refer to Van Herwaarden and Kapteyn (1981) and Buyze (1982). Examples of WFIs have been drawn in Figure 1. It is seen that exp (μ) is the income which is evaluated by 0.5. So, the higher an individual's μ, the higher incomes have to be to receive a favorable rating from this individual. The parameter σ determines the slope of one's WFI: a high σ-value implies a flatly sloping WFI.

In the sample used in this paper, WFIs have been measured by asking respondents in a survey the following question:

"What after tax family income would	very bad	Dfl. ....
you consider, in your circumstances,	bad	Dfl. ....
to be very bad? And bad, insufficient,	insufficient	Dfl. ....
sufficient, good and very good?"	sufficient	Dfl. ....
	good	Dfl. ....
<u>Please enter an amount on each line</u>	very good	Dfl. ....

Care has been taken that before answering this question, the respondent has gained a good understanding of the notion of after tax family income. Actually, he has been asked to compute his own after tax family income.[1)]

To illustrate how a respondent's answers to this question are used to measure his WFI, a hypothetical response has been plotted in Fig. 2. In Fig. 2, the verbal labels "very good", "good", etc. have been associated with the midpoints of the six equal intervals that partition the [0,1]-scale. In other words, the income response to "very good" is supposed to satisfy approximately $U_n(y) \approx \frac{11}{12}$, the response to "good" is supposed to satisfy approximately $U_n(y) \approx \frac{9}{12}$, etc. Given this assumption, the answers to the question provide us with a scatter of six points through which we can fit a lognormal function $\Lambda(\cdot;\mu_n,\sigma_n)$ by means of OLS. In this way, the parameters μ_n and σ_n of respondent n are estimated.

Obviously, an important assumption in this procedure is that the verbal labels correspond to equal intervals of the [0,1]-scale. This so-called equal interval hypothesis has been tested by Buyze (1982) and Antonides et al. (1980). The general conclusion is that the hypothesis is not exactly true but that it provides a reasonable approximation.

It will be argued in Section 5 that a possible non-validity of the equal interval assumption will bias our empirical test of the relative utility hypothesis towards rejection. A discussion of this point is postponed until Section 5. For the moment, the equal interval hypothesis is taken for granted.

1) From now on, we will refer to after tax family income simply as "income".

Figure 1. The individual welfare function of income for some values of μ.

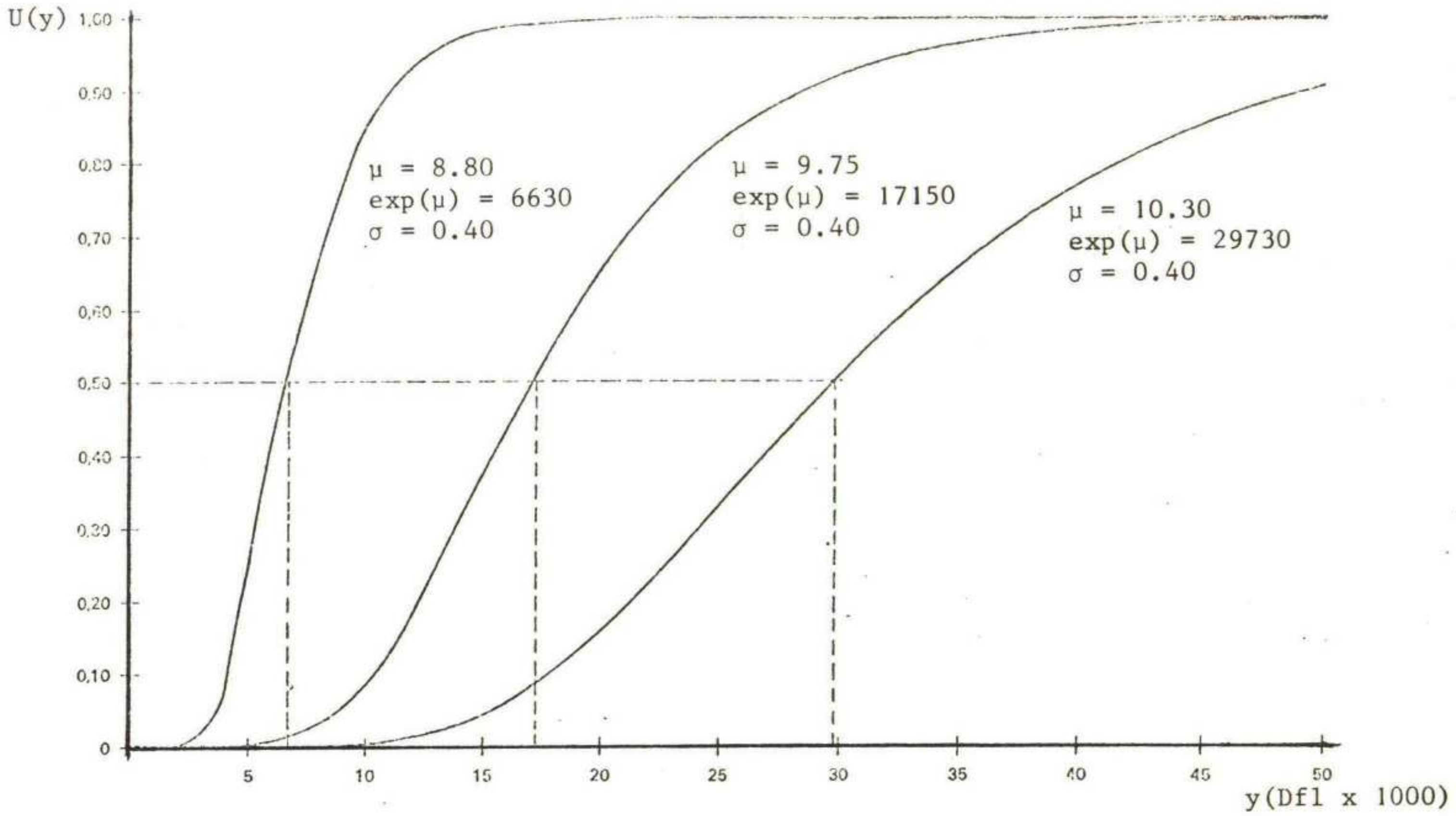
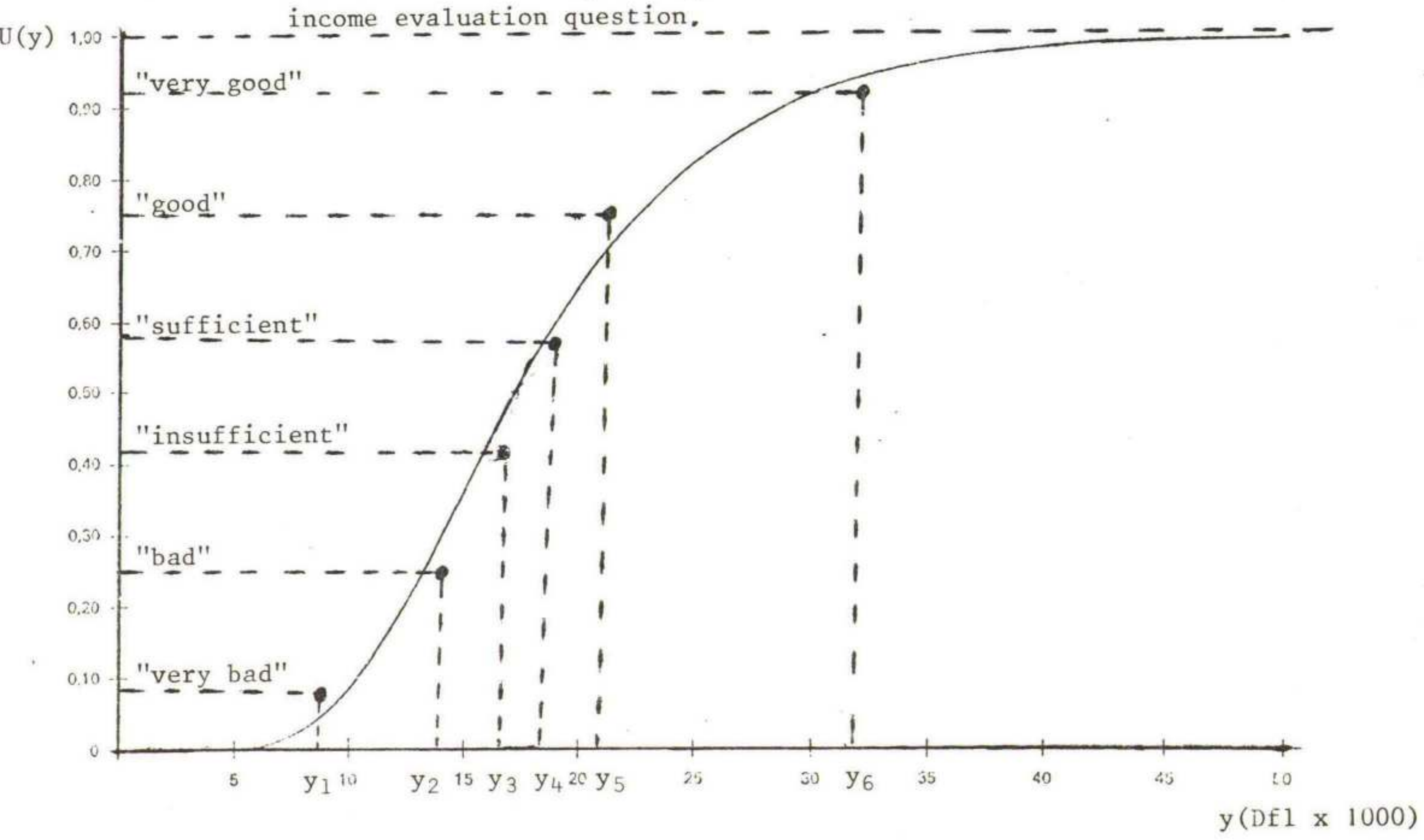


Figure 2. The estimation of μ_n and σ_n. y_1, y_2, .., y_6 are the answers on the income evaluation question.



3. Relative Utility

In line with the various theories mentioned in the introduction, Kapteyn (1977) has formulated a theory which assumes that utility is completely relative. For expositions of his so-called theory of preference formation we refer to Kapteyn (1977, 1980) or Kapteyn et al. (1980). Here we shall only present a simplified version which can be tested against the data at hand.

The basic idea is that an individual's WFI is nothing else than a perceived income distribution. That is, an individual evaluates any income level by its ranking in the income distribution which he perceives. To operationalize this idea, we have to explain what is meant by a perceived income distribution. To that end some notation is introduced:

Let there be N individuals in society. Time is measured in years, $t = -\infty, \ldots, 0$, where $t = 0$ represents the present. At each moment of time an individual n ($n = 1, \ldots, N$) is assumed to assign non-negative reference weights $w_{nk}(t)$ to any individual k in society ($k = 1, \ldots, N$), $\Sigma_{k=1}^{N} w_{nk}(t) = 1$. The reference weights indicate the importance individual n attaches to the income of individual k at time t. Obviously, quite a few of the $w_{nk}(t)$ will be zero. On the other hand, $w_{nn}(t)$, i.e., the weight that individual n attaches to his own income at time t, may be substantial. The set $(w_{n1}(t), \ldots, w_{n,n-1}(t), w_{n,n+1}(t), \ldots, w_{nN}(t))$ will sometimes be referred to as n's social reference group at time t.

Furthermore, let $y_k(t)$ be the income of individual k at time t. The reference weights now allow for the definition of a perceived income distribution at time t. Denote this function by $F_n(y|t)$, then its definition is

$$(3.1) \qquad F_n(y|t) \equiv \sum_{\{k; y_k(t) \leq y\}} w_{nk}(t) .$$

The $F_n(y|t)$ for any t can be aggregated to one presently perceived income distribution, $F_n(y)$. To that end a non-negative memory function $a_n(t)$ is introduced, which describes individual n's weighting of perceived consumption over time,

$$(3.2) \qquad \sum_{t=-\infty}^{0} a_n(t) = 1 \ , \quad n = 1,\ldots,N.$$

The presently perceived distribution function $F_n(y)$ can now be defined as

$$(3.3) \qquad F_n(y) \equiv \sum_{t=-\infty}^{0} a_n(t)\ F_n(y|t) \ .$$

As indicated above, the preference formation theory claims that this perceived income distribution equals the utility function $U_n(y)$ of the individual. It is this claim that we want to test in this paper.

The development of the argument so far has been in terms of individual incomes, whereas our data refer to family income (cf. the wording of the survey question above). It may be expected that a family with children needs more income than a single person to reach the same utility level, so it stands to reason to reformulate the preference formation theory in terms of incomes per equivalent adult. Let $f_k(t)$ be the number of equivalent adults in family k at time t. The income per equivalent adult in this family at time t is denoted by

$$(3.4) \qquad \tilde{y}_k(t) \equiv y_k(t)/f_k(t) \ .$$

The reformulation of $U_n(y)$ in terms of incomes per equivalent adult amounts to a transformation of the income scale: y is replaced by $\tilde{y} \equiv \frac{y}{f_n}$ and

e^{μ_n} by $\frac{e^{\mu_n}}{f_n}$.[1)] Consequently,

$$(3.5) \qquad U_n(y) = N(\ln y; \mu_n, \sigma_n) = N(\ln\left(\frac{y}{f_n}\right); \mu_n - \ln f_n, \sigma_n)$$

$$= N(\ln \tilde{y}; \tilde{\mu}_n, \tilde{\sigma}_n) = \tilde{U}_n(\tilde{y}) .$$

Replacing $y_k(t)$ and y in (3.1) and (3.3) by $\tilde{y}_k(t)$ and $\tilde{y}$, we obtain the perceived distribution of incomes per equivalent adult $\tilde{F}_n(\tilde{y})$.

The theory of preference formation now states

$$(3.6) \qquad \tilde{U}_n(\tilde{y}) = \tilde{F}_n(\tilde{y}); \quad n = 1,\ldots,N; \quad \tilde{y} \in [0,\infty).$$

Equation (3.6) implies that utility is a completely relative concept. The utility of a certain income per equivalent adult is obtained by comparing it with the perceived distribution of incomes per equivalent adult.

To test the theory we derive from (3.6) implications for variations in μ and σ over individuals, which can be confronted with the data at hand.

Denote the first log-moment of $\tilde{F}_n(\tilde{y})$ by $\tilde{m}_n$.

$$(3.7) \qquad \tilde{m}_n = \int_0^\infty \ln \tilde{y} \, d\tilde{F}_n(\tilde{y}) = \sum_{t=-\infty}^{0} a_n(t) \sum_{k=1}^{N} w_{nk}(t) \ln \tilde{y}_k(t) .$$

The equality of the two distribution functions implies the equality of the first two log-moments:

$$(3.8) \qquad \mu_n = \ln f_n + \tilde{m}_n + \varepsilon_n$$

$$= \ln f_n + \sum_{t=-\infty}^{0} a_n(t) \sum_{k=1}^{N} w_{nk}(t) \ln \tilde{y}_k(t) + \varepsilon_n$$

1) For convenience, we generally omit arguments equal to zero, so $f_n \equiv f_n(0)$.

and

$$(3.9) \qquad \sigma_n^2 = \sum_{t=-\infty}^{0} a_n(t) \sum_{k=1}^{N} w_{nk}(t) \, [\ln \tilde{y}_k(t) - \tilde{m}_n]^2 + \delta_n ,$$

where measurement errors in μ_n and σ_n^2 and errors in the equations are taken into account by means of the i.i.d. distributed disturbance terms ε_n and δ_n, with zero means and variances σ_ε^2 and σ_δ^2.

In principle, (3.8) and (3.9) are in a form suitable for estimation from panel data, the results of which should give us insight into the validity of the preference formation theory. However, without further restrictions there are far too many parameters to be estimated (particularly the N(N-1) independent reference weights). In order to facilitate estimation, a few more assumptions and definitions are needed. We assume that $w_{nn}(t)$ is the same for all individuals and constant over time, i.e., all individuals give themselves the same constant weight. We write $\beta_2 \equiv w_{nn}(t)$ and $\beta_3 \equiv \sum_{k \neq n} w_{nk}(t) = 1-\beta_2$. The function $\ln f_k(t)$ is specified as $\beta_0 + \beta_1 \ln fs_k(t)$ where $fs_k(t)$ is the number of members of family k at time t. The memory function $a_n(t)$ is assumed to be the same for everyone and is specified as $a_n(t) = (1-a)a^{-t}$. Furthermore, we define

$$(3.10) \qquad q_{nk}(t) \equiv w_{nk}(t)/\beta_3 , \quad k \neq n$$

$$\equiv 0 \qquad , \quad k = n$$

$$(3.11) \qquad \bar{m}_n(t) \equiv \sum_k q_{nk}(t) \ln y_k(t) ,$$

$$(3.12) \qquad \bar{h}_n(t) \equiv \sum_k q_{nk}(t) \ln f_k(t) = \beta_0 + \beta_1 \{ \sum_k q_{nk}(t) \ln fs_k(t) \} \equiv$$

$$\equiv \beta_0 + \beta_1 \overline{hs}_n(t) \ ,$$

where $\overline{hs}_n(t)$ is defined implicitly. So, $\overline{m}_n(t)$ and $\overline{hs}_n(t)$ are the log-means of incomes and family sizes in family n's social reference group at time t.

All this makes it possible to rewrite (3.8) as

$$(3.13) \quad \mu_n = \ln f_n + (1-a) \sum_{t=-\infty}^{0} a^{-t} [\beta_2 \{ \ln y_n(t) - \ln f_n(t) \} + \beta_3 \{ \overline{m}_n(t) - \overline{h}_n(t) \}] + \varepsilon_n$$

This can be written in lagged form as

$$(3.14) \quad \mu_n - a\mu_n(-1) = \ln f_n - a \ln f_n(-1) + \beta_2(1-a)(\ln y_n - \ln f_n) + \beta_3(1-a)(\overline{m}_n - \overline{h}_n)$$

$$+ \varepsilon_n - a\varepsilon_n(-1) \ ,$$

or

$$(3.14) \ \mu_n = [1 - \beta_2(1-a)] \ln f_n - a \ln f_n(-1) + \beta_2(1-a)\ln y_n + \beta_3(1-a)\overline{m}_n - \beta_3(1-a)\overline{h}_n$$

$$+ a\,\mu_n(-t) + \varepsilon_n - a\,\varepsilon_n(-t) \ .$$

Using the expression for $\ln f_n$ we obtain

$$(3.16) \quad \mu_n = [1 - \beta_2(1-a)]\beta_1 \ln fs_n - a\beta_1 \ln fs_n(-1) + \beta_2(1-a)\ln y_n + \beta_3(1-a)\overline{m}_n$$

$$- \beta_3(1-a)\beta_1 \overline{hs}_n + a\mu_n(-1) + \varepsilon_n - a\varepsilon_n(-1) \ .$$

We observe that (3.16) has no constant term. If we allow for the fact that incomes in previous years have to be deflated by a price index it is easy to show that this does not influence the coefficients in (3.16), but only gives rise to a constant term. In the empirical application (3.16) has been estimated with a constant term included.

It is rather straightforward to use (3.9) and derive an expression for σ_n^2 similar to (3.16). However, that expression is non-linear in both parameters and variables. It will be seen in the next section that estimation of (3.16), which is non-linear in parameters but linear in variables, is already complicated. Estimation of a similar relation for σ_n^2 would involve problems of measurement errors in a non-linear model. Since we have not yet been able to solve the estimation problems posed by such a model satisfactorily, only (3.16) will be confronted with the data.

4. Estimation of the μ_n -equation

The data consists of the first two waves of a panel of 775 households in The Netherlands. The main breadwinner of each household was interviewed in March 1980 and the same person was reinterviewed in March 1981. The items in the questionnaire included questions to measure the respondent's WFI, the after tax family income, family composition, and a number of demographic and socio-economic characteristics. On the basis of this information (3.16) is estimated.

The main problem with the estimation of (3.16) is that $\bar{m}_n$ and $\bar{hs}_n$ are unobservable. To solve this problem we have to make explicit assumptions about the process which generates the references weights. Our first, rather innocuous, assumption is

$$(4.1) \qquad q_{nk} = p_{nk} + \delta_{nk}.$$

Here, p_{nk} is a parameter which may interpreted as the probability that individual[1] n meets individual k (we call that a contact between n and k); q_{nk} is the relative frequency of contacts between n and k in any given year, out of all the contacts by individual n; δ_{nk} is an error term. Both p_{nk} and q_{nk} are subjective, i.e. individual n may weigh certain contacts more than others and in particular there does not have to hold that $p_{nk} = p_{kn}$ or $q_{nk} = q_{kn}$. Given our interpretation of q_{nk} and p_{nk}, an obvious assumption is that q_{nk} follows a multinomial distribution. However, that assumption will not play a role in

1) In view of the data and given the development in the preceding section, the word "individual" is an abbreviation of "the main breadwinner of the household".

what follows. We do assume, though, that $E\,\delta_{nk} = 0$ and that δ_{nk} is independent of p_{nk} and of all incomes and family sizes in society.

Given (4.1) we can write

$$(4.2)\quad \bar{m}_n = \sum_k q_{nk}\ln y_k = \sum_k (p_{nk}+\delta_{nk})\ln y_k = \sum_k p_{nk}\ln y_k + \sum_k p_{nk}\delta_{nk} \equiv m_n^* + \upsilon_n,$$

with $m_n^* = \sum_k p_{nk}\ln y_k$ and $\upsilon_n = \sum_k \delta_{nk}\ln y_k$. The foregoing assumptions regarding the δ_{nk} imply the independence of m_n^* and υ_n. So far we have only replaced one unobservable, $\bar{m}_n$, by another, m_n^*. Let us now assume that society is partitioned in groups $G_1,\ldots,G_i,\ldots,G_I$ such that there exist constants P_i satisfying

$$(4.3)\qquad p_{nk} = \begin{cases} \dfrac{P_i}{N_i-1} & \text{if } n \in G_i,\ k \in G_i, \\[2ex] \dfrac{1-P_i}{N-N_i} & \text{if } n \in G_i,\ k \notin G_i, \end{cases}$$

where N_i is the number of individuals in group i. Notice from (4.1) and (4.3) that

$$(4.4)\qquad E \sum_{k\in G_i} q_{nk} = \sum_{k\in G_i} p_{nk} = P_i \quad , \; i = 1,\ldots,I.$$

Thus, P_i is the total reference weight that any individual n in G_i assigns, on average, to the other individuals $n\in G_i$. Assumption (4.3) therefore states that, on average, individuals within a group G_i give a total weight P_i to others in the same group and a total weight $(1-P_i)$ to individuals outside their own group.

As a consequence of this assumption, we can write

$$(4.5)\qquad m_n^* = \sum_k p_{nk} \ln y_k = \sum_{k \in G_i} p_{nk} \ln y_k + \sum_{k \notin G_i} p_{nk} \ln y_k =$$

$$= \frac{P_i}{N_i - 1} \sum_{\substack{k \in G_i \\ k \neq n}} \ln y_k + \frac{1-P_i}{N-N_i} \sum_{k \notin G_i} \ln y_k \equiv P_i y_n^* + (1-P_i)\tilde{y}_n^* \text{ , for } n \in G_i,$$

where y_n^* is the mean log-income of individuals in group G_i, other than n; $\tilde{y}_n^*$ is the mean log-income of individuals outside G_i. Let $\bar{Y}$ be mean log-income in society, so that $N\bar{Y}$ is total log-income in society. Then there holds

$$(4.6)\qquad (N_i - 1)y_n^* + (N-N_i)\tilde{y}_n^* + \ln y_n = N.\bar{Y} \qquad \text{for } n \in G_i.$$

Solving (4.6) for $\tilde{y}_n^*$ and inserting the result in (4.5) yields

$$(4.7)\qquad m_n^* = P_i . y_n^* + \frac{1-P_i}{N-N_i}(N-N_i)\tilde{y}_n^*$$

$$= P_i y_n^* + \frac{1-P_i}{N-N_i} N.\bar{Y} - \frac{1-P_i}{N-N_i} \ln y_n - \frac{1-P_i}{N-N_i}(N_i - 1)y_n^*$$

$$= \left\{ 1-(N-1).\frac{1-P_i}{N-N_i} \right\} y_n^* + \frac{1-P_i}{N-N_i} N\bar{Y} - \frac{1-P_i}{N-N_i} \ln y_n.$$

So far the P_i were taken as constants. Now we take them as random variables (so that the preceding analysis was conditional upon the realization of the random variables). We assume that the P_i are generated according to a process satisfying:

$$(4.8)\qquad \frac{1-P_i}{N-N_i} = \tilde{q} + \Delta_i, \qquad i = 1,\dots,I$$

where Δ_i is an i.i.d. random variable with mean zero and variance σ_Δ^2. The interpretation of $\tilde{q}$ is that it is the mean reference weight assigned by individuals to others outside their own group. Combining (4.7) and (4.8) we obtain

$$(4.9)\qquad m_n^* = [1-(N-1)\tilde{q}]y_n^* + \tilde{q}(N\bar{Y}-\ln y_n) + \Delta_i(N\bar{Y}-\ln y_n-(N-1)y_n^*)$$

Inserting (4.9) in (4.2) yields

$$(4.10)\qquad \bar{m}_n = [1-(N-1)\tilde{q}]y_n^* + \tilde{q}(N\bar{Y}-\ln y_n) + u_n\ ,$$

where $u_n = \upsilon_n + \Delta_i(N\bar{Y}-\ln y_n-(N-1)y_n^*)$.

According to (4.8) $\tilde{q}$ is of the order of magnitude of $1/(N-N_i)$, so that $\tilde{q}\ \ln y_n$ can be neglected without losing much precision, provided that groups are defined in such a way that $N - N_i$ is large.[1] Analogous to (4.10) we can derive a similar expression for $\overline{hs}_n$ (now neglecting $\tilde{q}\ \ln fs_n$):

$$(4.11)\qquad \overline{hs}_n = [1-(N-1)\tilde{q}]\ f_n^* + \tilde{q}\ N\bar{F} + v_n\ ,$$

where f_n^* is the mean log-family size of families in the group individual n belongs to, excluding his own family, and $\bar{F}$ is mean log-family size in society. Inserting (4.10) and (4.11) in (3.16) we obtain as an estimating equation:

$$(4.12)\qquad \mu_n = [1-\beta_2(1-a)]\beta_1\ \ln fs_n - a\beta_1\ \ln fs_n(-1) + \beta_2(1-a)\ln y_n$$

$$+ \beta_3(1-a)(1-\kappa)y_n^* - \beta_3(1-a)(1-\kappa)\beta_1 f_n^* + a\,\mu_n(-1) + \beta_0 + \zeta_n\ ,$$

1) Given that N is the number of families in society, $N - N_i$ will be large as long as the different groups are of comparable size.

where $\kappa \equiv (N-1)\tilde{q}$; $\beta_0 \equiv \tilde{q}N(\bar{Y}-\bar{F})$;

$$(4.13) \quad \zeta_n = \varepsilon_n - a\,\varepsilon_n(-1) + \beta_3(1-a)u_n - \beta_3(1-a)\beta_1 v_n$$

As a result of the various assumptions made above regarding the stochastic distribution of reference weights q_{nk}, (4.12) explains individual n's μ on the basis of his present income and family size, last year's μ, and the mean (log-) income and mean (log-) family size in group G_i, of which n is a member.

To estimate (4.12) a number of issues have to be dealt with. First of all we have to specify the groups in society which form the basis of the definition of y_n^* and f_n^*. We have partitioned the sample in groups of respondents with identical characteristics (i.e. the same education level, age bracket and employment status[1]). For these groups we have calculated the sample counterparts of y_n^* and f_n^* for each individual (i.e. within a group the mean log-income and log-family size varies slightly per respondent, because the respondent's own income and family size are not part of the definition of y_n^* and f_n^*).

Evidently, replacing y_n^* and f_n^* by their sample counterparts induces measurement error. If there are M_i observations for group i, then an obvious estimate of y_n^* is

$$(4.14) \quad \hat{y}_n \equiv \frac{1}{M_i-1} \sum_{\substack{k \in G_i \\ k \neq n}} \ln y_k . \qquad n \in G_i$$

1) Five education levels are distinguished, three employment situations, (self-employed, employee, not employed) and five age brackets (less than 30, 30-39, 40-49, 50-65, over 65). This leads to 51 groups in the sample with respondents who have identical characteristics.

This quantity measures y_n^* unbiasedly. For the variance of $\hat{y}_n$ we employ the usual estimator.

The construction of a proxy for f_n^* is analogous. So we have proxies for y_n^* and f_n^* with known measurement error covariance matrix. In principle this covariance matrix differs per group. For simplicity we have averaged all these matrices and used the result as our estimate of the error variance-covariance matrix for all observations.

From (4.13) it is clear that $\mu_n(-1)$ correlates with the error term. The covariance between $\mu_n(-1)$ and the error in the equation is one of the parameters to be estimated.

Assuming that the random variables involved all follow approximately a normal distribution, (4.12) can be estimated by means of maximum likelihood. To that end, the LISREL computer program (version IV) has been used. The LISREL-specification is given in the Appendix. We have ignored one major complication, namely that the measurement errors in the proxies for y_n^* and f_n^* are correlated across observations for those observations that pertain to the same social group. This neglect does not impair the consistency of the ML-estimates obtained by LISREL, but the asymptotic standard errors have to be viewed with some reservation.

5. Results and discussion

Two versions of the model have been estimated; in one version we assume the measurement error in y_n^* and f_n^* to be absent; in the other version we allow for errors in y_n^* and f_n^* with a known covariance matrix, as described above. The results are given in Table 1.

The differences between both columns are generally small. The parameter estimate most affected by the assumption on the errors in y_n^* and f_n^* is that of $\kappa \equiv (N-1)\tilde{q}$. Recalling that $\tilde{q}$ is the mean reference weight which individuals assign to others outside their own group, it is clear that $\kappa(N-N_i)/(N-1)$ is the total reference weight assigned on average to people outside one's group. Since N_i is small relative to N, $(N-N_i)/(N-1) \simeq 1$. Thus κ measures approximately the total reference weight assigned to others outside one's group and $1-\kappa$ the total weight assigned to others within the group. So $\kappa = 0.500$ implies that individuals within and outside one's group get about equal total weight. The estimate in the second column of Table 1 ($\kappa = 0.421$) suggests a somewhat higher total weight for individuals within the group. Both estimates turn out to be rather unreliable, so no significance can be attached to this difference.

Notice that since there are a lot more people outside each group than in it, the average weight assigned to individuals within one's group is substantially higher than the average weight assigned to individuals outside the group. This can be seen more clearly by using (4.8) to obtain

$$(5.1) \qquad \kappa = (N-1)\tilde{q} = E(N-1).\frac{1-P_i}{N-N_i} = E(1-P_i)/[(N-N_i)/(N-1)]$$

Table 1. Estimation results for equation (4.12)

Parameter	Without measurement errors in y_n^*, f_n^*	With measurement errors in y_n^*, f_n^* (cov. matrix known)
a	0.828	0.834
	(0.145)	(0.147)
β_1	0.114	0.114
	(0.039)	(0.039)
β_2	0.663	0.657
	(0.128)	(0.140)
β_3	0.337	0.343
	(0.128)	(0.140)
κ	0.500	0.421
	(0.462)	(0.502)
variance of ζ_n	0.029	0.029
	(0.005)	(0.007)
covariance $(\mu_n(-1),\zeta_n)$	-0.015	-0.015
	(0.007)	(0.007)
variance of error[b] in y_n^*	-	0.0066
variance of error[b] in f_n^*	-	0.0095
covariance of error[b] in y_n^* and f_n^*	-	0.0016
$1 - \text{var}(\zeta_n)/\text{var}(\mu_n)$	0.770	0.770
Degrees of freedom	1	1
χ^2	0.050	0.072

a Asymptotic standard errors in parentheses

b Specified a priori. See the end of Section 4.

Here $(N-N_i)/(N-1)$ is the total weight given to individuals outside one's own group G_i if everyone in society would get the same reference weight $1/(N-1)$; $1-P_i$ is the total weight actually given to individuals outside one's own group G_i. If the division in groups were irrelevant we would have $\kappa = 1$. The smaller κ is, the higher the average weight assigned to people within one's group relative to the weight assigned to people outside one's group. Of course, the rather high standard errors of the estimates of κ indicate that κ could very well be equal to one, so that our choice of characteristics to define groups may have been a poor one.

The estimates of β_2 and β_3 suggest that the total weight which an individual assigns to the incomes of all other people is about half the weight which he gives to his own income (in present and past). This contrasts with earlier results obtained by Kapteyn et al. (1980) who found β_3 to be approximately twice as large as β_2. There are two ways to explain the difference. First of all Kapteyn et al. use cross-section data and their analysis rests upon a number of strong assumptions, required to identify the model. Secondly, their analysis pertains to holiday expenditures rather than income. One would expect β_3 to differ between goods. The more conspicuous a good, the higher β_3 probably is. Since holidays are among the most conspicuous consumption items, the corresponding β_3 should be substantially higher than for income, which is an aggregate of all consumption possibilities, both conspicuous and unconspicuous ones.

The parameter β_1 measures the increase in a family's cost of living due to an increase in family size. If the size of the family increases by 1% then the cost of living of the family increases by β_1%. The low values of β_1 suggest substantial economies of scale in the operation of a family. In itself it is of interest to see how a purely subjective model provides estimates of seemingly "objective" quantities like cost of living differences. It has been argued elsewhere (e.g., Kapteyn and Van Praag (1980)) that the methodological basis of the present measurement method is identical to the one underlying conventional demand systems approaches to the measurement of differences in cost of living. Although the specification of $\ln f_n$ by $\beta_0 + \beta_1 \ln fs_n$ is very primitive, it is noteworthy that never before in cost of living studies account was taken of both preference interdependence and habit formation.

As to the exact numerical value of β_1 a caveat should be expressed. No measurement error in fs_n has been allowed. The definition of fs_n in the questionnaire (all persons living with the family plus relatives who are living elsewhere but receive at least 50% financial support from the family) is, moreover, somewhat ambiguous. Although there is one degree of freedom left, which would allow for the introduction of a measurement error in fs_n, attempts to do so yielded nonsensical results. This suggests that relaxing the assumption of no measurement error in fs_n brings the model on the brink of underidentification. Only additional waves of data will make it possible to investigate the effect of measurement error in fs_n in a statistically reliable way.

The estimate of a (approximately 0.83) suggests a fairly strong influence of past income distributions. For instance, weights given to years 0, -1, -2, etc. are: 0.17, 0.14, 0.12, 0.10, 0.08, 0.07, 0.06, 0.05, 0.04, 0.03, etc. So the present year receives a weight which is about six times as high as the weight given to an income ten years ago, but all past years combined get a total weight equal to 0.83 as compared to 0.17 for the present year. According to these results, a discussion of the relativity of utility framed exclusively in cross-sectional terms would be highly incomplete.

Due to the modest sample size, the availability of only two waves in the panel and the omission from the model of a relation similar to (3.16) for the explanation of σ^2, the numerical values of the estimates have to be viewed with some care. Also, the interpretation of the parameters and a discussion of implications have been given elsewhere (e.g., Kapteyn et al. (1980)). Still, the statistical performance of the model is quite promising. The value of $1 - \mathrm{var}(\zeta_n)/\mathrm{var}(\mu_n)$ indicates that the variance of the error term ζ_n contributes 23% to the total variance of μ_n. It is not possible to identify the separate contributions of the components of ζ_n given in (4.13). However, it seems reasonable to assume that a major part comes from u_n and v_n, which represents the imperfections in our definition of the reference groups.

For the sake of judging the quality of the theoretical model (3.8) it is important to obtain information on the variance of ε_n relative to the variance of μ_n. Under our stochastic assumptions it follows from (4.13) that

$$(5.2) \qquad \mathrm{var}(\zeta_n) = (1+a^2)\sigma_\varepsilon^2 + \mathrm{var}[\beta_3(1-a)u_n - \beta_3(1-a)\beta_1 v_n] > (1+a^2)\sigma_\varepsilon^2 ,$$

so that

$$(5.3) \qquad \sigma_\varepsilon^2 < \frac{\mathrm{var}(\zeta_n)}{1+a^2} = 0.017 ,$$

where the estimates from Table 1 have been used to obtain the last equality (both columns give the same result up to three decimal places). Consequently,

$$(5.4) \qquad 1 - \sigma_\varepsilon^2/\mathrm{var}(\mu_n) > 0.86$$

Thus, (3.8) appears to explain at least 86% of the variance in μ_n. Part of the unexplained variance has to be ascribed to measurement error in μ. Since, moreover, we have used a very crude proxy for the effects of family composition, the overall results indicate the need for better measurements and definitions of the variables involved, but the model itself appears to be basically correct.

To conclude, we recall the description of the measurement of μ and σ per individual. One remembers the crucial role played by the equal interval assumption. To the extent that this assumption introduces a systematic bias, this would probably not affect an explanation of the <u>variation</u> in μ. To the extent that the equal interval assumption introduces random measurement error in μ, this would attenuate the explanatory power of a relation like (3.16). So, if anything, the equal interval assumption may have lead to an underestimation of the explanatory power of the preference formation theory.

6. Conclusion

This paper has been devoted to an empirical analysis of a preference formation theory implying that utility is relative. The results are unambiguously supportive of the theory. The theory generalizes related earlier results mentioned in the introduction pertaining to static (cross-sectional) models. This is by no means the only possible test and we hope to use different data and different operationalizations of utility to carry out further tests. Highest on our list of priorities is the estimation of an equation for σ_n^2, similar to (3.16), although it is clear from Section 4 that such enterprise involves major difficulties.

To the extent that the utility concept used in this paper (the WFI) is a sufficiently close approximation to the indirect utility function defined in economic theory, it seems clear that utility functions are interdependent and subject to habit formation. This has far reaching consequences for both positive and normative economics. It may be held, of course, that direct questions about satisfaction measure something entirely different from the economic utility concept. Although, on intuitive grounds, we find this hard to accept, further research into the relation between verbal statements about satisfaction and economic behavior is evidently needed.

Appendix: Estimation Procedure and Data

The model (3.16), has been estimated by means of the computer program LISREL IV. This program gives the full information maximum likelihood estimates (under normality) of the following structural model

$$\text{(A.1)} \qquad B\,\eta = \Gamma\xi + \zeta \ ,$$

where η is an m-vector of (possibly unobservable) dependent variables, ξ is an n-vector of (possibly unobservable) independent variables, ζ is an m-vector of errors and B and Γ are $(m \times m)$ and $(m \times n)$ matrices of coefficients. It is assumed that ζ is uncorrelated with ξ and that B is non-singular.

Equation (3.16) is non-linear in parameters. In order to deal with these non-linearities within the LISREL format some auxiliarly variables had to be introduced. The complete specification reads as follows:

$$\text{(A.2)} \quad \begin{bmatrix} 1 & -a & 0 & -a & \gamma_2 & \gamma_3 \\ & 1 & & & & \\ & & 1 & & & \\ & & 1 & 1 & & \\ & & & & 1 & \\ & & & & & 1 \end{bmatrix} \begin{bmatrix} \mu_n \\ \eta_2 \\ \eta_3 \\ \eta_4 \\ \eta_5 \\ \eta_6 \end{bmatrix} = \begin{bmatrix} 0 & 0 & \beta_1 & \gamma_2 & \gamma_3 & 0 & 1 \\ 1 & & & & & & \\ & \beta_1 & & & & & \\ & & & & & & \\ & & \beta_1 & & & & \\ & & & & & \beta_1 & \end{bmatrix} \begin{bmatrix} \mu_n(-1) \\ \ln fs_n(-1) \\ \ln fs_n \\ \ln y_n \\ y_n^* \\ fs_n^* \\ \zeta_n \end{bmatrix}$$

$$\gamma_2 = \beta_2(1-a)$$
$$\gamma_3 = \beta_3(1-\kappa)(1-a)$$

The first element of η and the first four elements of ξ are observable and the other elements of η and ξ are unobservable. All variables are taken as deviations from their means. The lower triangle of the variance-covariance matrix of the independent variables Φ is given by

$$\Phi = \begin{bmatrix} \Phi_{11} & & & & & & \\ \Phi_{21} & \Phi_{22} & & & & & \\ \Phi_{31} & \Phi_{32} & \Phi_{33} & & & & \\ \Phi_{41} & \Phi_{42} & \Phi_{43} & \Phi_{44} & & & \\ \Phi_{51} & \Phi_{52} & \Phi_{53} & \Phi_{54} & \Phi_{55} & & \\ \Phi_{61} & \Phi_{62} & \Phi_{63} & \Phi_{64} & \Phi_{65} & \Phi_{66} & \\ \Phi_{71} & & & & & & \Phi_{77} \end{bmatrix}$$

where Φ_{ij}, $i,j \leq 4$, are the known variances and covariances of the observable independent variables and Φ_{5j} and Φ_{6j}, $j \leq 4$, are covariances between observables and (an estimate of) y_n^* resp. fs_n^*, and thus also known.
Φ_{55}, Φ_{65} and Φ_{66} can be adjusted by subtracting the error (co)variances.

The only empirical information LISREL requires, if the structural model is identified, is the variance-covariance matrix of the observable variables. This matrix may be computed from the correlations and standard deviations, given in Table A.1.

Table A.1. Sample means, standard deviations and correlations of the observable variables.

Variable	Mean	St. dev.	Correlation with						
			μ_n	$\mu_n(-1)$	ln fsn(-1)	ln fsn	$\ln y_n$	$\hat{y}_n$	$\hat{fs}_n$
μ_n	10.11	0.35	1						
$\mu_n(-1)$	10.07	0.37	.862	1					
$\ln fs_n(-1)$	1.01	0.52	.474	.349	1				
$\ln fs_n$	1.00	0.52	.477	.353	.938	1			
$\ln y_n$	10.31	0.42	.826	.643	.401	.417	1		
$\hat{y}_n$	10.30	0.29	.593	.597	.349	.353	.643	1	
$\hat{fs}_n$	1.00	0.31	.392	.561	.538	.538	.391	.597	1

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