Review of Income and Wealth
Series 49, Number 3, September 2003

USING PANEL DATA ON INCOME SATISFACTION TO ESTIMATE EQUIVALENCE SCALE ELASTICITY

BY JOHANNES SCHWARZE*

University of Bamberg, DIW (German Institute for Economic Research), Berlin, and IZA Bonn, Germany

In this paper, the equivalence scale elasticity will be estimated by using individual panel data on income satisfaction from the German Socio-Economic Panel Study (GSOEP). Satisfaction or happiness data have been more frequently used by economists in recent years to analyze individual well-being. The approach differs from other subjective approaches as respondents are requested to evaluate current income rather than income in hypothetical situations. The estimated scale elasticity is higher compared to those from other subjective approaches based on German data. In addition, panel data enable different scale use by the respondents to be controlled. It can be shown that elasticity decreases when unobserved fixed-effects are controlled for.

1. EQUIVALENCE SCALE ELASTICITY: AN UNSOLVED PUZZLE

Theoretical and empirical work has shown that measures of income inequality and income poverty depend heavily on the equivalence scale chosen (e.g. Buhmann *et al.*, 1988; Coulter *et al.*, 1992b; Burkhauser *et al.*, 1996). In general, equivalence scales are intended to measure the variation in income needed to bring households of different compositions to the same welfare level. The main arguments revolve around economies of scale in household formation and increasing utility when households choose to have children.[1] Buhmann *et al.* (1988) have shown that nearly all equivalence scales can be approximated by h^e, where h is household size and $e[0,1]$ is the scale elasticity parameter. Equivalent household income Y^e then can be expressed as $Y^e = Y/h^e$, where Y is total household income. If e equals 1, equivalent income equals per capita income, whereas e equal to 0 implies no adjustment for needs. The larger e is, the higher will be the scale rate relative to that for a single-person household.

In applied inequality analysis, researchers often make use of so-called expert scales, where different weights are assigned to different household members. Most of these scales depend not only on household size—as in the Buhmann *et al.* (1988) formulation—but also on other household characteristics, such as age. The scale proposed by the OECD, for example, assigns a weight of 1 to the first adult, 0.5 to each additional adult, and 0.3 for each child under 15 years of age. Computing the elasticity e from expert scales (see Buhmann *et al.*, 1988 and Table 4 below) shows values between 0.53 and 0.66 for the OECD scale and between

Note: For helpful comments I am grateful to two anonymous referees.

*Correspondence to: Johannes Schwarze, Professur für Sozialpolitik, Universität Bamberg, Feldkirchenstrasse 21, 96045 Bamberg, Germany (johannes.schwarze@sowi.uni-bamberg.de).

[1]This argument is true, however, only when people have children by choice and not when they are unplanned. Bojer and Nelson (1999) give a summary of the discussion.

0.82 and 0.87 for another expert scale derived from the German social assistance program.

It is widely accepted that there is no uniquely true equivalence scale as equivalence scales are a part of social evaluation (e.g. Coulter *et al.*, 1992a). Nevertheless, they should at least be based on data derived from observing individuals and households. Indeed, many researchers have tried to estimate equivalence scales from individual data based on economic theory.

Two strands of estimating equivalence scales can be distinguished. Based on consumer theory, the scale can be obtained from consumption or expenditure data by estimating a system of demand equations (for an overview, see, for example, Nelson, 1992, 1993; for an estimation for Germany, see Merz *et al.*, 1993). It has been shown, however, that this approach suffers from identification problems (see Blundell and Lewbel, 1991; Coulter *et al.*, 1992a; Johnson and Garner, 1995).

Another approach is to estimate the scale directly or indirectly from subjective income-evaluation data. The first approach was introduced by Van Praag (1968) and rests on certain assumptions, which lead to a cardinal log-normal Welfare Function of Income (WFI). The WFI is empirically obtained by asking respondents what amount of income they associate with different welfare levels, such as "very bad," "bad," "insufficient," "sufficient," "good," or "very good." From these income evaluation questions the parameters of the log-normal welfare function can be estimated. Given an arbitrary cost function the equivalence scale can then be derived. The WFI approach has been discussed extensively, particularly the cardinality assumption (e.g., Hartog, 1988; Seidl, 1994; see also the reply by van Praag and Kapteyn, 1994). However, as Van Praag (1991) has shown, the approach might also be used within an ordinal framework.

A second approach based on income evaluation data is the Subjective Poverty Line first proposed by Goedhart *et al.* (1977). This approach uses a Minimum Income Question (MIQ) such as "Which household income would you, in your circumstances, consider as absolutely minimal? That is to say that with less you could not make ends meet." It can be shown that the variation in minimum income is best explained by family size and current household income (for recent results, see e.g. Van den Bosch *et al.*, 1993; Garner and De Vos, 1995; and, for Germany, Plug *et al.*, 1997). Minimum income can be regarded as the realization of a point on a cost function from which equivalence scales can be derived.

Although both approaches contribute much to the fields of poverty research and equivalence scales, some problems still remain. Apart from theoretical assumptions, which need not be discussed here, both approaches often request respondents to evaluate their current income in relation to situations they have never experienced. However, in some surveys the MIQ is asked in a way which comes closer to the respondents' real situation (Garner *et al.*, 1997, discuss some of the problems comparing different questions of subjective assessment of economic well-being). Another problem is that nearly all of these analyses use cross-sectional data and thus are not able to control for different interpretations of the evaluation questions by the respondents. A critical review of the MIQ approach by Garner and de Vos (1995) shows that across countries respondents do not necessarily associate the same welfare levels in answering this question.

In this paper, another method based on subjective income evaluation is introduced. The equivalence scale elasticity will be estimated from individual panel data on income satisfaction. Satisfaction data are only seldom used to derive equivalence scales. An exception is the analysis of Morissette and Poulin (1991) who use Canadian cross-section data.[2] However, their approach is different from the present one, where the elasticity of the scale is estimated directly. Satisfaction with current household income is recorded on an ordinal scale (from completely dissatisfied to completely satisfied). Thus, the respondents are requested to evaluate their current household income. In addition, panel estimation methods allow for controlling different scale use by the respondents, which might be a problem when using satisfaction scales (see, e.g. Stinson, 1997). The basic idea of the paper—to estimate scale elasticity from income-satisfaction data—will be outlined in Section 2. Section 3 describes the data from the German Socio-Economic Panel Study (GSOEP) and presents the results. The results will be discussed in Section 4.

2. Estimating Scale Elasticity from Satisfaction Data

For a long time economists were very skeptical of satisfaction data because this type of data measures stated rather than revealed preferences. However, satisfaction data (or analogously happiness data) has been more frequently used by economists in recent years. Satisfaction data was used to analyze labor market questions (e.g. Clark and Oswald, 1994; Winkelmann and Winkelmann, 1998), public choice related items (e.g. Frey and Stutzer, 2000), income, income inequality and well being (e.g. Stanovnik, 1992; Schwarze and Härpfer, 2002), and many other topics (Frey and Stutzer, 2000 give an overview).

Altogether, analysis of satisfaction data done by economists but also much earlier and recent work by psychologists (see for an overview Diener *et al.*, 1999; Frey and Stutzer, 2000; Van Praag *et al.*, 2000; Frey and Stutzer, 2001) have shown, that satisfaction is a valid measure of individual well being. Frey and Stutzer (2000, p. 159) conclude: "Happiness is a 'subjectivist' measure of individual welfare, and is much broader than the way individual utility is normally defined. . . . While happiness is not derived from actual behavior, it is systematically and closely connected with generally accepted manifestations of well-being."

The present approach is based on a survey question on income satisfaction, which is included in a similar form in many household surveys today:

> How satisfied are you currently with the following areas of your life?
> (Please answer by using the following scale, in which 0 means completely dissatisfied, and 10 means completely satisfied.)
> How satisfied are you with your . . .
> health
> . . .
> household income
> . . .
> environmental conditions in your area?

[2]Unfortunately, the report by Morissette and Poulin (1991) was not published in a journal or a book. I am grateful to one referee for making the report available to me.

The basic hypothesis of this paper is that if individuals are to evaluate their household income, they evaluate Y^e rather than Y, because they anticipate increasing returns to scale or enjoy additional utility when they have children by choice. In other words, welfare derived from income is evaluated rather than income itself. An empirical evaluation of this assumption will be given later.

Assume that income satisfaction S^* is a continuous latent variable that cannot be observed. Assuming decreasing marginal utility (satisfaction) of income, the model can be written as:

(1) $$S_i^* = \beta_0 + \beta_1 \ln Y_i^e + X_i'\beta_2 + \varepsilon_i$$

where X is a vector of characteristics of the respondents, and ε is a well-behaved error term. Note that the size of the household is not included in X. Obviously, the specification of the model implies the assumption that family members share income equally (see Jenkins, 1991 for a discussion).

To estimate the elasticity, the basic hypothesis discussed above can be incorporated more explicitly in model (1). Remember that as a simple formulation $Y^e = Y/h^e$ model (1) can be written as

(2) $$S_i^* = \beta_0 + \beta_1 \ln\left(\frac{Y_i}{h_i^e}\right) + X_i'\beta_2 + \varepsilon_i$$

Rearranging (2) we have

(3) $$S_i^* = \beta_0 + \beta_1 \ln Y_i - \beta_1 e \ln h_i + X_i'\beta_2 + \varepsilon_i$$

e should take values between 0 and 1—note that this is not a necessary condition—and thus the estimated coefficient $\beta_1 e$ should be negative unless β_1 is expected to be positive. Parameter e, equivalence scale elasticity, can then be identified as $\beta_1 e/\beta_1$.

What is the difference between the approach presented here and other subjective approaches such as the MIQ or the WFI approach? In general, equivalence scales M are defined as the ratio of two cost or expenditure functions (see e.g. Coulter *et al.*, 1992a):

(4) $$M_i = \frac{c(u, p, a_i)}{c(u, p, a_r)}$$

where u denotes a reference well-being level, p is a price vector, a_i represents characteristics of household i, and a_r those of the reference household r. From the MIQ approach, e.g., M_i is derived as follows (see Van Praag and Van der Sar, 1988; Plug *et al.*, 1997). The responses to the MIQ are regressed on household size and current household income:

(5) $$\ln y\min_i = a_0 + a_1 \ln Y_i + a_2 \ln h_i + \varepsilon_i$$

If $y$min is treated as a common reference welfare level, the following equivalence scale can be derived from the related cost function:

(6) $$M_i = \frac{Y_i}{Y_r} = \frac{h_i^{a_2/(1-a_1)}}{h_r}$$

where the scale elasticity $a_2/(1 - a_1)$ depends on the estimated parameters for household size and current income. It would be possible to derive the scale from the income satisfaction approach analogously if a common level of well-being $\bar{S}_i^*$ could be defined (see, e.g. Morissette and Poulin, 1991). As this assumption would be very restrictive, this approach is not followed here. However, the equivalence scale derived from (2) is $(h_i/h_r)^{\beta_1^e/\beta_1}$. Thus, it is obvious that the satisfaction approach and the MIQ respectively WFI approach are close to each other because, in either case, the equivalence scale is derived from the estimated parameters on household size and household income.

The scale derived here is dependent only on household size. However, equivalence scales might reflect both economies of size and differences in household characteristics. Following Coulter *et al.* (1992b), the equivalence scale elasticity can then be written as $e = a + b(HC)$, where a is a basic scale parameter and $b(HC)$ is a function of household characteristics. An important characteristic of the household is the number of children. The utility derived from having children theoretically reduces the additional income necessary to maintain a given level of well-being.[3] The argument can be expressed by the specification $e = a - bk$, where k is the number of children. Given household size, elasticity will decrease with the number of children. Incorporating this relationship for e in (2), we have:

$$S_i^* = \beta_0 + \beta_1 \ln Y_i - \beta_1 a \ln h_i + \beta_1 b k_i \ln h_i + X_i' \beta_2 + \varepsilon_i \tag{7}$$

The parameters a and b can be identified as $\beta_1 a/\beta_1$ and $\beta_1 b/\beta_1$. The elasticity of the scale now depends on the number of children in the household and can be computed as $e = a - bk$.

The continuous latent variable S^*, however, cannot be observed. What can be observed instead is income satisfaction S measured on an ordinal scale from 0 to 10. S can derived from S^* as follows:

$$\begin{aligned} S &= 0 \quad \text{if } S^* \leq 0, \\ &= 1 \quad \text{if } 0 < S^* \leq \mu_1 \\ &= 2 \quad \text{if } \mu_1 < S^* \leq \mu_2 \\ &\ \ . \\ &\ \ . \\ &= 10 \quad \text{if } \mu_9 \leq S^* \end{aligned} \tag{8}$$

The μ's are unknown parameters which can be estimated. If it is assumed that ε is normally distributed an ordered probit model can be estimated. In the case of a logistically distributed ε the ordered logit model appears. Both can be estimated by maximum likelihood. In practice, there are virtually no differences between ordered logit and ordered probit (see, e.g. Greene, 2000).

Two problems remain. First, it has often been argued that people rate their income relative to the income of others rather than according to neoclassical utility theory. This might not be a problem as long as this effect is captured wholly by the estimated coefficient β_1, which measures the impact of income on satisfaction,

[3]Unfortunately, the data set offers no information on whether the birth of children are planned events or not. However, throughout this paper it is assumed that parents have their children by choice.

because β_1 should not be interpreted here. Second, the scale may be used by the respondents in a different way (this is analogous to the ordinal–cardinal debate in utility theory). This may lead to inconsistent estimations of the parameters and standard errors.

The panel data available here enable us to control for some of the problems mentioned above. For model (3), for example, the panel specification is:

$$S_{it}^{*} = \beta_0 + \beta_1 \ln Y_{it} - \beta_1 e \ln h_{it} + X_{it}'\beta_2 + PRICE_t\beta_3 + \eta_t'\beta_4 + \varepsilon_{it} \tag{9}$$

The model can be estimated as a random- or a fixed-effects model. In the random-effects case, the error term will be composed as $\varepsilon_{it} = \alpha_i + v_{it}$ where α_i is an individual random effect controlling for inter-individual differences in scaling and anchoring of the responses, and unobserved variables. v_{it} is the disturbance term with $E(v_{it}) = 0$. In the random-effects model, the unobserved heterogeneity is assumed to be independent from the covariates. However, this assumption of independence seems to be implausible in the current context. An alternative specification is the fixed-effects model where α_i is treated as a fixed effect and does not have to be independent from the covariates. However, little is known about fixed effects in nonlinear models (see, e.g. Arellano and Honoré, 2001). An exception is the fixed-effects binary logit model. Thus, we follow the approach by Winkelmann and Winkelmann (1998) and collapse the satisfaction variable into a satisfied/dissatisfied binary variable[4]. First, a pooled ordered logit model is estimated, in which the variance-covariance matrix is clustered by individuals to relax the assumption of independence.[5] The results of this model will then be compared with the results from a pooled binary logit model, where the variance-covariance matrix is also clustered by individuals. Finally, a fixed-effect binary logit model will be estimated. Therefore, consider the following latent model (see, e.g. Arellano and Honoré, 2001):

$$S_{it}^{*} = \alpha_i + z_{it}'\beta + v_{it} \quad i = 1, \ldots, N, t = 1, \ldots, T, \tag{10}$$

where all covariates are included in z for convenience. α_i is the constant over time fixed effect. What we observe is:

$$S_{it} \begin{cases} 1 & \text{if } S_{it}^{*} > 0 \\ 0 & \text{else.} \end{cases} \tag{11}$$

Assuming that v_{it} is distributed independently logistic, it follows that

$$P(S_{it} = 1 \mid z_{it}, \alpha_i) = \frac{\exp(\alpha_i + z_{it}'\beta)}{1 + \exp(\alpha_i + z_{it}'\beta)}. \tag{12}$$

This fixed-effects model can be estimated by conditional maximum likelihood (Arellano and Honoré, 2001). In particular, the probability of a sequence of outcomes ($S_{i1}, \ldots, S_{iT}$), conditional to $S_i = \Sigma_{t=1}^{T} S_{it}$, is

[4]The consistency of the resulting binary logit estimator does not depend on the choice of the breaking point (see Chrouchley, 1995).

[5]Thus, the pooled data are not examined as a pure series of cross-sections.

$$(13) \qquad P(S_{it},\ldots,S_{iT} \mid z_{it},\ldots,z_{iT},\alpha_i,S_i) = \frac{\prod_{t=1}^{T} \exp(z_{it}'\beta S_{it})}{\sum_{d \in D_i} \prod_{t=1}^{T} \exp(z_{it}'\beta d_t)}$$

where D_i is the set of all possible combinations of S_i ones and $T - S_i$ zeros, is independent of α_i. Note that the sample has to be restricted to observations for which S_{it} changes. To test for fixed individual effects a Hausman-test will be performed.

When estimating equivalence scales from cross-sectional data, the price vector in the cost function can be ignored. However, prices have to be taken into account when using panel data. Until now, little has been known about whether respondents evaluate nominal or real income when asked to evaluate their household income. As a first attempt, a variable *PRICE* is included in the regression. Two possible specifications will be tested, price levels and inflation rates. The panel specification also includes as additional variables fixed time effects η_t.

3. Data and Results

The data used here come from the GSOEP, which is a representative longitudinal micro-database covering a wide range of socio-economic information on random selected households in Germany. The first round of data was collected from approximately 6,000 families in the western states in 1984. After German reunification in 1989, the GSOEP was extended by about 2,200 families from the eastern states.[6] For all estimates, an unbalanced panel design covering the years 1992–99 is used. Respondents who answered at least twice are included. In addition, the sample is restricted to those respondents who filled out the household questionnaire. This restriction is necessary because only those persons gave information on both overall net household income requested by the household questionnaire and satisfaction with household income requested by the personal questionnaire. Inflation rates are computed from official statistics using a consumer price index. The inflation rates are computed separately for East and West Germany.

Table 1 first shows estimates for the coefficients of model (1) by pooled ordered logit regressions. A series of estimates is provided for various values of equivalent income computed with different values for scale elasticity *e*. All regressions include the socio-economic characteristics of the respondent, such as age, age squared, sex, employment status, education, nationality, ethnic characteristics, and fixed time effects, as additional variables.

The top part of Table 1 shows that the estimated coefficient for equivalent income depends on the elasticity set by the researcher. Moreover, it can be seen that the explanatory power of the regression (measured by a pseudo-R^2) also depends on the elasticity. R^2 first increases with decreasing elasticity. When *e* is set

[6]The GSOEP data used in this study are available as a "scientific use" file (see Wagner, Burkhauser, and Behringer, 1993). For further information please contact the German Institute for Economic Research (DIW), Berlin: http://www.diw.de/soep/.

TABLE 1

THE IMPACT OF EQUIVALENT HOUSEHOLD INCOME ON INCOME SATISFACTION (MODEL 1 ESTIMATED BY POOLED ORDERED LOGIT)

	Equivalent Income Y^e is computed using . . .					
Variable	$e = 1.0$	$e = 0.8$	$e = 0.5$	$e = 0.3$	$e = 0.2$	$e = 0.0$
ln Y^e	1.2753	1.6687	2.1005	2.0810	1.9805	1.6883
	(0.031)	(0.034)	(0.037)	(0.036)	(0.035)	(0.032)
Log-likelihood	–104840	–103879	–102631	–102429	–102547	–103048
Pseudo-R^2	0.0469	0.0557	0.0670	0.0689	0.0678	0.0632
ln Y^e	2.1044	2.1044	2.1044	2.1044	2.1044	2.1044
	(0.037)	(0.037)	(0.037)	(0.037)	(0.037)	(0.037)
ln$h$	1.3972	0.9763	0.3449	–0.0758	–0.2863	–0.7072
	(0.036)	(0.032)	(0.030)	(0.031)	(0.032)	(0.035)
Log-likelihood	–102420	–102420	–102420	–102420	–102420	–102420
Pseudo-R^2	0.0689	0.0689	0.0689	0.0689	0.0689	0.0689

Notes: No. of observations/respondents = 50416/9929. Variance-covariance matrix is clustered by individuals. Standard errors appear in parentheses. All regressions include age, age squared, sex, employment status, education, nationality of the respondents, and time effects as additional variables.

Source: GSOEP 1992–99.

lower than 0.3, however, R^2 decreases. These results support the assumption that individuals evaluate Y^e rather than Y, when they are asked for satisfaction with household income.

The regression results in the bottom part of Table 1 also include the size of the household. Here, the coefficient for equivalent household income shows the "true" effect of income on satisfaction with income: the coefficient is always the same, regardless of the elasticity chosen. This is also true for the value of R^2. Given household income the size of the household obviously serves as a correcting factor. When elasticity is equal to 1 (i.e. equivalent income corresponds to per capita income), the estimated sign of the household size variable is significantly positive: given per capita household income, satisfaction with income increases with household size because of scaling effects, or because of utility derived from having children. The estimated coefficient for the household size variable decreases, though it is still positive, until the elasticity is 0.5. When e is set lower than 0.3, the estimated sign for the household size variable becomes negative. Obviously, scaling effects implied by the equivalent income are now higher, as anticipated by the respondents. In summary, it can be argued that the elasticity which could explain most of the variance of income satisfaction lies between 0.3 and 0.5.

The elasticity can be estimated using model (3). The estimated coefficients are shown in Table 2. As before, all regressions include age, age squared, sex, education, employment status, nationality and ethnic characteristics of the respondents, and fixed time effects as additional variables. They are not shown here. The coefficients of interest are β_1 and $\beta_1 e$. Both are at least significant at the 5 percent level for all models presented. As expected, β_1 has a positive sign and $\beta_1 e$ has a negative sign. The scale elasticity e is 0.336 as estimated from the pooled ordered logit model. Including inflation rates does not change the result. The estimated coefficients for inflation rates are not significant. The same result can be obtained when including price levels instead of inflation rates.

TABLE 2

ESTIMATES OF EQUIVALENCE SCALE ELASTICITY FROM LONGITUDINAL DATA (MODEL 3)

Variable	Pooled Ordered Logit[a]		Pooled Binary Logit[a,b]	Fixed Effects Binary Logit[b]
β_1 (ln Y)	2.1044	2.1051	2.0880	2.1308
	(0.037)	(0.037)	(0.047)	(0.067)
$\beta_1 e$ (ln$h$)	–0.7072	–0.7078	–0.6974	–0.6354
	(0.035)	(0.035)	(0.041)	(0.080)
Inflation	–	0.0035	0.0073	0.0036
		(0.004)	(0.006)	(0.008)
Log-likelihood	–102420	–102420	–30028	–11041
Pseudo-R^2	0.0689	0.0689	0.1396	0.0723
No. of observations/ respondents	50416/ 9929	50416/ 9929	50416/ 9929	30418/ 5006
Hausman-Test (chi^2,df)	–	–	257.5(13)	
$e = \beta_1 e/\beta_1$	0.336	0.336	0.334	0.298
	(0.038)	(0.038)	(0.046)	(0.068)

Notes: Standard errors appear in parentheses. All regressions include a constant, age, age squared, sex, employment status, education, nationality of the respondents, and time effects as additional variables.

[a]Variance-covariance matrix is clustered by individuals.

[b]Dependent variable: 1 if satisfied (satisfaction response above 6 on the 0–10 scale), 0 if dissatisfied.

Source: GSOEP 1992–99.

In the pooled binary logit model the dependent variable is coded as 1 if the satisfaction response is above 6. This is equivalent to classifying the respondents into those who report above- and below-average satisfaction. The estimated coefficients are nearly identical to those estimated by the ordered logit model whereas the estimated standard errors are higher due to a loss of efficiency (see Table 2). The resulting scale elasticity is 0.334. Thus, the binary logit seems to provide a good approximation of the ordered logit model and thus can be compared with the fixed effects binary logit model. The fixed-effects estimation (see also Table 2) yields a scale elasticity of 0.298, which is somewhat lower than the elasticity derived from the other models. When testing the binary logit model versus the fixed-effects model using the Hausman-Test, the fixed-effects assumption cannot be rejected. Thus, it can be argued that unobserved individual effects play a crucial part when deriving equivalence scales from subjective evaluation data. This might also be true for other subjective data as in the WFI or MIQ approaches. Note, however, that further research is requested in relation to whether unobserved characteristics of the household or of the respondents are of importance.

Estimated elasticity shown by Table 2 is constant across household size. Thus, it does not matter whether a household consists of four adults or two adults and two young children. However, if there is extra utility derived from having children or economies of scale are higher for children than for adults, this result might be misleading. Thus, Table 3 depicts estimated coefficients related to model (7), where the number of children enters the regression in the form of an interaction effect. The estimated parameters $\beta_1 a$ and $\beta_1 b$ have expected signs and both are statistically significant as estimated from all three models. Equivalence scale elasticity can

TABLE 3

ESTIMATES OF EQUIVALENCE SCALE ELASTICITY FROM LONGITUDINAL DATA (MODEL 7)

Variable	Pooled Ordered Logit[a]	Pooled Binary Logit[a,b]	Fixed Effects Binary Logit[b]
β_1 ($\ln Y$)	2.1653	2.1546	2.1514
	(0.038)	(0.048)	(0.067)
$\beta_1 a$ ($\ln h$)	−0.9172	−0.9061	−0.7553
	(0.045)	(0.052)	(0.091)
$\beta_1 b$ ($k\ln h$)	0.0967	0.0940	0.0788
	(0.013)	(0.015)	(0.027)
Inflation	0.0035	0.0073	0.0033
	(0.005)	(0.006)	(0.008)
Log-likelihood	−102338	−29977	−11037
Pseudo-R^2	0.0697	0.1411	0.0726
No. of observations/respondents	50416/ 9929	50416/ 9929	30418/ 5006
Hausman-Test (chi^2,df)	–	280.1(14)	
$a = \beta_1 a/\beta_1$	0.423	0.421	0.351
	(0.050)	(0.061)	(0.079)
$b = \beta_1 b/\beta_1$	0.0446	0.0441	0.0366
	(0.007)	(0.009)	(0.014)
$e = a - bk$	$0.423–0.0446k$	$0.421–0.0441k$	$0.351–0.0366k$

Notes: Standard errors appear in parentheses. All regressions include a constant, age, age squared, sex, employment status, education, nationality of the respondents, and time effects as additional variables.

[a]Variance-covariance matrix is clustered by individuals.

[b]Dependent variable: 1 if satisfied (satisfaction response above 6 on the 0–10 scale), 0 if dissatisfied.

Source: GSOEP 1992–99.

be computed as $e = a - bk$, where k is the number of children in the household. Parameter a is something like a baseline elasticity that will be lowered b times for each child in the household. Thus, the elasticity is higher for a four-adult household than for a household with two adults and two children. The estimated results from the ordered logit model give a value of 0.423 for parameter a and a value of 0.044 for parameter b. The estimated values from the binary logit model are nearly the same. The value for parameter a as computed from the fixed-effects logit model estimates is 0.351 and thus lower than that computed from the other models. The computed value for b (0.0366) is also lower.

The scale elasticity estimated here is based on the underlying assumption of equivalence scale exactness; that is, the equivalence scale will be the same for all levels of income and utility (see, e.g. Johnson and Garner, 1995). As a first attempt to prove this assumption, an interaction term between household income and household size was included in the regressions. However, in no case was this effect significant.

4. DISCUSSION

How does the estimated elasticity fit the extensive literature on equivalence scales? Buhmann *et al.* (1988) summarized the existing literature and found that estimated scale elasticity covers a wide range, between 0.2 and 0.8. Expert scales imply the highest scale elasticity with mean around 0.7. Expenditure-oriented econometric estimates are most often higher (between 0.23 and 0.57 with mean 0.40) than estimates based on subjective evaluation data (between 0.12 and 0.36 with mean 0.24). However, some of the more recent studies on subjective scales report higher elasticities (see, e.g. Stanovnik, 1992; Van den Bosch *et al.*, 1993; Garner and de Vos, 1995).[7]

Table 4 compares different equivalent scales for various types of households. All scales shown are either used in German research on income inequality and poverty or are computed using German data. Three so-called expert scales—a scale proposed by the OECD, a scale related to the social assistance scheme in Germany, and the US-Poverty line scale—are compared with an econometric expenditure scale for Germany (see Merz *et al.*, 1993), two subjective scales estimated along the WFI and the MIQ approach for Germany (see Plug *et al.*, 1997), and the subjective scale approach presented in this paper. The top part of Table 4 depicts the weights assigned to each household type, and the bottom part shows the elasticity derived from the weights (or vice versa).

Using income-satisfaction data, estimated equivalence scale elasticity—as computed from the ordered logit regression results—is a constant value of 0.34. Controlling for children, elasticity is 0.42 and constant for households consisting of adults only. Elasticity is estimated to be lower for households where some of the members are children, and elasticity decreases slightly as the number of children increases. When controlling for unobserved fixed effects the elasticity becomes lower. However, elasticity estimated in the present paper is higher than for the other subjective scales that were also based on the GSOEP. Plug *et al.* (1997) have estimated an elasticity of 0.20 based on the WFI approach and one of 0.23 based on the MIQ approach (see Table 4).

If estimated elasticity is compared with that computed from expert scales, it can be seen, first, that for almost all expert scales the elasticity is higher and, thus, expert scales obviously underestimate the economies of scale that flow from individuals living together in a household. Second, it can be seen that almost all expert scales increase with household size—some are even bouncing around (for a discussion see Citro and Michael, 1995). This seems to be misleading from a theoretical point of view because increasing elasticity means diseconomies of scale rather than economies of scale.

In summary, it can be shown that income satisfaction data might be an interesting alternative to estimate equivalence scales from subjective data. Compared with other subjective approaches, respondents evaluate current household income, rather than hypothetical situations as in the WFI or the MIQ approaches. Satisfaction data are easy to collect and the number of non-respondents and implausible answers are low compared with the WFI or MIQ approaches. Compared with the other subjective approaches, a higher elasticity is derived from satisfaction data. Further, it has been shown that unobserved fixed effects seem to play a crucial

[7]The equivalence scale estimated by Garner and De Vos (1995) using the MIQ implicates an elasticity of 0.49 for the U.S. Following the income satisfaction approach of Morissette and Poulin (1991) an elasticity of even 0.6 can be computed. This is probably due to the fact that they regress satisfaction on yearly disposable income whereas it is not clear what income the respondents have evaluated.

TABLE 4

COMPARISON OF DIFFERENT EQUIVALENCE SCALES COMPUTED FROM GERMAN DATA OR USED IN GERMAN RESEARCH FOR CERTAIN TYPES OF HOUSEHOLD: WEIGHTS AND ELASTICITY

Type of Household/ Scale	1 Adult	2 Adults	3 Adults	4 Adults	2 Adults 1 Child	2 Adults 2 Children	2 Adults 3 Children
Weights							
Per capita (= h)	1.0	2.0	3.0	4.0	3.0	4.0	5.0
OECD Scale[a]	1.0	1.5	2.0	2.5	1.8	2.1	2.4
Social Assist. Germany[b]	1.0	1.8	2.6	3.4	2.45	3.1	3.75
US-Poverty Line[c]	1.0	1.29	1.57	2.01	1.55	1.99	2.35
Econom. Expenditure[d]	1.0	1.49	1.73	1.89	1.61	1.72	1.84
Subjective scales:							
WFI[e]	1.0	1.15	1.25	1.32	1.25	1.32	1.39
MIQ[f]	1.0	1.17	1.28	1.37	1.28	1.37	1.44
Satisfaction[g]: e	1.0	1.26	1.45	1.59	1.45	1.59	1.72
$e = a - bk$	1.0	1.34	1.59	1.79	1.51	1.59	1.60
Satisfaction with Fixed-effects[h]: e	1.0	1.23	1.39	1.51	1.39	1.51	1.61
$e = a - bk$	1.0	1.28	1.47	1.63	1.41	1.47	1.48
Elasticity (e)[i]							
OECD Scale[a]	–	0.58	0.63	0.66	0.54	0.53	0.54
Social Assist. Germany[b]	–	0.84	0.87	0.88	0.82	0.82	0.82
US-Poverty Line[c]	–	0.36	0.41	0.50	0.39	0.49	0.53
Econom. Expenditure[d]	–	0.57	0.50	0.46	0.43	0.39	0.38
Subjective scales:							
WFI[e]	–	0.20	0.20	0.20	0.20	0.20	0.20
MIQ[f]	–	0.23	0.23	0.23	0.23	0.23	0.23
Satisfaction[g]: e	–	0.34	0.34	0.34	0.34	0.34	0.34
$e = a - bk$	–	0.42	0.42	0.42	0.38	0.34	0.29
Satisfaction with Fixed-effects[h]: e	–	0.30	0.30	0.30	0.30	0.30	0.30
$e = a - bk$	–	0.35	0.35	0.35	0.31	0.28	0.24

Notes:

[a]First adult has weight 1.0, every further adult 0.5, children 0.3.

[b]First adult has weight 1.0, further adults 0.8, children 1–7 years old 0.5, children 8–14 years old 0.65, children 15–18 years old 0.90.

[c]See Merz *et al.*, 1993.

[d]Equivalence scale estimated along an extended linear expenditure system (source: Merz *et al.*, 1993).

[e]Based on the Income Evaluation Question Approach estimated from GSOEP data 1992 (source: Plug *et al.*, 1997).

[f]Based on the Minimum Income Question Approach estimated from GSOEP data 1992 (source: Plug *et al.*, 1997).

[g]Estimated parameters from ordered logit model (GSOEP 1992–99, see Tables 2 and 3).

[h]Estimated parameters from fixed effects logit model (GSOEP 1992–1999, see Tables 2 and 3).

[i]In the present paper, e is estimated directly. For the other scales, e is computed as ln(*weights*)/ln(h).

part when deriving scale elasticity from subjective data. Further research should be done to analyze the origin of the fixed-effects. The satisfaction data approach could also be tested for a more flexible specification in future. In addition, it has to be proved further whether the underlying assumption of equivalence scale exactness holds.

REFERENCES

Arellano, M. and B. Honoré, "Panel Data Models: Some Recent Developments," in J. J. Heckman and E. Leamer (eds), *Handbook of Econometrics*, Volume 5, Elsevier, Amsterdam, 3229–96, 2001.

Blundell, R. and A. Lewbel, "The Information Content of Equivalence Scales," *Journal of Econometrics*, 50, 49–68, 1991.

Bojer, H. and J. A. Nelson, "Equivalence Scales and the Welfare of Children: A Comment on 'Is There Bias in the Economic Literature on Equivalence Scales?'," *Review of Income and Wealth*, 45, 531–4, 1999.

Buhmann, B., L. Rainwater, G. Schmaus, and T. J. Smeeding, "Equivalence Scales, Well-Being, Inequality, and Poverty: Sensitivity Estimates Across Ten Countries using the Luxembourg Income Study (LIS) Database," *Review of Income and Wealth*, 34, 115–42, 1988.

Burkhauser, R. V., T. M. Smeeding, and J. Merz, "Relative Inequality and Poverty in Germany and the United States Using Alternative Equivalence Scales," *Review of Income and Wealth*, 42, 381–400, 1996.

Chrouchley, R., "A Random-effects Model for Ordered Categorical Data," *Journal of the American Statistical Association*, 90, 489–98, 1995.

Citro, C. F. and R. T. Michael (eds), *Measuring Poverty. A New Approach*, National Academy Press, Washington, 1995.

Clark, A. E. and A. J. Oswald, "Unhappiness and Unemployment," *Economic Journal*, 104, 648–59, 1994.

Coulter, F. A. E., F. A. Cowell, and St. P. Jenkins, "Equivalence Scale Relativities and the Extent of Inequality and Poverty," *The Economic Journal*, 102, 1067–82, 1992a.

———, "Differences in Needs and Assessment of Income Distributions," *Bulletin of Economic Research*, 44, 77–124, 1992b.

Diener, E., E. Suh, R. Lucas, and H. Smith, "Subjective Well-Being: Three Decades of Progress," *Psychological Bulletin*, 125, 276–303, 1999.

Frey, B. S. and A. Stutzer, "Maximizing Happiness?," *German Economic Review*, 1, 145–67, 2000.

———, "What can Economists Learn from Happiness Research?," Center for Economic Studies and Ifo Institute for Economic Research, Working Paper 503, Munich, 2001.

Garner, Th. I. and K. de Vos, "Income sufficiency v. poverty. Results from the United States and The Netherlands," *Journal of Population Economics*, 8, 117–34, 1995.

Garner, Th. I., L. Stinson, and St. Shipp, "Subjective Assessments of Economic Well-Being and Preliminary Findings from Miami," *Proceedings of the Joint Statistical Meetings* (http://www.amstat.org/sections/srms/Proceedings/papers/1997_156.pdf), 1997.

Goedhart, Th., V. Halberstadt, A. Kapteyn, and B. M. S. van Praag, "The Poverty Line: Concept and Measurement," *Journal of Human Resources*, 12, 503–20, 1977.

Greene, W. H., *Econometric Analysis*, 4th edition, Prentice-Hall International, London, 2000.

Hartog, J., "Poverty and the Measurement of Individual Welfare," *Journal of Human Resources*, 23, 243–66, 1988.

Jenkins, St. P., "Poverty Measurement and the Within-Household Distribution," *Journal of Social Policy*, 20, 357–83, 1991.

Johnson, D. S. and Th. I. Garner, "Unique Equivalence Scales. Estimation and Implications for Distributional Analysis," *Journal of Income Distribution*, 4, 215–34, 1995.

Merz, J., Th. Garner, T. M. Smeeding, J. Faik, and D. Johnson, "Two Scales, One Methodology—Expenditure Based Equivalence Scales for the United States and Germany," Syracuse University, Cross-National Studies in Aging, Program Project Paper No. 8, Syracuse, 1993.

Morissette, R. and S. Poulin, "Income Satisfaction Supplement: Summary of Four Survey Years, English language version," Statistics Canada, Canada, 1991.

Nelson, J. A., "Methods of Estimating Household Equivalence Scales: An Empirical Investigation," *Review of Income and Wealth*, 38, 295–310, 1992.

———, "Household Equivalence Scales: Theory versus Policy?," *Journal of Labor Economics*, 11, 471–93, 1993.

Plug, E. J. S., P. Krause, B. M. S. van Praag, and G. G. Wagner, "Measurement of Poverty—Exemplified by the German Case," in Ott, N. and G. G. Wagner (eds), *Income Inequality and Poverty in Eastern and Western Europe*, Springer-Physica, Heidelberg, 69–89, 1997.

Schwarze, J. and M. Härpfer, "Are People Inequality Averse, and Do They Prefer Redistribution by the State? Evidence from German Longitudinal Data on Life Satisfaction," Institute for the Study of Labor (IZA), Discussion Paper No. 430, Bonn, 2002.

Seidl, Ch., "How Sensible Is the Leyden Individual Welfare Function of Income?," *European Economic Review*, 38, 1633–59, 1994.

Stanovnik, T., "Perception of Poverty and Income Satisfaction," *Journal of Economic Psychology*, 13, 57–69, 1992.
Stinson, L., "Using the 'Delighted/Terrible' Scale to Measure Feelings About Income," *Proceedings of the Joint Statistical Meetings* (http://www.bls.gov/ore/abstract/st/st960100.htm), 1997.
Van den Bosch, K., T. Callan, J. Estivill, P. Hausman, B. Jeandidier, R. Muffels, and J. Yfantopoulos, "A Comparison of Poverty in Seven European Countries and Regions using Subjective and Relative Measures," *Journal of Population Economics*, 6, 235–59, 1993.
Van Praag, B. M. S., *Individual Welfare Functions and Consumer Behavior*, Amsterdam, 1968.
———, "Ordinal and Cardinal Utility," *Journal of Econometrics*, 50, 69–89, 1991.
Van Praag, B. M. S., P. Frijters, and A. Ferrer-I-Carbonell, "A Structural Model of Well-being," University of Amsterdam, Tinbergen Institute, Tinbergen Institute Discussion Paper TI 2000-053/3, Amsterdam, 2000.
Van Praag, B. M. S., A. J. M. Hagenaars, and H. van Weeren, "Poverty in Europe," *Review of Income and Wealth*, 28, 345–59, 1982.
Van Praag, B. M. S. and A. Kapteyn, "How Sensible is the Leyden Individual Welfare Function of Income? A Reply," *European Economic Review*, 38, 1817–25, 1994.
Van Praag, B. M. S. and N. L. Van der Sar, "Household Cost Functions and Equivalence Scales," *Journal of Human Resources*, 23, 193–210, 1988.
Wagner, G. G., R. V. Burkhauser, and F. Behringer, "The English Language Public Use File of the German Socio-Economic Panel," *Journal of Human Resources*, 28, 429–33, 1993.
Winkelmann, L. and R. Winkelmann, "Why are the Unemployed so Unhappy? Evidence from Panel Data," *Economica*, 65, 1–15, 1998.