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# The estimation of pensioner equivalence scales using subjective data \*

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## Abstract

This paper uses panel data on pensioners' subjective evaluations of their financial positions to construct equivalence scales for pensioners. A pensioner couple is estimated to require an income 44% higher than a comparable single pensioner to reach the same standard of living. This is significantly less than the equivalence scale value implied by the ratio of state pension rates, the McClements equivalence scale value and the scale value derived from Engel curve estimation for food expenditure using the same data source. The estimated equivalence scale value is robust to variations in the definition of the pensioner sample, the measurement of income and the econometric model used.

**Key words:** equivalence scales, pensioner incomes, pensions, subjective data, ordered response models, Engel curves.

**JEL classification:** D12, D63, H55, J14, C23, C25, C14.

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# 1 Introduction

What level of household income is required to give a pensioner couple the same level of welfare as a single pensioner on a given level of income? Equivalence scale questions of this type are of key importance to many policy judgements. Such a scale is required, for example, to set the relative magnitudes of state pensions for couples and singles and in the setting of other household composition-contingent state benefits. More generally, the construction of consistent inequality or poverty measures requires comparison of households with different incomes and demographics, and hence an equivalence scale (implicitly or explicitly) with which to deflate household income.

Indeed welfare comparisons across households of different demographic composition are central to public policy analysis. The measurement of welfare (or a proxy for it) lies at the heart of public policy analysis (Slesnick, 1998). Analysis of inequality and poverty, tax and benefit system shifts, government intervention programs, pension and health care reform, and other public policies must eventually examine how these policies affect the well-being of individuals. Inferring individual living standards from household data requires an estimate of the economies of scale. There is however considerable disagreement on how to estimate income relativities of the type in the iso-welfare question posed at the start.

The UK basic state retirement pension contains an equivalence scale value of 1.6 for pensioner couples relative to single pensioners. This value was set when the state pension was introduced in 1948 and has remained unchanged (apart from rounding) over the 60 years since. The economies of scale in consumption for pensioners may well have changed over the last 60 years and the appropriate relativity is an important policy question. Most of the empirical literature on equivalence scales has focused on the costs of children. There has been much less investigation of economies of scale for adults and even less focusing on those for

pensioners. This paper investigates the estimation of the appropriate equivalence scale for pensioners.

The most commonly used approach to the estimation of equivalence scales uses expenditure data in a revealed preference framework. Engel's method using food expenditure shares is the oldest and most widely known such method. However there is an identification problem with this approach – equivalence scales cannot be identified from demand functions without additional strong identifying assumptions (see for example, Pollak and Wales, 1979, Pollak, 1991, Blundell and Lewbel, 1991). An alternative approach that has been suggested in the literature uses subjective data. This approach is the main focus here. This paper uses data on pensioners' evaluations of their financial position from 14 waves of the British Household Panel Survey to estimate the equivalence scale value for pensioners. Estimates using Engel's method and food expenditure information from the same data source are also presented to provide comparison.

The next section lays out the framework for equivalence scale estimation in general and also that specific to using subjective evaluations of the respondent's financial position. Section 3 describes the data used. Section 4 describes the econometric model and estimators used. Results are presented in Section 5 and conclusions in Section 6.

## 2 The equivalence scale framework

Couples typically need more resources than singles to achieve the same standard of living, but less per capita. An equivalence scale attempts to measure this difference, and more generally differences between households of different demographic type.<sup>1</sup> It allows comparisons to be made between households with different demographic structures by, for example, dividing the incomes of couples by a particular

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<sup>1</sup>See Browning (1992) and Nelson (1993) for surveys of the literature on equivalence scales with particular reference to those for children.

scale value to bring them into *equivalence* (in welfare terms) with those of singles.

The wide disagreement between the values of available equivalence scales has been pointed out by Atkinson and Bourguignon (1990) among others. They cite a survey which tabulates 44 different estimated scale values and in terms of a couple compared to a single person gives a range from 1.06 to 2.04.<sup>2</sup> Using these for comparison of incomes across households is roughly equivalent to using unadjusted household income at one end of the range and using per-capita income (i.e. deflating by household size) at the other end. In the current context, the composition of pensioner poverty, for example, is sensitive to the choice of equivalence scale. As an illustration, if the bottom quartile is used as the poverty threshold, 43% of pensioners in poverty are single pensioners when based on per-capita income, while 87% are when based on unadjusted household income.<sup>3</sup>

The UK semi-official equivalence scale until very recently was based on the work of McClements (1977), which gave values close to those implied by the levels of Supplementary Benefits in place at the time. Defects in the model, estimation method, and data were pointed out by Muellbauer (1979) and others since. However the scale has remained in widespread use for around 30 years. This scale gives a value of 1.64 for a couple relative to a single adult. The annual “households below average income” (HBAI) report by the Department of Work and Pensions (DWP) had used the McClements scale since the start of the series of reports. For the 2005/06 report (published in May 2007) the Department switched to using the “modified OECD scale” used by Eurostat. It uses a value of 1.50 for a couple relative to a single adult. The methodological change in this high-profile report series (and elsewhere) has re-ignited the long-running debate on the choice of equivalence scale and the appropriate methodology for constructing it. Reducing the equivalence scale value for couples relative to singles (as for example in

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<sup>2</sup>These are typically based on analysis of those of working age.

<sup>3</sup>The use of the pensioner lower quartile for each wave corresponds on average to 60% of the overall median, the poverty threshold standardly used by the UK government.

this shift) has the effect of reducing the equivalised incomes of singles relative to couples and hence increasing the relative percentage of singles to couples below any poverty threshold.

The most common methodology for constructing equivalence scales is based on expenditure data, but, as stated in Section 1, requires strong identifying assumptions.<sup>4</sup> Engel's method uses the proportion of income spent on food as the household (inverse) welfare proxy.<sup>5</sup> Its use derives from the initial empirical observations that food share declines with income holding household size constant and increases with household size holding income constant. As Deaton (1997) points out, it is important to be clear about the identifying assumptions on which any method rests. In the current context the identifying assumption is that a pensioner couple and a single pensioner are equally well-off if, and only if, they devote the same fraction of their expenditure or income to food. Although in common usage, the approach is a controversial one. Deaton and Paxson (1998) note that “estimates of economies of scale that are derived by Engel's method have no theoretical underpinning and are identified by an assertion that makes no sense” (p. 903). They suggest replacing the Engel approach by a more appropriate model in which economies of scale are attributed to public goods, but find that the implications of this model are clearly contradicted by the data.

Data on individuals' subjective evaluations of their financial position provide an alternative way to estimate equivalence scale relativities.<sup>6</sup> Two main types of subjective information have been utilized in this branch of the literature. The

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<sup>4</sup>Another common approach uses “expert scales” based on the opinions of a panel of experts in the area. Criticisms of this approach include: that it lacks theoretical justification and its basis is ad hoc; that the resultant scale is influenced by the value judgements of the experts; and that, because experts differ in their opinions, very different equivalence scale values have been produced by different panels of experts.

<sup>5</sup>In addition to the share of income spent on food, other proxies of this type have been suggested on similar grounds, such as the share of income saved or the share of income spent on “necessities” (see Nelson, 1993, for further discussion).

<sup>6</sup>See Bradbury (1989) for a review of the early literature on this approach. The use of additional subjective data is one of the possible responses suggested by Blundell and Lewbel (1991) to the underidentification problem stated in Section 1.

first, sometimes labelled the “Leyden method”, asks respondents how much income they think a family of their type requires to reach a specified level of satisfaction or well-being (see for example Kapteyn and van Praag, 1976, and van Praag, 1991). The hypothetical nature of the questions has been pointed to as a disadvantage of this approach – respondents may have little experience of such levels.<sup>7</sup>

The second type of data used asks respondents to evaluate their level of satisfaction with their own income or standard of living. (Dubnoff et al., 1981, is an early example, Bellemare et al., 2002, Charlier, 2002, and Schwarze, 2003, more recent ones.) An advantage of using this type of data is that the respondents evaluate their own financial positions rather than hypothetical situations.

This paper uses this latter approach and focuses on the equivalence scale value for pensioner couples relative to single pensioners. In this context the approach supposes that a monotonic function (which does not need to be specified) of an individual’s welfare level or standard of living,  $w$ , can be modelled by the following linear function:

$$g(w_{it}) = \gamma_1 \ln z_{it} + \gamma_2 d_{it} + q'_{it} \delta + \varepsilon_{it} \quad (1)$$

where  $z_{it}$  is household income,  $d_{it}$  is a dummy variable taking the value 1 for pensioner couples and 0 for single pensioners and  $q_{it}$  is a vector of other influences (possibly including other demographic characteristics of the household).<sup>8</sup> The disturbance term  $\varepsilon_{it}$  is taken to represent other unobserved influences.

The equivalence scale relativity for the two groups is then given as follows. Define the ratio  $\lambda$  to be such that, holding all other factors constant, a pensioner couple with income  $\lambda z^o$  would have the same level of welfare as a single pensioner

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<sup>7</sup>Leyden method equivalence scales based on hypothetical questions of this type tend to be much flatter than those from other methods (see for example Bradbury, 1989).

<sup>8</sup>The standard specification (not specifically for pensioners) would use an equation with  $d$  replaced by variables for the numbers of adults and children (possibly age-specific) in the household.

with income  $z^o$ . Then  $\gamma_1 \ln z^o + q'\delta + \varepsilon = \gamma_1 \ln(\lambda z^o) + \gamma_2 + q'\delta + \varepsilon$  and hence

$$\lambda = \exp\left(-\frac{\gamma_2}{\gamma_1}\right). \quad (2)$$

This gives an equivalence scale value for a pensioner couple relative to a single pensioner, i.e. it expresses the number of “single equivalents” for the couple, or one plus the equivalence scale value for the second person in the couple.

### 3 Data description

The empirical analysis in this paper is based on data from the British Household Panel Survey (BHPS), which contains a nationally representative sample of households whose members are re-interviewed each year.<sup>9</sup> Waves 1–14 of the survey are used in this paper. Interviews for wave 1 were conducted mainly in 1991, those for wave 14 mainly in 2004. Only data from the main BHPS are used.

#### 3.1 The definition of income

Standard practice in the literature is to measure an individual’s current economic position by the equivalised value of weekly net (disposable) *household* income of the household to which he or she belongs. An individual’s standard of living depends not only on his or her own income, but also on the income of other members of the household. The variable  $z$  is therefore taken to be current net (disposable) household income expressed in per week terms.<sup>10</sup> An annual net income variable is also examined in the analysis below to investigate the robustness of the estimates.

Net income is defined as in the DWP’s HBAI reports. Full details of the construction of this variable on the BHPS are given in Levy et al. (2006). It is the sum across all household members of: cash income from all sources (income

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<sup>9</sup>See Taylor (2007) and <http://www.iser.essex.ac.uk/ulsc/bhps/> for details.

<sup>10</sup>The time period over which current income components are measured is the month prior to the interview or the most recent relevant period (except for employment earnings which are “usual earnings”). See Levy et al. (2006) for details.

from employment and self-employment, investments and savings, private and occupational pensions, and other market income, plus cash social security and social assistance receipts and private transfers (e.g. maintenance)) minus direct taxes (income tax, employee National Insurance Contributions, local taxes such as council tax and the community charge) and any occupational pension contributions. A strength of this data source is the very detailed information provided on the components of household income.

### **3.2 Self-evaluated financial position**

BHPS respondents are asked to evaluate their current financial position. In particular they are asked: “How well would you say you yourself are managing financially these days?” and given a 5-point scale of answers to select from, ranging from “living comfortably” to “finding it very difficult”. The distribution of responses in the pensioner sample is given in the first row of Table 1.<sup>11</sup> 40% of pensioners respond that they are living comfortably and a further 28% that they are “doing alright”. At the other end 5% say that they are finding it quite or very difficult financially, and 28% say that they are “just about getting by”.

[Table 1 near here]

Distributions by household type, gender, and income are also given in Table 1. Subjective financial evaluations are typically higher for pensioner couples than for single pensioners and higher for male single pensioners than female single pensioners. 43% of pensioner couples are “living comfortably” compared with 35% of single pensioners and 40% of single men are compared with 33% of single women. There are considerable differences by income quartiles. For pensioner

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<sup>11</sup>A pensioner is taken to be a person of state pension age or above (65 for men, 60 for women). A pensioner couple contains at least one pensioner. The sample used in Table 1 is restricted to households containing only either a single pensioner or a pensioner couple. This excludes roughly 10% of the sample, mainly pensioners living with their children. The term single pensioner refers to the fact that they are living in a single-person household rather than their marital status.

couples 66% of those in the top income quartile are “living comfortably” compared with 22% in the bottom quartile. At the other end 54% of those in the bottom quartile are “just about getting by” or finding it difficult or very difficult compared with 9% in the top decile. There are similar differences for single pensioners.

Before using the data to estimate a suitable model, Figure 1 presents kernel density estimates of the probability density functions of the log of household net income for pensioner couples and single pensioners. Each panel of the figure presents these for a given value of the subjective financial position measure.<sup>12</sup> Within each of the three panels the density estimates are fairly consistent with the hypothesis that the income difference between pensioner couples and single pensioners is adequately represented by a horizontal shift in the density function, i.e. with a constant difference in the mean of log income, particularly for the top two categories of the subjective measure. This issue is also investigated in the context of the econometric model used below.

[Figure 1 near here]

## 4 Econometric models and estimation

The responses on financial position described in the previous section and used to estimate the parameters in equation (1) are discrete and ordered. The commonly used ordered probit model is taken as the starting point here. A range of alternative models and estimators is also considered below. The ordered probit model can be viewed as based on a linear equation for a latent dependent variable:

$$y_{it}^* = x'_{it}\beta + \varepsilon_{it} \quad (3)$$

for  $i = 1, \dots, N$ , and  $t = 1, \dots, T_i$ , where, in terms of equation (1),  $x = (\ln z, d, q')'$  is a vector of regressors,  $\beta = (\gamma_1, \gamma_2, \delta')'$  is a vector of parameters and  $\varepsilon_{it}$  is a

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<sup>12</sup>The Epanechnikov kernel with Silverman’s plug-in estimate for bandwidth is used. The bottom two categories of the subjective measure have much smaller sample sizes.

normally distributed random error term, assumed to be independently distributed across  $i$ . Time-variation in  $\beta$  is also considered. The observed dependent variable,  $y_{it}$  takes one of the values  $\{1, 2, \dots, J\}$ , and is related to  $y_{it}^*$  by:

$$y_{it} = j \text{ if } \alpha_{j-1} \leq y_{it}^* < \alpha_j \quad (4)$$

with the  $\alpha_j$  being parameters such that  $\alpha_1 < \alpha_2 < \dots < \alpha_{J-1}$  (and where  $\alpha_0 = -\infty$  and  $\alpha_J = +\infty$  are introduced to simplify the notation). The model is estimated by Maximum Likelihood. In this class of models a transformation of the cumulative probabilities is taken to be a linear function of the  $x$ -variables with only the intercept in this function differing across the categories:

$$\Phi^{-1} \{\Pr[y_{it} \leq j]\} = \alpha_j - x'_{it}\beta \quad (5)$$

where  $\Phi$  is the CDF of a standard normal. The impact of a particular  $x$ -variable on the transformed cumulative probability is taken to be the same for each threshold. The appropriateness of this “parallel lines” assumption in the current context and robustness of results to relaxing it are examined below.

Under the assumption that  $\varepsilon_{it}$  normally distributed, the pooled ordered probit estimator is consistent and asymptotically normally distributed without requiring either that  $x_{it}$  be strictly exogenous or that the  $\varepsilon_{it}$  be serially independent. An important feature is that it retains this consistency and asymptotic normality even if the  $\varepsilon_{it}$  are arbitrarily serially correlated. However, except under strong additional assumptions, the usual standard errors and test statistics are not valid – robust standard errors that account for the “clustering” are required.

The model can incorporate unobserved individual-specific heterogeneity by specifying  $\varepsilon_{it} = c_i + u_{it}$ . Amongst other things this would allow for the possibility that different individuals interpret the categories differently and hence have different thresholds between them, providing these differences are constant over time. The standard random effects ordered probit model adds the assumptions

that  $c_i|x_i \sim N(0, \sigma_c^2)$  and that  $y_{i1}, \dots, y_{iT}$  are independent conditional on  $(x_i, c_i)$  and can be estimated by conditional ML. In this specification (under the normalization of a unit variance for  $u$ ) the correlation of the composite error term,  $\varepsilon_{it}$ , across any two time periods is given by  $\rho = \sigma_c^2/(\sigma_c^2 + 1)$ . This is a very restrictive form for the correlation structure of the  $\varepsilon_{it}$ .

Wooldridge (2002, p. 486) makes the useful observation that pooled ordered probit can be used to consistently estimate  $\beta/(\sigma_c^2 + 1)^{1/2}$ , the average partial effect, without requiring the assumption that  $y_{i1}, \dots, y_{iT}$  are independent conditional on  $(x_i, c_i)$ . This is sufficient in the current context, since only an estimator of the ratio of two elements of  $\beta$  is required for the estimation of the equivalence scale ratio  $\lambda$  (see (2)). Pooled probit therefore provides a consistent estimator of  $\lambda$  without this conditional independence assumption (which the random effects ordered probit estimator requires). Given these properties, the pooled ordered probit estimator is used as the main estimator in the next section, but random effects estimates are also presented for comparative purposes.

A “correlated random effects” ordered probit model can be used to allow  $c_i$  to be correlated with  $x_i$ . Following Mundlak (1978), the mean of the distribution of  $c_i|x_i$  is taken to be a linear function of  $\bar{x}_i$ , the average of the  $x_{it}$  over  $t$ .<sup>13</sup> This is straightforward to implement by adding  $\bar{x}_i$  to the list of regressors. However again consistency is lost if the serial correlation is more general than that given by the equi-correlation assumption, i.e. if the  $u_{it}$  are not independent. The problem of potential correlation between  $c_i$  and  $x_i$  can also be addressed in linear models by using a “fixed effects” estimator. The fixed effects estimator for probit and related non-linear models suffers from the incidental parameters problem and is not consistent for fixed  $T$ . However Greene (2004) provides much useful evidence on the finite sample properties of fixed effects estimators in non-linear models and this evidence suggests that the fixed effects estimator may be of interest for

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<sup>13</sup>The Chamberlain (1982) alternative using the full set of  $x_{it}$  is also considered below.

comparison in the current context.

The above estimators all assume a probit form, that is to say they take  $\varepsilon$  in equation (3) to be normally distributed. Alternative specifications of this distribution (both parametric and semi-parametric) are also considered in order to examine the robustness of the estimate of  $\lambda$  to this distributional choice. Three alternative parametric models are considered: the ordered logit model and two models based on skewed distributions. I also examine two semi-parametric estimators, not requiring a specific parametric distributional assumption for  $\varepsilon$ . The first is constructed using the “semi-nonparametric” (SNP) series estimator of Gallant and Nychka (1987). The second uses a flexible polynomial transformation to normality (see Stewart, 2006, for fuller details of both estimators). Both estimators approximate the unknown density using polynomial series expansions. Providing it satisfies certain smoothness conditions, the unknown density can be approximated arbitrarily closely in both estimators by increasing the choice of polynomial order.

## 5 Results

### 5.1 Main specification

Before using the ordered probit specification described in the previous section, a further non-parametric examination of the data is potentially useful. Kernel regressions for cumulative probabilities of the subjective measure as a function of the log of household net income are presented for pensioner couples and single pensioners in Figure 2.<sup>14</sup> Separately for couples and singles these estimate the relationships between these cumulative probabilities and log income without imposing any functional form. For each pensioner group the distance between the estimated lines is fairly constant. Hence these kernel regressions are broadly

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<sup>14</sup>As with Figure 1 splits at the lower thresholds are not used because of the small relative frequencies in the bottom two categories.

supportive of the parallel lines assumption (as in equation (5)) that underlies the ordered probit models. This issue is also investigated below in the context of the parametric model.

[Figure 2 near here]

Estimates for the basic pooled ordered probit specification (with cluster-adjusted standard errors) are given in the first row of Table 2. This specification gives an estimate of the equivalence scale value,  $\lambda$ , of 1.440 (with a cluster-adjusted standard error of 0.052): pensioner couples require a net household income that is 44% higher than comparable single pensioners to produce the same self-evaluation of their financial position. This is significantly lower than the equivalence scale value implied by the ratio of state pension rates (1.60) and the McClements scale ratio (1.64), but not significantly different from the value of 1.50 used in the modified OECD scale.<sup>15</sup>

[Table 2 near here]

This estimate of  $\lambda$  is robust to various aspects of sample and variable definition. Income is defined as in section 3.1 and deflated using the “all items RPI excluding council tax”. A full set of wave dummies is included in each specification, so the price deflation effectively only adjusts for within-wave differences in interview date (and has little effect).<sup>16</sup> Controls are also included for gender, age and region.<sup>17</sup> Looking at the estimated effects of these controls, respondents’ evaluations of how

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<sup>15</sup>Tests of equality give p-values of 0.002, 0.0001 and 0.248 respectively.

<sup>16</sup>If either the all-items RPI or the average earnings index is used,  $\hat{\lambda} = 1.439 (.052)$ . If instead a set of month of interview dummies is included,  $\hat{\lambda} = 1.438 (.052)$  and the month dummies are individually and jointly insignificant.

<sup>17</sup>The selected specification contains two binary variables for ages 70-79 and ages 80+ (the base being <70). This is an acceptable parsimonious categorical representation: when tested against a fuller specification with 34 binary variables for individual integer ages, the test statistic has a p-value of 0.147. If a linear age term is used,  $\hat{\lambda} = 1.432 (.052)$ . Interactions between these three factors are found to be statistically insignificant. Tests of this give p-values of 0.108 for those between gender and age, 0.693 for those between gender and region, and 0.188 for those between age and region. The overall test of the interaction terms gives a p-value of 0.207.

they are managing financially rise with age, holding other factors (including income) constant. For the same level of these other factors (including income) those aged 70-79 are 4.6 percentage points less likely to be “just about getting by” or worse than those aged below 70, and those aged 80+ are 13.8 percentage points less likely. The effect of gender on these financial evaluations is insignificantly different from zero (and numerically small). Other things equal (including income), respondents in London and the south east and in Wales give significantly lower evaluations than those in the other regions of the country.

Pensioner couples are defined as those with at least one of the couple above the state pension age.<sup>18</sup> Only households containing just the pensioner couple or single pensioner are included in the sample in this specification.<sup>19</sup> The income measure used is current net household income for the month prior to interview or most recent relevant period converted into per week terms.<sup>20</sup> The income variable contains a small number of extreme values. The data were trimmed to exclude observations in the top and bottom 1% of the income distribution for each wave.<sup>21</sup>

Two factors that may be important to pensioners’ responses, housing and health, are excluded from the main specification, but subjected to an extensive examination below. Housing is excluded from this baseline model because there is likely to be considerable measurement error in the valuation of housing for owner-occupiers. Estimates which take account of housing wealth are considered in Section 5.3 below. Responses on health are potentially endogenous and are therefore not controlled for in this model. Estimates which do condition on

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<sup>18</sup>Restricting the sample to those who are themselves of pension age gives  $\hat{\lambda} = 1.435 (.051)$ .

<sup>19</sup>If pensioners in households containing other members in addition to the pensioner couple or single pensioner are added, with additional controls for the number of additional adults and the number of children in the household,  $\hat{\lambda} = 1.443 (.050)$ . If those who are the partners of pensioners are also added  $\hat{\lambda} = 1.448 (.050)$ .

<sup>20</sup>If annual net household income, with reference period the 12 months interval up to September 1 of the year of the interview, is used instead,  $\hat{\lambda} = 1.459 (.050)$ .

<sup>21</sup>This is also a commonly used procedure in the estimation of Engel curves (see for example Banks et al. (1997), Blundell et al. (1998)) and is used when Engel curves are estimated in section 5.6. If this is reduced to 0.5%,  $\hat{\lambda} = 1.429 (.055)$ , if increased to 2.5%  $\hat{\lambda} = 1.440 (.050)$ .

health are considered in Section 5.4 below. Robustness to the use of alternative econometric estimators is considered in Section 5.5 and comparison made with Engel method estimates using the same data in Section 5.6.

## 5.2 Robustness

Extensive robustness checks in addition to those already mentioned were conducted to investigate the sensitivity of this estimate of  $\lambda$  to various aspects of the specification. On the issue of functional form, the additive form of  $\gamma_1 \ln z_{it} + \gamma_2 d_{it}$  in equation (1) implies that the equivalence scale value,  $\lambda$ , is a constant and, in particular, that it does not vary with income. This would cease to be the case if either  $\gamma_1$  or  $\gamma_2$  were to vary with income. The latter amounts to considering an interaction term between  $\ln z_{it}$  and  $d_{it}$ . If this is added to the model, the hypothesis that its coefficient is zero gives a  $\chi^2(1)$ -statistic of 0.12 ( $p = .732$ ). The hypothesis that the income profiles for pensioner couples and single pensioners are parallel is not rejected.

The possibility that the latent variable is not linear in  $\log(\text{income})$ , i.e. that  $\gamma_1$  varies with income, was considered in two ways. First, a spline with single node at median income (i.e. two straight lines meeting at the median) was used. This gives a coefficient of 0.972 (.054) below the median and 1.091 (.055) above the median. The test of the hypothesis that these coefficients are equal gives a  $\chi^2(1)$ -statistic of 1.94 ( $p = .163$ ). The implied estimates of  $\lambda$  are 1.459 (.059) below the median and 1.400 (.055) above the median. As a second test, the square of  $\log(\text{income})$  is added to the main specification. The test that this has a zero coefficient gives a  $\chi^2(1)$ -statistic of 0.57 ( $p = .450$ ). Thus all these tests support the additive form in  $\ln z_{it}$  and  $d_{it}$  used in equation (1).

The main specification assumes that  $\lambda$  is constant over time. The second and third rows of Table 2 give estimates based on the first 7 waves and the last 7 waves respectively. The hypothesis that the two coefficient vectors are equal is rejected

in a likelihood-ratio test. However the two estimates of  $\lambda$ , 1.454 (.053) in the first period and 1.443 (.073) in the second, are very similar. The panel-robust Wald test of their equality gives a  $\chi^2(1)$ -statistic of 0.02 (p-value = .881). The hypothesis of inter-temporal constancy of  $\lambda$  is supported by the data. Taking a moving 7-year window produces estimates of  $\lambda$  that range from 1.429 to 1.489. Again in both these cases panel-robust Wald tests do not reject the hypothesis of inter-temporal constancy of  $\lambda$ . Several other tests based on interacted time dummies or trends come to the same conclusion.

The results are also robust to the combining of the lower categories, where the data is more sparse. These estimates are shown in the next two rows of Table 2. If those finding it quite difficult and those finding it very difficult are combined to give four categories,  $\hat{\lambda}$  hardly changes: 1.441. If both these are combined with “just about getting by” to give three categories,  $\hat{\lambda}$  is still similar at 1.468.

The ordered probit model assumes a parallel response function at each of the thresholds, i.e. the inverse-normal transformation of the cumulative probabilities differ in their intercepts, but are assumed to have common slope coefficients with respect to the explanatory variables as in equation (5). Diagnostic score tests for threshold heterogeneity in the main specification give cause to doubt this assumption.<sup>22</sup> That for threshold heterogeneity with respect to  $\ln z_{it}$  gives a  $\chi^2(1)$ -statistic of 27.09, that with respect to  $d_{it}$  gives a  $\chi^2(1)$ -statistic of 10.39, both strongly rejecting threshold constancy. The impact of allowing for this potential threshold variation is examined in the context of the three-category model considered in the previous paragraph (since data in the lower categories are rather sparse for estimation of separate  $\beta$ -vectors). The next two rows of Table 2 give binary probit estimates for the probabilities of 4 and 5 versus 1 to 3 and of 5 versus 1 to 4. The former increases the estimate of  $\lambda$  to 1.511, the latter reduces it to

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<sup>22</sup>The score test for threshold heterogeneity proposed by Machin and Stewart (1990) is used.

1.435.<sup>23</sup> The panel-robust Wald test of the equality of  $\lambda$  at these two thresholds has a p-value of 0.10, meaning that equality is not rejected at the 5% level, but the decision is somewhat marginal.<sup>24</sup>

Another issue that needs to be examined is the unbalanced nature of the panel data being used, i.e. the fact that there are missing observations for some individuals in some of the years.<sup>25</sup> Biases may result if the reasons for individuals entering or leaving the sample are correlated with the error term. Endogenous sample selection estimators can be extended to the panel context, but convincing exclusion restrictions are very hard to find. Alternatively a number of less formal approaches have been suggested for linear panel data models and can be extended to address the issue in the current context. The simple tests suggested by Verbeek and Nijman (1992) for the linear model can be adapted. They define a response indicator variable,  $r_{it} = 1$  if  $(y_{it}, x_{it})$  is in the sample and  $r_{it} = 0$  if not, and then test for endogenous selection by adding functions of these to the model under consideration.

Two of their suggestions in particular may be useful here. The first tests whether the error term is correlated with the probability of an individual dropping out in the next period by adding  $r_{it+1}$  to the model. If this is done for the main specification, the panel-robust Wald test gives a p-value of 0.166 and the hypothesis of exogenous selection is not rejected.

A second test uses  $b_i = \prod_{t=1}^T r_{it}$ , a binary indicator equal to 1 if and only if individual  $i$  is observed in all periods (i.e. if individual  $i$  would be a member of the balanced panel subsample). If this variable is added to the main specification, its coefficient is significant (p-value = 0.003), giving evidence of endogenous selection.

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<sup>23</sup>Estimates based on the generalised ordered probit model (which allows separate coefficient vectors at each threshold) are almost identical.

<sup>24</sup>The panel-robust estimator of the variance matrix of the stacked parameter vector is produced in the usual way, giving estimates of the covariances between separate models as well as of those within the same model (White, 1982).

<sup>25</sup>The sample is unbalanced as a result of selection on age and trimming of the household income distribution as well as non-response.

This test is connected to a quasi-Hausman test suggested by Verbeek and Nijman, which is based on the difference between the estimates of  $\beta$  using the full sample and the balanced panel subsample. The variant used here is based on the pooled ordered probit estimates and uses a panel-robust test statistic. Applied to the main specification this gives a  $\chi^2(22)$ -statistic of 42.32 ( $p = .006$ ). However the balanced panel estimate of  $\lambda$  given in Table 2 does not change greatly: from 1.440 to 1.412. The test of equality of the two estimates of  $\lambda$ , i.e. of a lack of selection bias in the estimation of  $\lambda$ , gives a  $\chi^2(1)$ -statistic of 0.14 ( $p = .712$ ).<sup>26</sup>

### 5.3 Home ownership

Assets are likely to be an important influence on the financial well-being of pensioners as well as current income. They provide both potential consumption and additional security. For many pensioners the most important is housing.

There is a clear association between the responses to the question on how they are managing financially and owner-occupation. 47% of owner occupiers say they are “living comfortably” compared with 24% of those living in rented accommodation. At the other end of the distribution, 26% of owner occupiers are “just about getting by” or finding it difficult or very difficult compared with 46% of those living in rented accommodation.

The impact of this on the estimated equivalence scale value is investigated by incorporating an estimate of the imputed rent from owner occupation into an extended definition of income. Imputed rent (or more accurately the net imputed return on housing equity) can be estimated as the income flow from converting the housing equity into an annuity, net of property taxes and mortgage interest. In practice 6% of net housing equity is a commonly-used construction. This is used in this paper, but the sensitivity of the results to the rate used is also

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<sup>26</sup>Estimates using two intermediary samples were also considered. If individuals with only a single observation are excluded,  $\hat{\lambda} = 1.449 (.054)$ . If individuals who enter late or who exit and re-enter are excluded (i.e. only exits without re-entry are allowed),  $\hat{\lambda} = 1.396 (.061)$ . Neither is significantly different from the estimate on the full sample.

examined. Net housing wealth is calculated for home owners as the estimated value of their home less the estimated value of any outstanding mortgage debt. 78% of pension couples and 47% of single pensioners own their own home. Of these, about one in ten have some outstanding mortgage debt. The value of the property is derived mainly from the respondent's expectation of what they would get for their home if sold at that time. There is then some imputation of missing values from other available information. The outstanding mortgage debt is estimated from information on the amount originally borrowed, the year the mortgage started and the years left to run on the mortgage.

Given the small proportion of pensioners that have outstanding mortgage debt, the low share of the property value that the outstanding debt typically represents, and the difficulty of estimating it accurately, a case can also be made for using gross housing wealth. If imputed rent is estimated as 6% of gross housing wealth (converted to a weekly basis) and added to household income, the estimate of  $\lambda$  in the main specification is 1.446 (.055), i.e. little changed.<sup>27</sup> Examining the sensitivity to the rate used, reducing it to 5% and 4% gives estimates of 1.455 (.053) and 1.462 (.051) respectively, while increasing it to 7% and 8% gives estimates of 1.436 (.058) and 1.426 (.060) respectively.

If net housing wealth is used (i.e. net of an estimate of any outstanding mortgage debt), the estimate of  $\lambda$  in the main specification is 1.433 (.054). Using rates from 4% to 8% as above gives estimates of  $\lambda$  that range from 1.414 to 1.449. Overall the inclusion of imputed rent in the definition of income for owner occupiers has little effect on the estimate of  $\lambda$ .

## 5.4 Disability and poor health

Disability and state of health more generally affect the standard of living that can be attained at a given level of income. In addition to the extra costs of living

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<sup>27</sup>If the model is estimated for owner occupiers only, the estimate of  $\lambda$  is 1.393 (.059), again little changed.

associated with disability, responses to questions about an individual's financial position may be influenced by how they are feeling in general terms, which in turn may be influenced by the individuals' state of health. Since the incidence of disability and poor health is higher among pensioners than among the working age population, these factors may be particularly important for this group. The same approach that is used to allow for differences in household size can be used to allow for disability or differences in quality of health. Zaidi and Burchardt (2005) find evidence that the estimated costs of disability derived in this way are substantial and rise with the severity of the disability.

The impact of a number of health variables was examined. The first measure is a binary indicator of disability.<sup>28</sup> This has a highly significant negative effect when added to the main specification (t-ratio = -6.92). Someone who is disabled is estimated to require an income 29% higher to produce the same self-evaluation of their financial position. However the impact on the estimate of  $\lambda$  is small, changing to 1.450 (.052). If the model is estimated solely on those who are not disabled, the estimate of  $\lambda$  is 1.484 (.055), again relatively little changed.

Self-evaluated health status over the last 12 months "compared to people of your own age" (on a 5-point scale) was also used (as 4 dummy variables). The estimated effects here too are large and highly significant. However the impact on the equivalence scale value,  $\lambda$  is negligible – an estimate of 1.443 (.053).<sup>29</sup>

Alternative health controls were constructed using the responses to the General Health Questionnaire (GHQ). This is a battery of questions originally developed as a screening instrument for psychiatric illness. The BHPS self-completion ques-

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<sup>28</sup>There is a change in wording of the question involved between the waves. In waves 1-11 the respondent was asked "are you registered as a disabled person?", while in waves 12-14 the question was "do you consider yourself to be a disabled person?" Using this as a single measure assumes that the impact of the change in question wording is captured by the wave dummies.

<sup>29</sup>Including both these controls and the binary indicator for disability gives an estimate of  $\lambda$  of 1.448 (.053). Adding additional controls for whether the individual has any of the health problems or disabilities from a list and whether they have made use of any of the health and welfare services in the previous year do not alter this estimate of  $\lambda$ .

tionnaire contains a reduced 12-question version of the GHQ. If the Likert scale constructed from the GHQ is added to the main specification, the estimate of  $\lambda$  is 1.460 (.053).<sup>30</sup> Overall, while the estimates indicate that those with adverse health indicators require significantly higher income to generate the same self-evaluated financial position, incorporating them does not alter the estimated equivalence scale relationship between pensioner couples and single pensioners.

## 5.5 Alternative econometric estimators

This section examines the sensitivity of the equivalence scale results to the econometric model and/or estimator used. In the model incorporating unobserved individual-specific heterogeneity (by specifying  $\varepsilon_{it} = c_i + u_{it}$ ) discussed in section 4, the pooled ordered probit estimator provides a consistent estimator of  $\lambda$  without the need for the conditional independence assumption that is required for the consistency of the random effects ordered probit estimator. However if it is the case that  $y_{i1}, \dots, y_{iT}$  are independent conditional on  $(x_i, c_i)$ , then the random effects ordered probit estimator provides an efficiency gain over the pooled ordered probit estimator. Random effects ordered probit estimates are considered next. In the standard (or “uncorrelated”) random effects model  $c_i$  is assumed uncorrelated with  $x_i$ . The results of applying this estimator to the “at least two waves” and “balanced panel” samples, for which pooled ordered probit results were presented in section 5.2, are given in the next block of Table 2.

The coefficient estimates from the pooled ordered probit and random effects ordered probit estimators are not directly comparable due to the different scalings. However the estimates of  $\lambda$  are comparable since the scaling factor gets cancelled out. In the unbalanced panel (but requiring at least two waves) the estimate of  $\lambda$  falls from 1.449 to 1.430 (and its standard error rises). In the balanced panel the

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<sup>30</sup>If the Caseness scale is used, the estimate is almost identical at 1.459 (.053). Using large sets of dummy variables constructed from the GHQ responses in various ways rather than these scales gives estimates of  $\lambda$  that range from 1.460 to 1.463.

estimate of  $\lambda$  rises from 1.412 to 1.485 (again with a rise in standard error).

Both differences are smaller than the standard errors involved.

Correlation between  $c_i$  and  $x_i$  can be allowed by using the Mundlak approach described in section 4. If the  $\bar{x}_i$  are added to the last model as indicated by this approach, the estimate of  $\lambda$  changes very little: from 1.485 to 1.488. Alternatively, following the Chamberlain approach,  $c_i$  can be specified as a function of the entire  $x_i$  vector. Adding these variables to the same model (less those that give rise to perfect collinearity) gives an estimate of  $\lambda$  of 1.490, again very similar to the uncorrelated random effects estimate.

The fixed effects ordered probit estimates for the same balanced panel sample are given in the next row of Table 2. The implied estimate of the equivalence scale value,  $\lambda$ , is 1.479 (.161) and is similar to both the random effects and pooled estimates. The Greene (2004) Monte Carlo results for the binary probit case indicate that the biases in the fixed effects probit and pooled probit estimators (in the correlated individual effects case) are in opposite directions to one another. If this finding is more generally valid, the closeness of estimates for the current analysis would suggest that the magnitudes of these biases may be small and in particular that any bias in the pooled ordered probit estimator from correlated individual-specific effects may be small.

A diagnostic score test for non-normality (Chesher and Irish, 1987) in the main specification gives a  $\chi^2(2)$ -statistic of 41.34, strongly rejecting the assumption of normality. It is therefore important to consider alternative specifications of this distribution (both parametric and semi-parametric) and to examine the robustness of the estimate of  $\lambda$  to these alternatives. I start with alternative parametric models. Ordered logit estimation typically gives similar estimates for ratios of coefficients to ordered probit. For the main specification here the ordered logit estimate of  $\lambda$  is 1.449 (.052). Asymmetric distributions give a greater departure from normality. The ordered Gompertz model given in Table 2 is based on a

distribution which is skewed to the left,  $F(\varepsilon) = \exp(-\exp(\varepsilon))$ . Despite the asymmetry it gives a very similar estimate of  $\lambda$  of 1.447 (.053). If an alternative asymmetric distribution which is skewed to the right, namely the extreme value (or Weibull) distribution,  $F(\varepsilon) = 1 - \exp(-\exp(\varepsilon))$ , is used, the estimate of  $\lambda$  is 1.434 (.056), again very similar to that based on the normal distribution. Thus all four of the parametric distributional specifications produce very similar estimates of  $\lambda$ .

Turning to the results from using the semi-parametric estimators, the SNP estimator approximates the unknown density by a product of the square of a polynomial (of order  $K$ ) and a normal density. The model is estimated for a series of values of  $K$  and the final specification chosen by likelihood-ratio tests between them. When applied to the main specification used in section 5.1, the ordered probit model and model with  $K=3$  are rejected against the  $K=4$  model, while the  $K=4$  model is not rejected against models with higher  $K$ . The SNP estimator with  $K=4$  gives an estimate of  $\lambda$  in the main specification of 1.453 (.052), very similar to that from the ordered probit estimator.

The second semi-parametric estimator used, the polynomial-generalised ordered probit estimator, uses a flexible polynomial transformation to normality, similar to the family of distributions proposed by Ruud (1984). The form proposed by Stewart (2006), in which the derivative of the transformation is specified as the square of a polynomial (of order  $M$ ) to impose the required monotonicity, is used. When applied to the main specification used in section 5.1, the ordered probit model and models with  $M=1$  and 2 are rejected against the  $M=3$  model, while the  $M=3$  model is not rejected against models with higher  $M$ . The polynomial-generalised ordered probit estimator with  $M=3$  also gives an estimate of  $\lambda$  in the main specification of 1.453 (.052), identical to the SNP estimate, and again very similar to that from the ordered probit estimator.

## 5.6 Comparison with Engel method estimates

The Engel method can also be viewed as based on estimation of equation (1). In this case the monotonic function of welfare,  $g(w_{it})$ , is proxied (inversely) by the proportion of income spent on food:

$$s_{it} = \gamma_1^* \ln z_{it} + \gamma_2^* d_{it} + q'_{it} \delta^* + \varepsilon_{it}^* \quad (6)$$

where  $s_{it} = f_{it}/z_{it}$ , with  $f_{it}$  the household expenditure on food. This form of the Engel curve is commonly referred to as the Working-Leser specification. The coefficients  $\gamma_1^*, \gamma_2^*$  are expected to take the opposite signs to  $\gamma_1, \gamma_2$  respectively. The equivalence scale value of interest,  $\lambda$ , is again given by the equivalent of equation (2). As above, the analysis is conditional on household composition.

Although this form of Engel curve has been found to provide insufficient curvature to fit the data for certain commodity groups, it has also been found to provide a good approximation (with linearity in log income not rejected) for other groups such as food expenditure. Banks et al. (1997) find, using FES data for the UK and a variety of alternative parametric and nonparametric estimation techniques, that the Engel curve for food is very close to linear in log income. The kernel regression that they present for food is extremely close to linearity and in the quadratic parametric model the coefficient on the square of log income is insignificantly different from zero. Equation (6) therefore represents a useful starting point for the analysis of food share equations presented here.

BHPS respondents are asked approximately how much the household spends each week on food and groceries. In wave 1 respondents were asked to give an actual amount. In subsequent waves they were asked to select one of 12 intervals shown on a card. With this interval measure divided by the continuous weekly income variable, the data therefore provide lower and upper bounds on food share. If a distributional assumption is made, an interval regression estimator (as in Stewart, 1983) can then be used to estimate (6). Very few of the observations

in the sample used here are in the open-ended categories (just over 1% of the sample). In such circumstances, the Monte Carlo evidence suggests that using interval midpoints and linear regression gives very similar estimates to the use of interval regression with normal errors (Stewart, 1983). This is confirmed for the estimation of (6) by the two methods.

Estimates of  $\gamma_1^*$ ,  $\gamma_2^*$  and  $\lambda$  using the BHPS data described are given in Table 3. As far as I am aware, this is the first study that estimates subjective and Engel equivalence scales on the same panel of households. The first two rows of Table 3 give estimates at the household level using interval regression and midpoint regression. They are almost identical. The next two rows give the results of estimation at the individual level – to enable comparison with the subjective measure estimates. They too are almost identical – both to each other and to the household-level estimates. The estimates of  $\lambda$  are considerably higher than that based on the subjective measure. The individual-level sample is slightly smaller than that in the first row of Table 2. Estimation of the subjective measure model on this sample gives an estimate of  $\lambda$  of 1.444, little changed from Table 2. The panel-robust Wald test (see footnote 24) of the equality of the subjective and Engel estimates of  $\lambda$  gives a  $\chi^2(1)$ -statistic of 20.25 (p-value < 0.0001). Hence the null of equality is strongly rejected. The Engel method estimate of  $\lambda$  is significantly higher than that from the subjective measure model.

[Table 3 near here]

The functional form, and in particular the additive form of  $\gamma_1^* \ln z_{it} + \gamma_2^* d_{it}$  in equation (6) is examined by considering interaction and non-linearity terms. When an interaction term between  $\ln z_{it}$  and  $d_{it}$  is added to the model, its coefficient has a t-ratio of -0.30 (p = 0.763). The parallel income profiles assumption is not rejected. However when the linearity in log income assumption is tested by adding the square of log(income), it is highly significant (t = 12.31) and the lin-

earity hypothesis is strongly rejected. This contrasts with the evidence of Banks et al. (1997) referred to above. Since it implies that  $\lambda$  varies with income, it casts the Engel method of equivalence scale estimation into doubt for this data.

Estimation of equation (6) using various restricted ranges of income suggests that the variation in  $\lambda$  may not be all that large and in all cases considered the Engel estimate of  $\lambda$  is still appreciably bigger than that based on subjective data. For example, if only households with income below the median are used the estimate of  $\lambda$  is 1.715 (.053), while if only households with income above the median are used the estimate of  $\lambda$  is 1.768 (.047). If the upper quartile is excluded, the estimate of  $\lambda$  is 1.680 (.035). If the bottom quartile is excluded, the estimate of  $\lambda$  is 1.797 (.039).

The robustness checks conducted on the subjective model (in Section 5.2) on age restriction, addition of households with other members, extent of trimming, use of deflator, etc. were also conducted on this model. All had very little effect on the estimate of  $\lambda$ . The estimate based on waves 1-7 alone is 1.741 (.039), that based on waves 8-14 is 1.691 (.043). The test of their equality gives a  $\chi^2(1)$ -statistic of 1.18 (p-value = .277). The adjustment of income for owner-occupation and the inclusion of health controls (as in sections 5.3 and 5.4) both have little effect on the estimate of  $\lambda$ .

Excluding those with only a single observation, the midpoint regression estimate of  $\lambda$  is 1.725 (.036). Using an (uncorrelated) random effects estimator on this sample then gives an estimate of  $\lambda$  of 1.767 (.035). Restricting the sample to the balanced panel of those present in all 14 waves gives a midpoint regression estimate of  $\lambda$  of 1.618 (.080) and an (uncorrelated) random effects estimate of 1.690 (.076). Allowing for correlated individual-specific effects by adding Mundlak means or Chamberlain factors to the random effects estimator or by using the fixed effects estimator gives exactly the same estimate of  $\lambda$  at 1.725. Thus the estimate of  $\lambda$  is similar in all these cases and still appreciably larger than that

based on the subjective measure. It should of course be remembered that the inappropriateness of the Engel method implied by the rejection of the linear in log income hypothesis puts a caveat on these subsequent estimates.

## 5.7 Implications for poverty and inequality measures

In this section attention is turned to the implications of the equivalence scale value for certain features of the distribution of equivalised income for pensioners. As pointed out in section 2, the composition of pensioner poverty is potentially sensitive to the choice of equivalence scale value. It was noted there that the use of the pensioner lower quartile for each wave as a poverty threshold corresponds on average to 60% of the overall median, the poverty threshold standardly used by the UK government. In the DWP's HBAI annual reports, the chapter on pensioners also looks at thresholds corresponding to 50% and 70% of the overall median. The data given there indicate that these thresholds correspond in turn to the lowest 13% and 40% of pensioners respectively. The analysis here therefore examines the use of poverty thresholds defined as the lowest 13%, 25% and 40% of pensioner equivalised incomes (for each wave).

The focus here is on the composition of pensioner poverty and in particular on that between pensioner couples and single pensioners. Table 4 presents 3 sets of estimates based on equivalised income constructed using the equivalence scale values of 1.44 (the main estimate in this paper based on subjective evaluations), 1.60 (the ratio of state pension rates) and 1.72 (the estimate based on the Engel method in section 5.6).

[Table 4 near here]

The first block of Table 4 presents figures for the percentage of pensioners in poverty (under each of the thresholds described above) who are single pensioners. For example, for the 25% threshold (corresponding to 60% of the overall median),

the estimated equivalence scale value of 1.44 implies that 66% of pensioners in poverty are single pensioners, compared with 58% when the scale value implied by the ratio of state pension rates is used and 53% when the scale value derived from Engel curve estimation for food expenditure is used. A similar gradient across the equivalence scale values is exhibited when the other two poverty thresholds are examined.

To show this same feature in a slightly different way, the second block of Table 4 presents figures for the implied ratio of single to couple pensioner poverty rates. For the 25% poverty threshold, the equivalence scale value of 1.44 implies that the poverty rate for single pensioners is 2.7 times that for pensioner couples. This compares with 1.9 times when the scale implied by the ratio of state pension rates is used and 1.6 times when the scale value derived from Engel curve estimation for food expenditure is used.

Finally the third block of Table 4 presents different measures of the inequality in equivalised pensioner incomes. All show, as expected, that inequality is higher when the equivalence scale value of 1.44 is used than when that implied by the ratio of state pension rates is used, which in turn is higher than when the scale value derived from Engel curve estimation for food expenditure is used. For example the top decile of pensioner equivalised income is 4.1 times the bottom decile when equivalence scale value of 1.44 is used, compared with 3.9 times when that implied by the ratio of state pension rates is used and 3.8 times when the scale value derived from Engel curve estimation for food expenditure is used.

## 6 Discussion and Conclusions

This paper uses panel data on pensioners' subjective evaluations of their financial positions to estimate the equivalence scale relativity for pensioner couples relative to single pensioners. Pensioner couples are estimated to require a net household income that is 44% higher than comparable single pensioners to reach the same

standard of living.<sup>31</sup>

The estimate of the equivalence scale value,  $\lambda$ , presented is also shown to be robust to variations in a range of aspects of sample and variable definition; to the incorporation of imputed returns on housing equity; to the inclusion of a range of health and disability controls; to the use of uncorrelated random effects, correlated random effects, fixed effects and pooled ordered probit estimators; and to the use of a range of parametric and semi-parametric alternatives to the ordered probit estimator.

This estimate of  $\lambda$  is significantly lower than the equivalence scale value implied by the ratio of state pension rates (1.60), the McClements scale ratio (1.64) and the scale value derived from Engel curve estimation for food expenditure using the same data source (1.72), but not significantly different from the modified OECD scale (1.50).

Comparison with previous estimates in the literature using the same type of subjective data as in this paper require caution, since these studies all use data for different countries and different time periods, and most importantly are not restricted to pensioners. The early study of Dubnof et al. (1981) uses US data for 1972 and gives an estimate of  $\lambda$  of 1.48. Bradbury (1989) uses Australian data for 1983, producing an estimate of  $\lambda$  of 1.25. However both these studies are based on rather small samples and have very large standard errors as a result. The 95% confidence interval for the Bradbury estimate, for example, covers the entire range between 1.0 and 2.0 and beyond. The three more recent studies referred to in section 2 use data from the German Socio-Economic Panel and considerably larger samples. Bellemere et al. (2002) use the 1998 wave and get estimates of  $\lambda$  between 1.33 and 1.45 according to the econometric estimator used. Charlier (2002) uses a panel for 1984-91 and gets estimates of  $\lambda$  over the eight years between 1.31 and 1.49. The main estimate in this paper is therefore not out of line with

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<sup>31</sup>The 95% confidence interval for this figure is from 34% to 54%.

the ranges of estimates in these two papers. Schwarze (2003) uses a panel for 1992-99 and gets an estimate of  $\lambda$  of 1.26, but whether this is significantly less than the main estimate in this paper is hard to judge, since its standard error is not given. Judged overall the main estimate in this paper does not seem out of line with these other estimates, although it may be slightly towards the upper end of the range of estimates. However the caveat on this comparison that none of these previous estimates are specifically for pensioners needs to be reiterated and kept in mind.

Bradbury (1989) points out that equivalence scales may be flattened if respondents evaluate their incomes not just relative to “needs”, but also relative to “social reference groups” and if, in the current context, single pensioner households have different social reference groups to pensioner couple households. Examining the marital status of those in single pensioner households indicates that only 15% are classified as never married (73% are widowed, 12% divorced or separated). It therefore seems less likely that the two groups will have different social reference groups in the current context (particularly once age is controlled for).

A recent Rowntree Foundation report (Bradshaw et al., 2008) investigated the “level of income people think is needed to afford a socially acceptable standard of living in Britain today” combining two budget standards methodologies, based on group discussions among members of the public and on expert advisors. The minimum income standards derived for different demographic groups imply an estimate of  $\lambda$  for pensioner couples relative to single pensioners of 1.42, very similar to that estimated in this paper despite the very different methodologies.

The composition of pensioner poverty is shown in section 5.7 to be sensitive to the choice of equivalence scale value. If the bottom quartile is used as the poverty threshold, (corresponding to 60% of the overall median, the poverty threshold standardly used by the UK government), the estimate of 1.44 implies that 66% of pensioners in poverty are single pensioners, compared with 58% when the scale

value implied by the ratio of state pension rates is used and 53% when the scale value derived from Engel curve estimation for food expenditure is used. To put this another way, using the same poverty threshold, the equivalence scale value of 1.44 implies that the poverty rate for single pensioners is 2.7 times that for pensioner couples, compared to 1.9 times when the scale implied by the ratio of state pension rates is used and 1.6 times when the scale value derived from Engel curve estimation for food expenditure is used.

In conclusion, the estimates in this paper suggest that there may be a case in the UK for raising the (relative) state pension for single pensioners.

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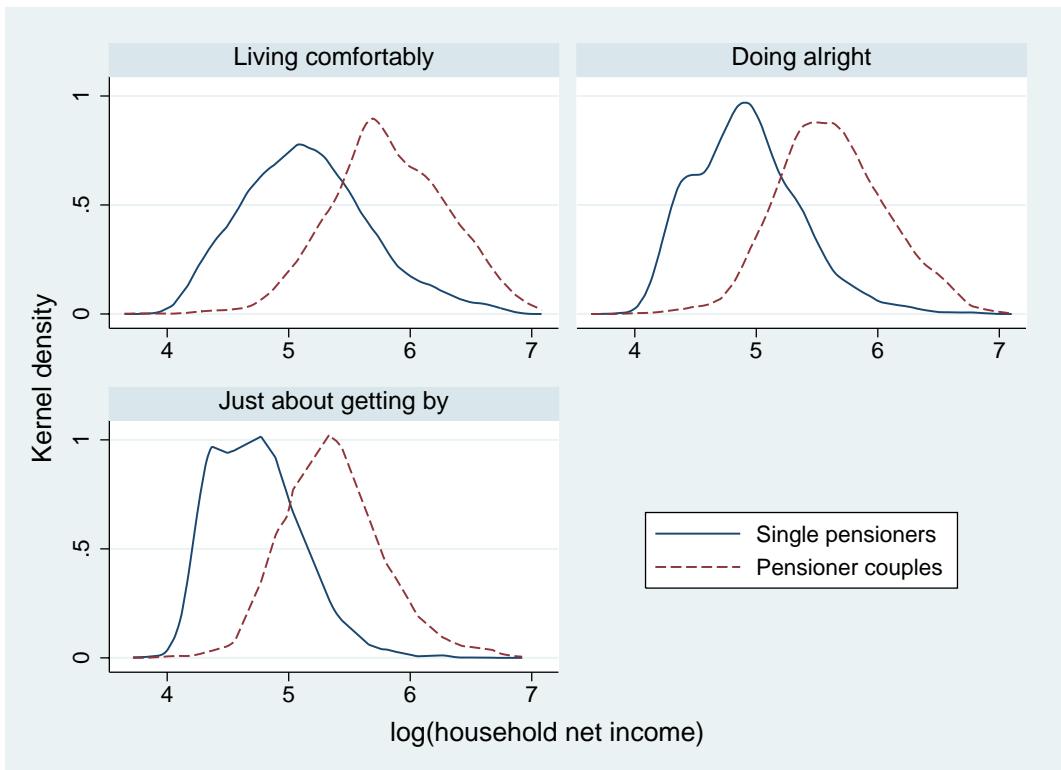


Figure 1: **Kernel densities for the log of household net income**

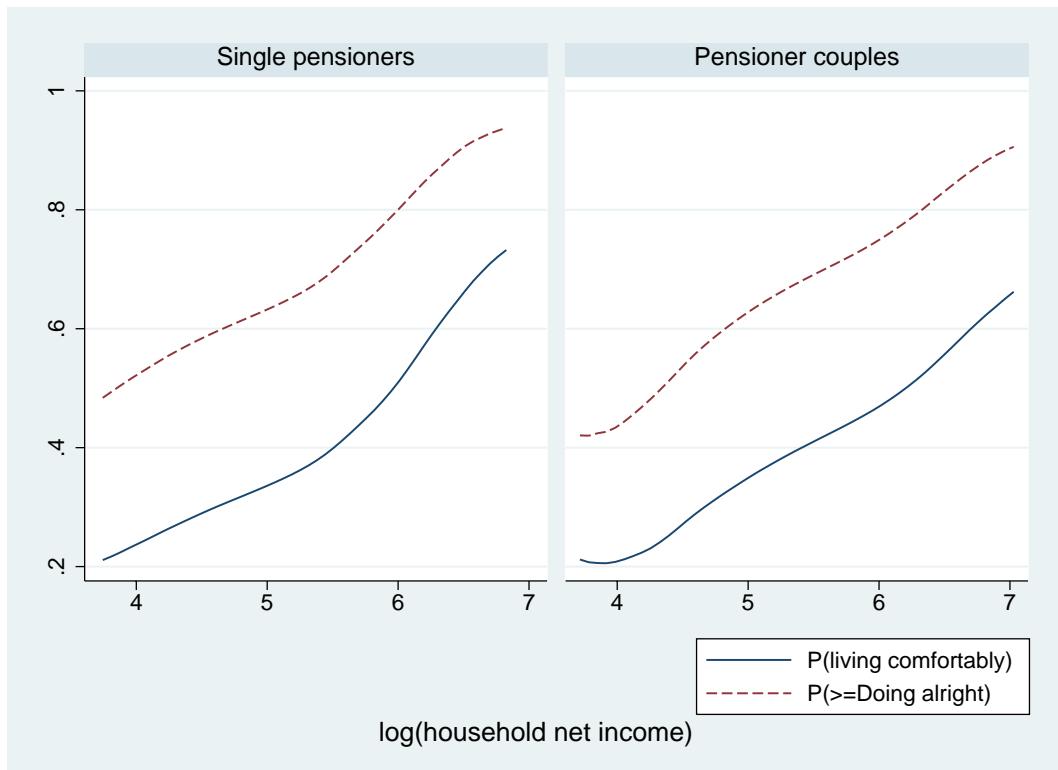


Figure 2: **Kernel regressions for cumulative probabilities**

**Table 1**  
**Descriptive Statistics for How Managing Financially**

	(5)	(4)	(3)	(2)	(1)
All pensioners	39.6	28.0	27.6	3.5	1.3
Pensioner couples	43.2	27.3	25.8	2.8	0.9
Single pensioners	34.7	28.9	29.9	4.6	1.8
Single men	39.9	27.5	26.9	4.1	1.6
Single women	33.1	29.4	30.8	4.7	1.9
Pensioner couples:					
Bottom income quartile	21.9	24.2	46.1	5.8	2.0
2nd. income quartile	33.9	30.4	31.1	3.6	1.0
3rd. income quartile	50.5	29.8	18.1	1.2	0.4
Top income quartile	66.4	24.8	8.0	0.6	0.2
Single pensioners:					
Bottom income quartile	18.2	25.9	42.5	9.5	3.9
2nd. income quartile	25.5	29.3	38.2	5.0	2.0
3rd. income quartile	35.8	33.3	26.7	3.0	1.2
Top income quartile	59.4	27.3	12.3	0.8	0.2

*Notes:*

1. Question asked: "How well would you say you yourself are managing financially these days?"  
 Responses: (5) Living comfortably, (4) Doing alright, (3) Just about getting by,  
 (2) Finding it difficult, (1) Finding it very difficult.
2. Sample size in top row = 21,780.

**Table 2**  
**Ordered Response Model Estimates for How Managing Financially**

	$\gamma_1$	$\gamma_2$	$\lambda$	N. obs	$\log L$
Main specification	1.037 (0.034)	-0.378 (0.042)	1.440 (0.052)	21780	-25083.14
Waves 1-7	1.179 (0.041)	-0.440 (0.048)	1.454 (0.053)	11285	-13323.49
Waves 8-14	0.887 (0.041)	-0.325 (0.050)	1.443 (0.073)	10495	-11641.37
Collapse to 4 cats	1.040 (0.034)	-0.380 (0.042)	1.441 (0.052)	21780	-24480.56
Collapse to 3 cats	1.059 (0.035)	-0.407 (0.044)	1.468 (0.054)	21780	-21582.91
Probit for $\geq 4$	1.121 (0.041)	-0.462 (0.050)	1.511 (0.059)	21780	-11943.74
Probit for = 5	1.013 (0.038)	-0.366 (0.046)	1.435 (0.058)	21780	-13075.40
At least 2 waves	1.036 (0.035)	-0.384 (0.043)	1.449 (0.054)	21379	-24574.82
Balanced panel	1.281 (0.079)	-0.442 (0.087)	1.412 (0.089)	5978	-6393.31
Imputed rent of owner occupiers added to income:					
Gross house value	0.999 (0.035)	-0.368 (0.043)	1.446 (0.055)	20635	-23636.10
Net equity	1.019 (0.034)	-0.367 (0.043)	1.433 (0.054)	20075	-22889.62
Disability/health	1.015 (0.034)	-0.376 (0.041)	1.448 (0.053)	21726	-24750.73
Random effects (uncorrelated):					
At least 2 waves	0.885 (0.030)	-0.316 (0.045)	1.430 (0.067)	21379	-20681.64
Balanced panel	1.042 (0.060)	-0.412 (0.070)	1.485 (0.092)	5978	-5273.98
Correlated individual effects (balanced panel):					
RE + Mundlak	0.820 (0.071)	-0.326 (0.098)	1.488 (0.168)	5978	-5220.93
RE + Chamberlain	0.832 (0.071)	-0.332 (0.099)	1.490 (0.167)	5978	-5173.10
Fixed effects OP	0.865 (0.072)	-0.339 (0.100)	1.479 (0.161)	5978	-4517.29
Gompertz model	0.975 (0.033)	-0.360 (0.039)	1.447 (0.053)	21780	-25263.38
SNP ( $K = 4$ )	1.090 (0.037)	-0.407 (0.045)	1.453 (0.052)	21780	-25029.06
PGOP ( $M = 3$ )	1.110 (0.040)	-0.415 (0.051)	1.453 (0.051)	21780	-25029.46

*Notes:*

1. Dependent variable as defined in Table 1.
2. Estimation by pooled ordered probit unless stated otherwise.
3. Panel-robust standard errors in parentheses.
4. All models also contain controls for age, gender, region of residence and year dummies.

**Table 3**  
**Engel Curve Estimates for Food Share**

	$\gamma_1^*$	$\gamma_2^*$	$\lambda$	N. obs	$\log L$
Main specification: household-level estimates:					
Interval regression	-0.179 (0.004)	0.097 (0.004)	1.716 (0.034)	15113	-24778.85
Midpoint regression	-0.182 (0.004)	0.099 (0.004)	1.724 (0.035)	15113	12340.88
Main specification: individual-level estimates:					
Interval regression	-0.180 (0.003)	0.097 (0.004)	1.714 (0.030)	21451	-36269.70
Midpoint regression	-0.183 (0.004)	0.100 (0.004)	1.723 (0.031)	21451	18257.69
Midpoint regressions on subsamples:					
At least 2 waves	-0.181 (0.004)	0.098 (0.004)	1.725 (0.036)	14817	12327.95
Balanced panel	-0.157 (0.008)	0.075 (0.008)	1.618 (0.080)	3514	3644.19
Random effects (uncorrelated):					
At least 2 waves	-0.213 (0.004)	0.121 (0.005)	1.767 (0.035)	14817	14949.21
Balanced panel	-0.195 (0.008)	0.102 (0.010)	1.690 (0.076)	3514	4406.34
Correlated individual effects (balanced panel):					
RE + Mundlak	-0.207 (0.009)	0.113 (0.016)	1.725 (0.128)	3514	4434.73
RE + Chamberlain	-0.207 (0.009)	0.113 (0.016)	1.725 (0.129)	3514	4462.54
Fixed effects reg.	-0.207 (0.009)	0.113 (0.016)	1.725 (0.128)	3514	4881.10

*Notes:*

1. Panel-robust standard errors in parentheses.
2. Models also contain controls for age, gender, region of residence and year dummies.

**Table 4**  
**Implications for Poverty and Inequality Measures**

	This paper	Equivalence scale value	
		State pension	Engel method
	1.44	1.60	1.72
% of pensioners in poverty who are single pensioners:			
13% poverty threshold	74.1	65.7	58.2
25% poverty threshold	65.8	58.2	53.1
40% poverty threshold	60.0	54.6	50.8
Ratio of single to couple pensioner poverty rates:			
13% poverty threshold	3.98	2.67	1.94
25% poverty threshold	2.68	1.94	1.58
40% poverty threshold	2.09	1.67	1.44
Inequality measures (among pensioners):			
Variance of logs	0.332	0.315	0.307
Coefficient of variation	0.725	0.713	0.707
Ratio of top decile to bottom decile	4.10	3.92	3.81

*Note:*

1. The three poverty thresholds are defined as the lowest 13%, 25% and 40% of pensioner equivalised incomes (for each wave) and correspond respectively to 50%, 60% and 70% of overall median income.