The welfare cost of means-testing:

Pensioner participation in income Support

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ABSTRACT

We estimate parametric and semi-parametric binary choice models of benefit take-up by British pensioners and use a revealed preference argument to infer the cash-equivalent value of disutility arising from stigma or complexity of the claims process. These implicit costs turn out to be relatively small, averaging about £ 3-4 per week across Income Support recipients. Using the Foster-Greer-Thorbecke measure of poverty among pensioners, we find that allowing for implicit claim costs incurred by benefit recipients raises the measured degree of poverty by not more than 13%.

NON-TECHNICAL SUMMARY

Means-testing is an obvious way of focusing welfare spending on those most in need whilst controlling the burden on public finances. The drawback of means-testing is that people who are entitled to receive welfare benefit may not come forward to claim it. There is evidence that this is an important feature of welfare programmes in many countries. Our focus is on older British pensioners (at least 5 years beyond official retirement age), over the period 1997-2002. This is a group with heavy reliance on means-tested benefits, through the Income Support (IS) (also known as Minimum Income Guarantee) programme. During this period, official estimates suggest that around a third of pensioners who appeared to be entitled to IS did not receive it.

Possible reasons include the social stigma that may be associated with welfare receipt, the effort or unpleasantness entailed in the claim process, the difficulty of acquiring and dealing with information about the benefit system, and fears associated with the unpredictability of the claim outcome. These barriers to benefit claiming can be thought of as potential (tangible or intangible) costs and the take-up decision is then seen as weighing expected benefits against these potential costs. Economic models of take-up are often criticised by non-economists as assuming implausible degrees of rationality and knowledge. However, this ignores the fact that it may be efficient to remain in ignorance of the details of welfare programmes if the costs of discovering and understanding their rules are very high and the potential benefits are moderate. It is hard to believe that, if welfare payments were raised to arbitrarily large amounts, a large number of the uninformed would not take some action to become better informed.

There are few specific estimates of the magnitude of claim costs in the research literature and no agreed principle for translating estimates of the takeup-entitlement relation into an estimate of the implied level of underlying claim costs. One of our aims is to develop a way of doing this and to estimate the cash-equivalent values of the tangible and intangible claim costs faced by different individuals. If the process of take-up gives rise to some form of disutility, then it is possible to construct an equivalent weekly cash amount which should be deducted from the observed net income of benefit claimants to give a true income-metric welfare measure. However, in making this income adjustment, we have to take account of the fact that there is self-selection into participation, so that claimants will tend to be those who experience lower than average levels of stigma and other claim costs and non-claimants tend to be those with high levels of claim cost. A further aim is to assess the potential impact of implicit claim costs on the measurement of poverty.

Implicit costs are found to be moderate for most IS recipients, typically around £3-4 per week (or about 13-15% of entitlement) for the average benefit recipient, and consequently the degree of measured poverty among pensioners increases by a modest but non-negligible amount - of up to 13% - when these claim costs are taken into account.

1 Introduction

Welfare programme participation (or, in British parlance, the take-up of means-tested benefits) has been the subject of much applied research. Studies by Moffitt (1983), Fry and Stark (1993), Blundell *et. al.* (1988), Duclos (1995), Bollinger and David (1997) and Keane and Moffitt (1998) are examples of the development of this literature. Means-testing is an obvious way of focusing welfare spending on those most in need whilst controlling the burden on public finances. The drawback of means-testing is that people who are entitled to receive welfare benefit may not come forward to claim it. There is evidence that this is an important feature of many welfare programmes in practice (see Kim and Mergoupis (1997) on AFDC and Food Stamps in the USA, Riphahn (2001) on the Social Assistance Programme in Germany and DWP (2001) on a range of programmes in the UK). Possible reasons include the social stigma that may be associated with welfare receipt and the effort or unpleasantness entailed in the claim process (Moffitt, 1983; Cowell, 1986). Other possible components of claim costs include the costs of information gathering and processing and the implicit risk premium associated with the unpredictability of the claim outcome (see, for example, Halpern and Hausman, 1986). Economic models of take-up are often criticised by non-economists as assuming implausible degrees of rationality and knowledge. This ignores the fact that it may be efficient to remain in ignorance of the details of welfare programmes if the costs of discovering and understanding their rules is very large and the potential benefits are moderate. It is hard to believe that, if welfare payments were raised to arbitrarily large amounts, a large number of the uninformed would not take some action to become better informed.

There are few specific estimates of the magnitude of claim costs in the literature, despite the interest in this issue and findings have been reported in various forms. For example, Duclos (1995) reports expected Supplementary Benefit claim costs of around £3-4 per week for single pensioners and figures as high as £30 per week for some other groups. Moffitt (1983) quotes an elasticity of AFDC participation with respect to entitlement of roughly 0.6, while Blundell, Fry and Walker (1988) report that a 50% increase in Housing Benefit entitlement for the average pensioner household generates a 7 percentage point increase in take-up (an elasticity of roughly 0.2). There is no simple principle established in the literature for translating results on the participation-entitlement elasticity into the implied level of underlying claim costs. One of our aims here is to develop a simple way of doing this.

In this paper, we estimate the cash-equivalent values of the tangible and intangible claim costs faced by different individuals. If the process of welfare participation gives rise to some form of disutility, then it is possible to construct an equivalent weekly cash amount which should be deducted from the observed net income of benefit claimants to give a true income-metric welfare measure. However, in making this income adjustment, we have to take account of the fact that there is self-selection into participation, so that claimants will tend to be those who experience lower than average levels of stigma and other claim costs and non-claimants tend to be those with high levels of claim cost. We have found no published empirical work that takes account of this self-selection in calculating individual-specific estimates of the claim costs incurred by claimants and non-claimants, despite the fact that take-up models are models of self-selection, nor has there been any comparison of alternative ways of making these individual-specific estimates. A further innovation of the paper is to assess the potential impact of implicit claim costs on the measurement of poverty.

Our application is to British pensioners at least 5 years beyond the official retirement age. Apart from the inherent interest in older pensioners as a relatively low-income group, this has the advantage that labour supply is virtually zero, so that labour market complications can be avoided. British pensioners rely heavily on means-tested income from the state, despite the fact that the state pension itself is not means-tested and private pension coverage is high by international standards. In the financial year 2002/2003 32% of British pensioners received some means tested state benefit (Department for Work and Pensions, 2004b). The scope of means-tested pensioner benefits was extended from October 2003 with the introduction of a new means-tested benefit, Pension Credit, to which around 50% of pensioners are expected to be entitled (Department for Work and Pensions (2002a)). Despite the high coverage of means-tested pensioner benefits, they are thought to suffer from a significant degree of non take-up. This is particularly so for Income Support[1] (IS) which provides general income maintenance. Official estimates are that in the financial year 2001/2002 about 33% of pensioners

[1]Income Support for pensioners became known as the Minimum Income Guarantee during the course of our sample period. We retain the older name.

who appeared to be entitled to IS did not receive it (Department for Work and Pensions, 2004a).

2 The state pension-benefit system in Britain

The state pays three main types of benefits to British pensioners: the flat-rate basic state pension, an earnings-related state pension and means-tested benefits. There are also disability-related benefits which are not means-tested. Most pensioners are entitled to the basic state pension earned through paying social security contributions during their working lives, but not all qualify for the full rate of pension. The state earnings related pension scheme (SERPS) was introduced in 1978. Entitlements to SERPS depend on contributions and past earnings. Pensioners retiring in 1998/9 were the first to retire with full SERPS rights. It is possible to opt out of SERPS and contribute to a private pension instead. Recent figures show that the average total state pension payment (basic pension, SERPS and other minor components) in March 2004 was marginally above the full basic state pension (Department for Work and Pensions, 2004c) but below the means-tested benefit level.

2.1 Means-tested benefit rules

During our sample period, there were three main means-tested benefits for pensioners: Income Support (IS) providing general income maintenance; Housing Benefit (HB), giving help with rent; and Council Tax Benefit (CTB) which reduces recipients' liability for local housing-related tax. The rules for calculation of entitlement to HB and CTB mean that pensioners entitled to IS will also be entitled to maximum HB, if they pay rent and CTB if they are liable for Council Tax. People not entitled to IS may be entitled to lower amounts of HB and CTB. Entitlement to each of the three benefits can be calculated independently. In this paper our concern is with IS. There are good reasons to focus specifically on IS. Its rules are independent of receipt of HB and CTB, so there is no bias introduced by analysing IS participation separately from HB and CTB. Moreover, HB has a very high participation rate (over 90%) and CTB entitlements are generally very small, so the decision to claim IS is the critical participation decision.

IS is assessed and paid to pensioner units – single pensioners or pensioner couples - which are our unit of analysis. In our sample period, irrespective of income, entitlement to IS is zero if the pensioner unit's financial holdings are above an upper threshold (£8,000, increasing in 2001/2002 to £12,000). Otherwise it is the difference between a guaranteed minimum (depending on age, disability and whether single or living with a partner) and assessable income (depending on the pensioner unit's income and capital). Assessable income does not include HB and CTB receipts, so entitlements to the three means-tested benefits are independent. For pensioners, the relevant disability-related addition to the guaranteed minimum is the Severe Disability Premium (SDP). Eligibility for the SDP is determined partly by receipt of Attendance Allowance (AA) or the care component of Disability Living Allowance (DLA), which are mutually exclusive, non means-tested disability-related benefits. Sources as well as levels of income therefore affect both assessable income and prescribed amounts. Like HB and CTB, AA and DLA are excluded from assessable income. Actual income from capital is also excluded. Instead a notional income from capital between a lower threshold (£3,000, increasing to £6,000 in 2001/2002) and the upper threshold is assumed at the rate of £1 a week for each £250 or part of £250 of capital between the two limits. In addition to the change in the capital limits, there was a substantial real-terms increase in the level of the income guarantee; these changes provide an important source of identifying variation. The main benefit rates prevailing over the sample period are set out in Appendix Table A1.

2.2 The claims process

To receive Income Support, pensioners must submit a claim to the Department for Work and Pensions (DWP). For most of the period we are concerned with, this entailed completing and taking a 40-page form to their local social security office[2]. Details of all sources of income, savings and relevant personal characteristics have to be provided. Attached to the IS form are supplementary forms covering HB and CTB. Consequently, applications for IS are almost always accompanied by applications for CTB and (for renters)

[2]In October 2001 the claim form was reduced to 10 pages although there was an increase in the number of pages of notes to be read in conjunction with the form.

HB, and IS is almost never received in isolation.[3]

If found to be entitled to IS, payment is often made with the state pension so the pensioner does not need to attend a social security office every week. In principle changes in circumstances which might affect entitlement are supposed to be reported immediately so that payment can be adjusted accordingly. In practice re-assessments are less frequent. The process of claiming the new Pension Credit, which subsumed IS in 2003, is intended to be less stigmatising, with less frequent re-assessments and more of the onus for initiating claims placed on DWP. However in the period considered here, few pensioners would have expected the experience of claiming IS to be enjoyable or hassle-free.

3 Take-up: evidence from the 1997-2002 Family Resources Survey

3.1 The Family Resources Survey

The Family Resources Survey (FRS) is a continuous cross-sectional survey of British households carried out on behalf of the DWP during April 1997 to March 2002. In principle, the FRS gives all information necessary to assess each FRS pensioner unit's entitlement to IS and establish whether they are receiving IS. We have applied the following process of error detection and correction before using the data (and before making the sample deletions listed below). The first step was to reverse data edits and imputations made by the DWP, affecting benefit receipts, private pension income and capital holdings, because we detected some inconsistencies in edits to benefit data and because some of the imputation procedures (such as substitution of sample means for missing values) are inappropriate for our purposes. The next stage involved detecting inconsistencies in benefit data and reconciling them where possible. Potential errors in recorded receipts of social security benefits are generally easier to identify than errors in other sources of income or in capital because specified benefit rates and eligibility rules allow consistency checks to be made. Missing values for benefit receipt were imputed

[3] In our sample period only around 1.36% of IS recipients received neither HB nor CTB. Only 0.26% of those receiving IS+HB did not also receive CTB; among renters, only 1.08% of those receiving IS+CTB did not also receive HB.

where a correct value could be identified unambiguously. For example, some pensioners in the FRS are able to supply a breakdown of their state pension payments which helps disentangle different benefits received as one combined payment. In other cases it is clear that a payment of IS is included in their pension payment and there is double counting if a separate amount of IS is also recorded. Where it was not possible to correct an inconsistency or to impute a missing value on any reliable basis, the value was left missing. This was true for all missing values for private pension and capital holdings where there is no reliable way to impute an individual-specific value. Full details of this data cleaning process can be found in Hancock and Barker (2005).

Two different versions of the dataset are used, differing in the entitlement measure used[4]. Sample 1 uses simulated entitlement for all households and makes no use of recorded benefit amounts, beyond the receipt/non-receipt distinction; for entitled households, sample 2 substitutes recorded benefit where available, for simulated entitlement, provided it does not exceed the guaranteed minimum for the benefit unit and provided the respondent consulted some form of IS documentation when answering the survey question[5]. We focus on older pensioners, defined as single people at least 5 years over state retirement age (60 for women and 65 for men) or couples where both partners are at least 5 years above retirement age. There are several reasons for this: they are a group with a high poverty rate; they have very little labour market involvement to complicate the welfare participation issue; and having been retired for a relatively long time, their adjustment to post-retirement circumstances is likely to be complete[6]. This contrasts with the dynamic modelling issues faced by Blank and Ruggles (1996) and Anderson and Meyer (1997). Note that there is currently no UK longitudinal

[4]An earlier version of this paper used both the pre-cleaned and the cleaned datasets. Data cleaning made relatively little difference to the results.

[5]Note that FRS respondents are encouraged by interviewers to consult bank records, payment books, etc, where possible to give accurate figures. We also tried an additional sample which substituted recorded IS, where available, for simulated entitlement provided it did not exceed the guaranteed minimum for the benefit unit. In such cases and in cases where IS receipt was missing simulated entitlement was used instead. The results are very similar to sample 2 reported in this paper.

[6]Note the insignificant role of age in our preferred semiparametric estimates reported below. Small numbers of FRS respondents (around 0.8% of the sample) had claims pending. We include these cases and treat them as recipients. There is no detectable difference in the results if they are dropped from the sample.

survey that is adequate for the purposes of simulating means-tested welfare entitlements.[7] This means that we have no history of the evolution of individual welfare entitlements over time, and this would preclude any convincing attempt to model the dynamics of participation behaviour.

The subsamples used for our analysis contain 4,003 (sample 1) or 4,129 (sample 2) after deleting households which: were not entitled (16,251); contained multiple benefit units (3,708); were still repaying a mortgage (230); received allowances from an absent spouse (2); had employment or self-employment income (36); did not respond to survey questions on a core variable such as recorded IS receipt, pension or non-assessable income (2,503 and 2,215); or which gave rise to other miscellaneous data-quality concerns (40 and 202). These deletions are less serious than they might at first appear. Most are simple exclusions of pensioners known to be non-entitled, for whom participation is not an issue. Multi-unit households are excluded because of the difficulty of simulating their entitlement. We exclude the tiny number of entitled earners, for whom take-up is complicated by labour supply, the few mortgagors and tiny number of beneficiaries from an absent spouse, because of the measurement problems associated with the calculation of mortgage interest and assessable income, respectively. The most serious of the deletions is likely to be the cases lost through item non-response, which we assume to be ignorable. Given the careful data cleaning and sample selection, we believe that the potentially serious problem of measurement error has been avoided as far as possible in the samples used for analysis. All data on benefits and income are adjusted to 2002 prices using the official Consumer Price Index.

3.2 Take-up rates

Table 1 shows estimated IS take-up rates for the two samples. These rates are largely unaffected by the choice of sample. There is some variation by category of pensioner, but the typical rate of non-participation is roughly one third (36.87% and 35.75% in samples 1 and 2 respectively). Although not directly comparable, these are close to the official estimate of 33% reported

[7] The widely-used British Household Panel Survey does not collect annual data on financial assets, which are necessary to calculate entitlements.

by DWP for 2001/2002. Single females appear to have higher take-up rates than single males and couples. Take-up rates are also higher in both samples for pensioners in the younger age groups. Pensioners who left full-time education after the age of fourteen have lower take-up rates. Take-up rates are higher for pensioners with registered disabilities but the is no significant difference for those in receipt of disability benefit. Take-up varies considerably with housing tenure, renters having much higher rates than home owners. Pensioners in the 2001/02 survey who became entitled for the first time after the reforms have very low take-up rates in both samples. This is partly due to their typically small levels of entitlement, but may also involve some lag in adjustment to the new benefit rules. We pursue this further in the applied work reported below.

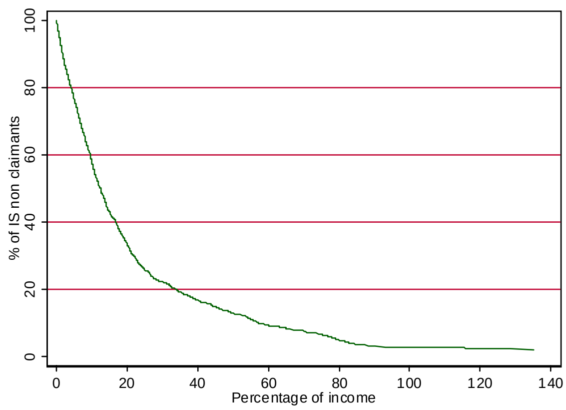
Table 1 Percentage take-up rates by demographic groups (1997/8-2001/02 FRS estimation sample)

	Take-up rate (std. error)	
	Sample 1	Sample 2
Sample size	4003	4129
Full sample	63.13 (0.76)	64.25 (0.75)
Single male	56.24 (2.08)	57.14 (2.05)
Single female	66.83 (0.86)	67.95 (0.84)
Couple	47.60 (2.33)	48.72 (2.31)
Head under 70	77.70 (2.02)	77.86 (2.00)
Head 70-79	64.00 (1.14)	65.11 (1.11)
Head 80-89	59.63 (1.22)	60.86 (1.19)
Head 90+	52.55 (3.57)	56.13 (3.41)
Education $<$ 14	66.14 (3.44)	67.01 (3.38)
Education equal to 14	64.31 (0.89)	65.52 (0.87)
Education $>$ 14	58.69 (1.64)	59.54 (1.62)
Not receiving disability benefit	63.45 (0.91)	63.45 (0.91)
Receiving disability benefit	62.37 (1.40)	65.96 (1.30)
Not registered disabled	62.29 (0.83)	63.06 (0.82)
Registered disabled	67.82 (1.89)	70.48 (1.77)
Owner occupier	46.81 (1.47)	48.30 (1.46)
Renter	70.27 (0.87)	71.18 (0.85)
Newly entitled 01/02	38.19 (2.44)	38.53 (2.59)
Not newly entitled 01/02	65.88 (0.79)	66.66 (0.77)

Let p be a given percentage of original pre-IS net income and let $S(p)$ be the proportion of IS non-claimants who are entitled to an amount of

IS in excess of the proportion of their income. In formal terms, $S(p) = H(b/p)$, where $H(.)$ is the distribution function of the income distribution among entitled non-claimants, b is entitlement and p is some positive income fraction. Figure 1 plots $S(p)$. Even though many pensioners who do not take up IS are entitled to small benefits, over half of the non-claimants could increase their income by at least 10% and more than 25% of non-claimants could increase their income by at least 20% if they were to take-up their entitlements. Thus non-participation in the IS programme has important consequences for a large minority of potential recipients.

Figure 1 Distribution of IS non-claimants by size of unclaimed benefit as a proportion of original income



4 The compensating variation approach

Our analysis is based on the idea that individuals will claim the benefit to which they are entitled whenever they see it as being in their best interests, after allowing for all costs associated with benefit claim and receipt. These claim costs can be financial (such as the cost of travel to the social security office), tangible but non-financial (for example the time or physical difficulty involved), social (for example social stigma) or psychological (such as feelings of inadequacy or shame induced by dependency). Lack of information can also be viewed as a claim cost equivalent, in a specific sense outlined below, to the cost of ignorance.

Note that our notion of claim costs is broad enough to encompass a wide range of factors. Someone who suffers difficulty in coping with the process of claiming benefit because of physical or mental impairment is seen as suffering from high claim costs. Those with access to external assistance from family, neighbours or other carers are likely to find it easier to make a claim than similar people with no such support. Thus claim costs depend on personal characteristics and circumstances as well as factors like the design of applications procedures. It is therefore very important to allow for wide variations in claim costs across benefit units.

4.1 The basic take-up model

Consider first the simplest case of static choice under certainty. Let the long-term welfare of the benefit unit be represented by a utility function $U_0(Y;\mathbf{X},V)$, where Y is net income in the absence of means-tested benefit, $\mathbf{X}$ is a vector of observable characteristics and V represents unobservable characteristics which vary randomly across benefit units. When means-tested benefits are claimed, there is a possible shift in welfare represented by a transformed utility function $U_1(Y+B;\mathbf{X},V)$, where B is the additional benefit income. The shift from U_0 to U_1 is induced by some form of claim costs. Under the assumption of strict rationality, the condition for take-up to occur is:

$$U_1(Y+B;\mathbf{X},V) > U_0(Y;\mathbf{X},V) \tag{1}$$

Since utility is monotonic and continuous in income, this can be rewritten:

$$B > U_1^{-1}\left(U_0(Y;\mathbf{X},V);\mathbf{X},V\right) - Y \tag{2}$$

where the function $U_1^{-1}\left(\,.\,;\mathbf{X},V\right)$ is U_1 inverted with respect to its first argument. Note that, if the functions U_0 and U_1 are identical, $U_1^{-1}\left(U_0;\mathbf{X},V\right)-Y$ is equal to 0 and benefit is claimed whenever the entitlement is strictly positive. When $U_1^{-1}\left(U_0;\mathbf{X},V\right)-Y$ is positive, it is the compensating variation: the cash equivalent of any barriers acting as a disincentive to take-up.

4.2 The model as reduced form

It is important to realise that the model outlined above is applicable as a reduced form in a much wider range of cases. In practice, take-up behaviour may involve non-stationarity, uncertainty and perceptions that change over time in response to new information. There is no possibility of a convincing structural model for the UK, since there are no longitudinal datasets which are adequate for benefit simulation and no direct observation of information and perceptions. Our approach is to regard the binary choice model as a reduced form construct, which permits the derivation of a compensating variation. A first example emphasises that expressing claim costs as an equivalent annual amount does not entail an assumption that they are actually incurred in that form. For example, suppose there is an up-front 'hassle' involved in the initial claim, C_0, followed by a lower annual renewal hassle, C_1. This stream of costs can be expressed as an equivalent annual amount, in much the same way that a capital sum can be annuitised. When receiving benefit, lifetime discounted utility is $U(Y+B,\mathbf{X},V,C_0)+\sum_{t=1}^{T}\rho^t U(Y+B,\mathbf{X},V,C_1)$ $=U(Y+B,\mathbf{X},V,C_0)+[\rho(1-\rho^T)/(1-\rho)]U(Y+B,\mathbf{X},V,C_1)$, where we have assumed static circumstances and a known lifetime T. Call this function $U_1(Y+B,\mathbf{X},V)$. Then condition (2) applies. The only difference between this case and conventional annuitisation is the lack of a secondary market, implying that the annuitisation process depends on the subjective discount factor, ρ, rather than a market rate.

A second example demonstrates the use of the model as an approximation in a case with uncertainty, where information is acquired sequentially, with updating of perceptions. Suppose the pensioner initially has perceptions of entitlement represented by a prior distribution $f_0(\widetilde{B}|B,\mathbf{X})$, where $\widetilde{B}$ is perceived entitlement and B is actual entitlement. The individual's post-retirement circumstances persist for a sequence of periods $1...t$, during which a stochastic flow of new information $I_1...I_t$ is received. With Bayesian updating, perceptions at t are represented by a posterior distribution $f_t(\widetilde{B}|B,\mathbf{X},I_1...I_t)=f_0(\widetilde{B}|B,\mathbf{X})l(\widetilde{B};I_1...I_t|B,\mathbf{X})$ where $l(.)$ is the likelihood reflecting the information acquired and the individual's understanding of the relationship between that information and true entitlement. The dependence of the likelihood on $B,\mathbf{X}$ reflects between-individual variations, including differences in access to information. After t periods, the expected utility of claiming is $E_tU_1(Y+\widetilde{B},\mathbf{X},V)$, where E_t is the expectation with respect to the distribution $f_t(.)$. Thus, at time t, take-up will be observed if B is positive and $E_t U_1(Y + \widetilde{B}, \mathbf{X}, V) > U_0(Y, \mathbf{X}, V)$. The left hand side of this inequality has the general form $\overline{U}_1(Y, B, \mathbf{X}, V, I_1...I_t)$. Under the reasonable assumption that increasing true entitlement shifts the posterior distribution $f_t(\widetilde{B}|B, \mathbf{X}, I_1...I_t)$ rightwards, $\overline{U}_1$ will be increasing in B and the analogue of the take-up condition (2) is $B > \overline{U}_1^{-1}(U_0, Y, \mathbf{X}, V, I_1...I_t)$. In this expression, $V, I_1...I_t$ are unobservable stochastic terms. Since different individuals will have different length sequences of information, elapsed time (represented mainly by age) will appear in the mean of $\overline{U}_1^{-1}(U_0, Y, \mathbf{X}, V, I_1...I_t)$ conditional on $Y, \mathbf{X}$, even if preferences do not evolve with age. In this case, the compensating variation will reflect the cost of risk and imperfect information in addition to stigma and hassle costs.

4.3 The econometric specification

Empirically, the best fit has been obtained by working with the logarithm of benefit entitlement (see also Blundell *et. al*, 1988). We thus approximate the log of the right-hand side of (2) directly by a linear stochastic function $\mathbf{Z}\boldsymbol{\alpha} + V$, where $\mathbf{Z}$ is a vector of variables constructed from $(Y, \mathbf{X})$, rather than using explicit specifications for U_0 and U_1. The stochastic term V now represents the effect of a combination of unobservable preference parameters and information available to the individual but unobserved by the analyst. The condition for take-up is:[8]

$$\ln B > \mathbf{Z}\boldsymbol{\alpha} + V \tag{3}$$

The conditional take-up probability is then:

$$\begin{aligned} \Pr(\text{take-up}|B, \mathbf{Z}) &= \Pr(V < \ln B - \mathbf{Z}\boldsymbol{\alpha}) \\ &= F\left(\frac{\ln B - \mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right) \end{aligned} \tag{4}$$

where σ^2 =var(V) and $F(.)$ is the distribution function of the random variable V/σ. The probability (4) amounts to a standard binary response model of

[8]This approximation is exact if U_1 can be written as $U_0(Y + B - e^{\mathbf{Z}\boldsymbol{\alpha}+V})$. Following the same approach, if we assume that subjective claim costs are proportional to the amount received, then $U_1 = U_0(Y + B - Be^{\mathbf{Z}\alpha+V})$ and the take-up condition is $\mathbf{Z}\boldsymbol{\alpha} + V < 0$, so that take-up does not depend on the scale of entitlement. This is clearly rejected empirically.

discrete choice, using $\ln B$ and $\mathbf{Z}$ as explanatory variables. Note that B is directly observable, from application of the IS rules to the data. There are no parameters requiring estimation in the function $B(Y, \mathbf{X})$. In the model (4), the coefficients of $\ln B$ and $\mathbf{Z}$ are $1/\sigma$ and $-\boldsymbol{\alpha}/\sigma$ respectively, so that $\boldsymbol{\alpha}$ can be estimated as minus their ratio. Given $\boldsymbol{\alpha}$, a conditional distribution of claim costs $C = \exp(\mathbf{Z}\boldsymbol{\alpha} + V)$ can be constructed for each individual benefit recipient. All that is required is a specific form for the function F.

Participation models raise many endogeneity issues. We are modelling participation conditional on entitlement and original income. These variables might be endogenous in the sense that people who know themselves to suffer particularly from stigma (implying negative V) will take steps to accumulate relatively high pension entitlements and other assets, which will in turn raise post-retirement income and reduce IS entitlement (implying low B). This positive correlation between V and B might be thought to generate an upward bias in the coefficient $(1/\sigma)$. However, the problem is not this simple. Under endogeneity, income Y is negatively correlated with V; it is also negatively correlated with B. Under these circumstances, the biases in the coefficients of $\ln B$ and Y cannot be signed *a priori* since each has two components of opposite sign. Moreover, the potential biases are moderated by the fact that the model is fitted only to those with strictly positive entitlement. Anyone whose fear of stigma is sufficiently large to increase their pension income or assets above the critical level will not be included in the estimation sample and will make a smaller bias contribution than would be the case under exogenous sample selection. A further consideration is that the important decisions governing pension income and asset accumulation were typically made many years earlier than the IS participation decision and often involved little real choice - the basic state pension scheme was all that was available to most of this cohort of poorer pensions. For these reasons, we are confident that endogeneity biases in our participation model are likely to be small. Moreover, convincing analysis of a model endogenising pensions and assets (and, potentially, housing and education also) would require long-horizon longitudinal data that does not currently exist for the UK.

4.4 Identification

In general terms, the model is of the form:

$$\Pr(\text{take-up}|Y,\mathbf{X}) = G(B(Y,\mathbf{X}), Y, \mathbf{X}) \tag{5}$$

where G is a function with range $[0,1]$ and $B(Y,\mathbf{X})$ represents the rules of the IS programme. If all the variables in $(Y,\mathbf{X})$ can appear indirectly through a simple benefit rule $B(Y,\mathbf{X})$ and also directly in their own right, then it is clear that the model is nonparametrically unidentified despite the fact that $B(.)$ is a known function. However, the changes in the rules of the IS system in 2000 and 2001 breaks the exact functional relationship between B and $(\mathbf{X}, Y)$ in the sample and generates independent identifying variation. Moreover, there are further restrictions that help to resolve this identification problem. We are usually content to make a smoothness assumption about the direct effect on behaviour of personal characteristics such as age, income and wealth. There are, several discontinuities and kinks built into the IS rules: (i) discontinuities in the guaranteed minimum with respect to age (at 75 and 80); (ii) several discontinuities in the guaranteed minimum with respect to the amount of disability benefit; and (iii) a kink in the definition of notional income with respect to capital (at £3,000, rising to £6,000 in 2001/2). A smoothness assumption on the direct impact of age, capital and the disability benefit element of income will theoretically suffice to ensure identification, provided the minimum acceptable degree of smoothness can be imposed appropriately. Exclusion restrictions can also be used to identify the model. If one or more of the variables determining B can be excluded *a priori* from the model, then the separate impacts of B and $(Y,\mathbf{X})$ can be distinguished empirically. Our final specification embodies several such restrictions. Some of these are data-driven, but we have assumed *a priori* that financial capital has a direct effect on take-up behaviour only through the contribution of observed investment returns to net income, which does not affect the calculation of IS entitlement.

4.5 Implicit claim costs

Our aim is to construct estimates of implicit claim costs: the compensating variation required to offset stigma and other barriers to participation. Once $F(.)$, σ and $\boldsymbol{\alpha}$ are known, estimates of individual claim costs can be constructed in various ways. It is not appropriate to use the unconditional mean $\mathbf{Z}_i\boldsymbol{\alpha}$ as most other researchers have done, since this does not make use of the information we have about the actual take-up decision of unit i. Instead we should condition the prediction of claim costs for claimants on the take-up event $\mathbf{Z}\boldsymbol{\alpha} + V < \ln B$. For non-claimants costs can be estimated by conditioning on the event $\mathbf{Z}\boldsymbol{\alpha} + V \geq \ln B$.

A natural approach is to use a conditional expectation. For the ith IS recipient:

$$\begin{aligned} \widehat{C}_{1i} &= E(\exp(\mathbf{Z}_i\boldsymbol{\alpha} + V_i)|V_i < \ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}) \\ &= e^{\mathbf{Z}_i\boldsymbol{\alpha}} \left[\int_{-\infty}^{\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}} e^V dF(V) \right] \left[\int_{-\infty}^{\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}} dF(V) \right]^{-1} \end{aligned} \tag{6}$$

In the special probit case where F is the standard normal distribution function, this yields the following expression (Aitchison and Brown 1957, page 87):

$$\widehat{C}_{1i} = \exp\left(\mathbf{Z}_i\boldsymbol{\alpha} + \frac{\boldsymbol{\sigma}^2}{2}\right) \Phi\left(\frac{\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha} - \boldsymbol{\sigma}^2}{\boldsymbol{\sigma}}\right) \Big/ \Phi\left(\frac{\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}}{\boldsymbol{\sigma}}\right) \tag{7}$$

where $\Phi(.)$ is the standard normal distribution function. For a non-claimant:

$$\begin{aligned} \widehat{C}_{1i} &= E(\exp(\mathbf{Z}_i\boldsymbol{\alpha} + V_i)|V_i \geq \ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}) \\ &= \exp\left(\mathbf{Z}_i\boldsymbol{\alpha} + \frac{\boldsymbol{\sigma}^2}{2}\right) \left[1 - \Phi\left(\frac{\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha} - \boldsymbol{\sigma}^2}{\boldsymbol{\sigma}}\right)\right] \Big/ \left[1 - \Phi\left(\frac{\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}}{\boldsymbol{\sigma}}\right)\right] \end{aligned} \tag{8}$$

An alternative is to use a conditional median estimate, $\widehat{C}_2$, which satisfies $\Pr\left(\mathbf{Z}_i\boldsymbol{\alpha} + V_i < \ln \widehat{C}_{2i} \,|V_i < \ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}\right) = 0.5$. Using Bayes' rule for claimants:

$$F\left(\frac{\ln \widehat{C}_{2i} - \mathbf{Z}_i\boldsymbol{\alpha}}{\sigma}\right) \Big/ F\left(\frac{\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}}{\sigma}\right) = 0.5 \tag{9}$$

and thus:

$$\widehat{C}_{2i} = \exp\left\{\sigma\left[\mathbf{Z}_i(\boldsymbol{\alpha}/\sigma) + F^{-1}\left(\frac{1}{2}F\left(\frac{\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}}{\sigma}\right)\right)\right]\right\} \tag{10}$$

For non-claimants, the condition $\Pr(\ln B_i < \mathbf{Z}_i\boldsymbol{\alpha}+V_i < \ln \widehat{C}_{2i} \,|\mathbf{Z}_i\boldsymbol{\alpha} + V_i > \ln B_i) = 0.5$ gives:

$$\widehat{C}_{2i} = \exp\left\{\sigma\left[\mathbf{Z}_i(\boldsymbol{\alpha}/\sigma) + F^{-1}\left(\frac{1}{2}\left[1 + F\left(\frac{\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}}{\sigma}\right)\right]\right)\right]\right\} \tag{11}$$

Note that $\widehat{C}_{1i}$ and $\widehat{C}_{2i}$ always lie below the unconditional mean and median of $\exp(\mathbf{Z}_i\boldsymbol{\alpha} + V)$ for participants and above for non-participants. Claimants will, on average, tend to be those who suffer lower than average levels of claim costs and conversely for non-claimants. The relationship between implicit claim costs and the coefficient of $\ln B_i$ is important. As $\sigma \to \infty$, the impact of entitlement on take-up vanishes. If we adjust $\boldsymbol{\alpha}$ so as to keep the take-up probability constant at some value P, then $\lim_{\sigma\to\infty} \mathbf{Z}_i(\boldsymbol{\alpha}/\sigma) = -F^{-1}(P)$. Consider the median (10). Since $F^{-1}(\frac{1}{2}P) - F^{-1}(P) < 0$, $\lim_{\sigma\to\infty} \widehat{C}_2 = 0$. This occurs because the leftward shift in the median induced by the truncation condition $C_i < B_i$ is greater, the larger is σ. Conversely, $\lim_{\sigma\to\infty} \widehat{C}_{2i} = +\infty$ for non-claimants: as we increase σ, $\mathbf{Z}_i\boldsymbol{\alpha}$ must increase towards $\ln B_i$ in order to keep the take-up probability constant. Thus the entitlement coefficient is critical in this type of model. A small value will imply modest implicit claim costs for those who do take-up the benefit, but very much larger costs for those who do not. A large coefficient implies large claim costs for claimants and a weaker distinction between claimants and non-claimants.

5 Estimates

5.1 The binary take-up model

We apply two different estimators of the binary take-up model. One is the familiar probit model, based on the assumption that the distribution function $F(.)$ is standard normal. The second is the semi-parametric estimator of Klein and Spady (1993) which, in its simplest form, maximises the following quasi-log-likelihood:

$$\max_{\boldsymbol{\alpha}} \ln L(\boldsymbol{\alpha}) = \sum_{i=1}^{n} \left\{\tau_i \ln\left(\widehat{F}(\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha})\right) + (1-\tau_i)\ln\left(1 - \widehat{F}(\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha})\right)\right\} \tag{12}$$

where: summations are over the set of n pensioner households for whom $B_i > 0$; τ_i is the dependent variable, equal to 1 for IS participation and 0 for non-participation; and $\widehat{F}(.)$ is a nonparametric kernel estimate of the regression function of y_i on $\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}$. We use the Gaussian kernel:

$$\widehat{F}(\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}) = \frac{\sum_{j\neq i} \phi\left(h^{-1}[(\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}) - (\ln B_j - \mathbf{Z}_j\boldsymbol{\alpha})]\right) y_j}{\sum_{j\neq i} \phi\left(h^{-1}[(\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}) - (\ln B_j - \mathbf{Z}_j\boldsymbol{\alpha})]\right)} \tag{13}$$

where $\phi(.)$ is the standard normal density function. Note that $\widehat{F}$ is not normalised to have zero mean and unit variance. Scale and location are normalised by fixing the coefficient of $\ln B_i$ at unity and excluding the intercept term from the linear form $\mathbf{Z}\boldsymbol{\alpha}$. This choice of normalisation does not affect the construction of implicit cost estimates. We experimented with a variety of fixed and adaptive bandwidths (the latter using the Breiman *et. al.* (1977) method). The results were remarkably insensitive to the particular choice used. The results reported below are based on a fixed bandwidth h equal to 0.6.

Table 2 gives estimates of the stigma/claim cost coefficients $\boldsymbol{\alpha}$. The variables appearing in the model are defined and summarised in Appendix Tables A2 and A3. For the probit model the estimates are calculated as minus the coefficients of the relevant variables divided by the coefficient of $\ln B_i$ (full coefficients are given in Appendix Table A3). The estimates are the outcome of an extensive process of specification search. The chosen form is superior (in likelihood terms) to other models with alternative functional forms for income and entitlement. There has been some attention paid by sociologists to neighbourhood influences on welfare participation behaviour, with the conclusion that high local rates of poverty, welfare dependency and density of population lead to higher rates of take-up, because of the lesser impact of social stigma and better local information and support, reducing claim costs (Hirschl and Rank, 1999). We are only able to match survey respondents to large regions (at Standard Region level) rather than neighbourhoods and there are, consequently, no locational effects detectable. We are also able to accept our specification against models with fuller demographic structure and a more general specification involving the ages and education levels of both members for 2-person households. Annual dummy variables were also included and found to be insignificant with χ^2 statistics of 3.33 and 5.50 for samples 1 and 2 respectively and a 5% critical value of 9.488. To guard against pre-test bias, we have worked from 'general' models to 'specific' ones, using a conservative criterion, retaining explanatory variables with asymptotic t-ratios in excess of 1.0.[9] Besides log entitlement, the main factors generating high claim costs emerge as income per head, education, status as a recipient of disability benefit, owner-occupation and newly entitled status.

The estimated effect of income is always significant at the 5% level but varies considerably over the two samples. For the probit model estimated on sample 1 data, the coefficient implies a large 11% increase in expected claim costs for each additional £1 of original income. This falls to just over 4% when the data from sample 2 is used. For the Klein-Spady estimates the range is similar: a 13% impact on sample 1 data but under 5% for sample 2 data.

Education has a very large effect. Having schooling past age 14 is estimated to cuadruple expected claim costs on the basis of models estimated from sample 1 data. The expected claim costs are lower when sample 2 is used, nevertheless having schooling past the age of 14 more than doubles claim costs. Although better-educated people may have greater capacity to negotiate the intricacies of the benefit system, on this evidence they must also typically be more vulnerable to stigma or tend to be in circumstances entailing greater costs of claiming.

The two disability variables reflect the household's status as a recipient of a (medically assessed but non-means-tested) disability benefit and or as one containing a registered disabled person. These have respectively positive and negative impacts on expected claim costs. Note that registering as a disabled person is voluntary and has no direct implications for benefit entitlement, but may bring other benefits such as subsidised transport, unrestricted car parking, etc. Unfortunately, we cannot observe the true physical state of the household members, so these two variables summarise a combination of factors. One might interpret the coefficient of the former variable as an indicator of physical impairment which increases the physical difficulty of coping with the IS claims process and thus increases implicit claim costs (roughly twofold). The latter variable might be interpreted as an indicator of low vulnerability to stigma: those who are willing to seek formal recognition of disability may also tend to be more willing to accept an IS-dependent

[9]To save space, we do not reproduce intermediate estimates here. Further details are available from the authors on request.

status and thus have lower expected claim costs (by around 58%). In the absence of direct information on physical capacity, such interpretations are necessarily speculative.

Housing tenure is closely linked to social status as well as wealth. Being a homeowner greatly increases the barriers to IS take-up, increasing estimated mean claim costs as much as eight-fold. This finding is likely to reflect the relatively poor access that homeowners have to information, help and advice, which is available (through housing associations and local authority housing offices) to the great majority of renters.[10]

In general, the probit and the Klein-Spady estimates have similar qualitative implications in all the samples considered here. However there is an important difference for age, which plays a significant role in the probit. Using the probit model, claim costs are estimated to increase with age, although at a decreasing rate. If accepted, this result would be hard to rationalise. It seems unlikely that people who claim benefit when younger would cease to do so when they reach a critical age. If the age effect arises through the acquisition of information through time, one would most plausibly expect the take-up rate to be increasing. Adjustment models based on random durations of periods of need (see Anderson and Meyer, 1997) are inappropriate here and again suggest rising take-up rates. The most plausible interpretation would be that the age variable reflects a cohort effect implying a gradual upward drift in take-up rates over time but this conflicts with the absence of a trend in IS take-up among pensioners at the macro level (DWP (2004a) and earlier issues). The issue is resolved once the more flexible semiparametric approach is used, since age becomes insignificant for all samples. This last finding provides a good illustration of the often-neglected proposition that misspecification of distributional form can cause serious biases in binary response models.

Finally, we have included in the model a dummy variable to reflect the possibility that there is some delay in adjusting to changes in the rules of the benefit system. In April 2001 there was a major revision in the IS rules, which

[10] Since renters who are entitled to IS are also entitled to both HB and CTB and owners entitled to IS are also entitled to CTB we also estimated two additional models using the cleaned dataset. The first model used total entitlement to all benefits as the entitlement amount while in the second, total entitlement was used only for renters and IS entitlement for owner-occupiers. The model with entitlement to IS produced a better fit.

significantly extended entitlement. A dummy variable is used to identify cases of individuals who are entitled in the year they are interviewed but who would have been non-entitled (with unchanged circumstances) under the previous year's rules. This variable has a strongly significant coefficient, implying the existence of adjustment lags.

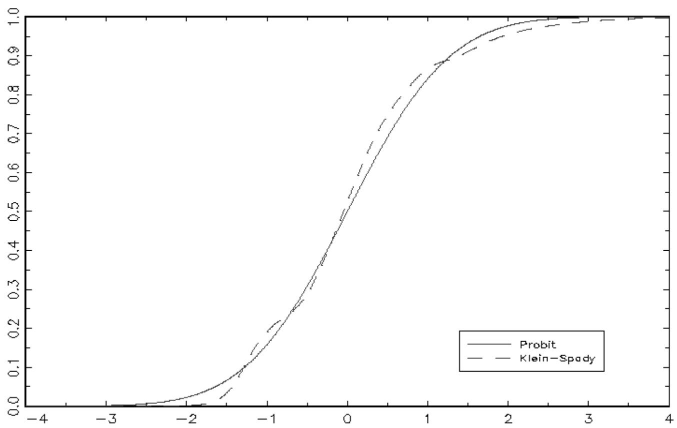
Table 2 Parametric and semi-parametric coefficient estimates (Scaled coefficients $\widehat{\boldsymbol{\alpha}}$)

	Sample 1[1]		Sample 2[2]	
	Probit	**Klein-Spady**	**Probit**	**Klein-Spady**
Variable	$\widehat{\boldsymbol{\alpha}}$ $\lvert\mathbf{t}\rvert$	$\widehat{\boldsymbol{\alpha}}$ $\lvert\mathbf{t}\rvert$	$\widehat{\boldsymbol{\alpha}}$ $\lvert\mathbf{t}\rvert$	$\widehat{\boldsymbol{\alpha}}$ $\lvert\mathbf{t}\rvert$
Single male household	-2.267	-2.809	-1.064	-1.232
	(3.104)	(2.770)	(2.702)	(2.371)
Single female household	-3.528	-3.688	-1.947	-2.026
	(4.328)	(3.431)	(4.960)	(4.022)
Age/10	8.746	-1.985	6.029	-0.738
	(2.056)	(0.386)	(2.235)	(0.220)
$(\text{Age}/10)^2$	-0.481	0.183	-0.331	0.092
	(1.800)	(0.561)	(1.946)	(0.435)
Income per person	0.102	0.122	0.041	0.047
	(4.422)	(3.542)	(4.649)	(3.865)
Head educated past 14	1.389	1.325	0.930	0.918
	(3.627)	(2.757)	(4.084)	(3.088)
Disability benefit	0.832	0.600	0.715	0.634
	(2.289)	(1.210)	(3.037)	(1.951)
Registered disabled	-1.248	-1.028	-0.840	-0.872
	(2.735)	(1.769)	(3.006)	(2.293)
Owner occupier	3.441	2.916	2.257	2.140
	(5.968)	(4.605)	(8.144)	(6.343)
Rent free	2.299	2.283	1.257	1.383
	(2.638)	(2.164)	(2.384)	(2.089)
Newly entitled	2.281	1.775	1.585	1.500
	(4.047)	(2.726)	(4.480)	(3.359)

[1] $n = 4003$. [2] $n = 4129$.

Figure 2 shows the distribution functions $\hat{F}(.)$ for the probit and the Klein-Spady models in sample 2. To make these comparable, the probit probability, $\Phi(.)$, is plotted against the standardised Klein-Spady estimate. The most striking difference between the two distributions is the fatter upper tail of the Klein-Spady estimate and a local concentration at around -1 standard deviations in the lower tail.

Figure 2 Estimated distribution functions for the probit and Klein-Spady models.



5.2 Estimates of the implicit stigma/claim costs

5.2.1 Claim costs incurred by claimants

Table 3 shows means and medians of the estimated claim costs for the subsample of pensioners receiving Income Support. These estimates are constructed using expressions (6) and (10), which give quite different results because of the skewness in the lognormal distribution for C. The semi-parametric estimates give substantially higher estimated claim costs than the probit model, regardless of the method used to construct the implicit costs. The results are rather sensitive to the choice of sample, with larger costs estimated for samples 2, where recorded rather than simulated benefit receipt is used when possible. Even using the preferred semi-parametric estimates the average estimated claim cost for IS recipients is moderate, averaging around £3.97 per week in the preferred sample 2 (or 15% of mean entitlement) and £3.40 in sample 1 (13% of mean entitlement).

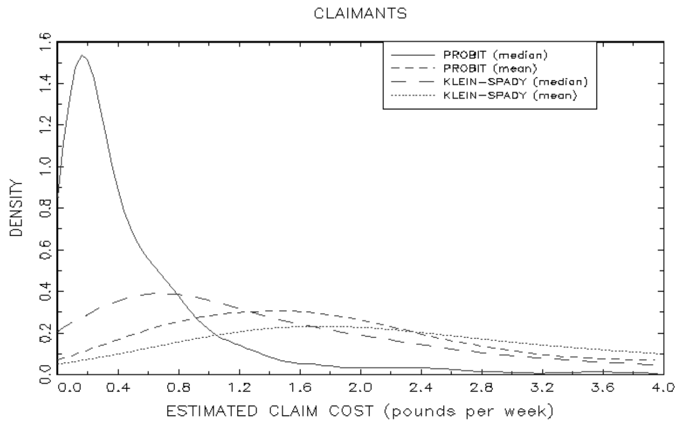
Table 3 Summary measures of estimated stigma/claim costs for Income Support recipients (£ per week) Standard errors in brackets

		Mean	Median
Probit:	Sample 1[1]	1.70	1.20
conditional mean		(0.26)	(0.19)
method ($\widehat{C}_1$)	Sample 2[2]	2.78	1.96
		(0.27)	(0.19)
Probit:	Sample 1[1]	0.10	0.04
conditional median		(0.01)	(0.00)
method ($\widehat{C}_2$)	Sample 2[2]	0.53	0.29
		(0.17)	(0.11)
Klein-Spady:	Sample 1[1]	3.40	2.20
conditional mean		(0.08)	(0.25)
method ($\widehat{C}_1$)	Sample 2[2]	3.97	2.61
		(0.08)	(0.19)
Klein-Spady:	Sample 1[1]	1.85	0.95
conditional median		(0.05)	(0.23)
method ($\widehat{C}_2$)	Sample 2[2]	2.07	1.20
		(0.05)	(0.17)
Entitlement	Sample 1[1]	25.51	15.19
to IS among		(0.48)	(0.60)
IS recipients	Sample 2[2]	26.74	15.95
		(0.47)	(0.61)

[1]$n = 2527$; [2]$n = 2653$;

Figure 3 shows the empirical distribution of these estimated claim costs for the subset of pensioners within sample 2 who are observed to be in receipt of IS. The Klein-Spady estimates imply greater dispersion, especially when the conditional mean method is used to construct the implicit costs.

Figure 3 Kernel estimates of the distributions of stigma/claim costs for IS recipients (sample 2)

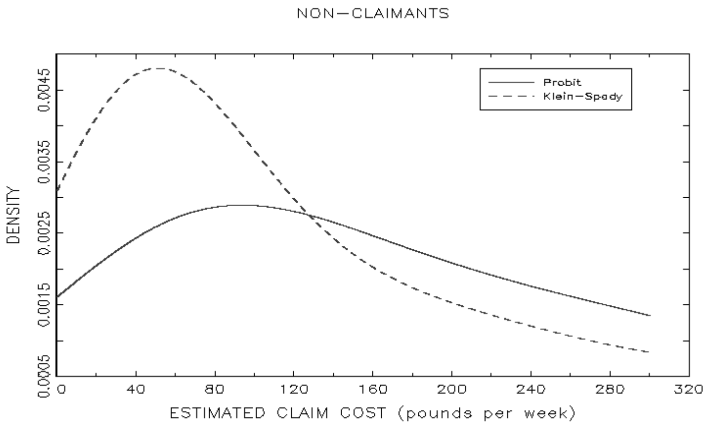


How do these estimates compare with others in the literature? There are no directly comparable figures available, since other researchers have not taken account of the conditioning on observed take-up which is appropriate. For example, Blundell *et. al.* (1988 p.72) estimated claim costs by finding the level of entitlement at which the take-up probability is 0.5. This approach ignores the self selection problem which is overcome by expressions (10) and (11). Duclos (1995 p. 409) finds some illustrative expected costs of claiming Supplementary Benefits (SB) in Britain using the 1985 FES for benefit units with different characteristics. Among the cases depicted for pensioners, take-up costs range from over £3 per week for single pensioners to over £20 for couples. These expected costs are however not conditional on the take-up event. It is possible to estimate the scale of claim costs using published estimates of take-up models. The analysis closest to our own is the work on Housing Benefit (HB) by Blundell *et. al.*, using Family Expenditure Survey data for 1984. From the published probit coefficients and sample means relating to retired/unoccupied respondents (Blundell *et. al.*, 1988, pages 73-74), we can apply the predictors (7) and (10) to estimate implicit claim costs for the average 1984 pensioner claimant. Respectively, these come to £1.79 and £1.08 (updated to 2002 prices) using the conditional mean and median methods. These are comparable with our 1997-2002 estimates for IS.

5.2.2 Claim costs faced by non-claimants

The claim costs faced by those who do not participate in the IS programme are impossible to estimate reliably. For participants, claim costs are bounded by the amount of entitlement B but for non-participants, they are unbounded. The conditional mean method in particular is numerically unstable because it is heavily influenced by the tail behaviour of the function $F(.)$, which is not well-determined statistically. To get good estimates of the upper tail of the claim costs distribution, we would need to observe some cases with very large amounts of entitlement but this is prevented by the design of the benefit system. To illustrate this, Figure 4 compares the distributions of estimated claim costs of IS non-participants for the probit and Klein-Spady models. Among non-participants, the estimates suggest a highly skewed distribution, with a long upper tail. This is especially true for the probit model, which lacks the flexibility of the semi-parametric approach.

Figure 4 Kernel estimates of the distributions of stigma/claim costs for IS non-recipients (sample 2)



6 Implications for poverty measurement

How much difference does allowance for claim costs make to the empirical measurement of pensioner poverty? To answer this satisfactorily we need to make use of the whole distribution of claim costs, rather than its mean or median. We have a poverty line $T(\mathbf{X})$ which may depend on the demographic characteristics of the benefit unit. Ignoring implicit claim costs, we count a pensioner unit as being in poverty if their total net income $Y + B$ falls below the threshold where B is now defined as actual IS receipt. Define S to be the number of individuals in the benefit unit. We use the poverty measure of Foster *et. al.* (1984), denoted here FGT. This measure weights individuals in poverty according to their distance below the poverty threshold. We set the poverty-aversion parameter to 2, so that the definition is:

$$FGT = \frac{E\left[S\,Q\left(Y, B, \mathbf{X}\right)\right]}{E\left(S\right)} \tag{14}$$

where:

$$Q(Y, B, \mathbf{X}) = \begin{cases} \left(1 - \frac{Y+B}{T(\mathbf{X})}\right)^2 & \text{if } Y + B \leq T(\mathbf{X}) \\ \\ 0 & \text{otherwise} \end{cases} \tag{15}$$

A baseline estimate of this measure can be computed by replacing the expectations in (14) with sample averages:

$$\widehat{FGT} = \sum_{i=1}^{n} S_i Q(Y_i, B_i, \mathbf{X}_i) \,/\, \sum_{i=1}^{n} S_i \tag{16}$$

This measure can be adjusted for claim costs by using the estimated costs directly. In this case, the function $Q(Y, B, \mathbf{X})$ in (16) is substituted by

$$Q^*(Y, B, \widehat{C}, \mathbf{X}) = \begin{cases} \left(\frac{T(\mathbf{X})-Y}{T(\mathbf{X})}\right)^2 & \text{if } Y \leq T(\mathbf{X}) \\ & \text{ and } B = 0 \\ \left(\frac{T(\mathbf{X})-Y-B+\widehat{C}}{T(\mathbf{X})}\right)^2 & \text{if } Y + B - \widehat{C} \leq T(\mathbf{X}) \\ & \text{ and } B > 0 \\ 0 & \text{otherwise} \end{cases} \tag{17}$$

Alternatively, we can use an analytical adjustment for claim costs. In general this is preferable since it gives a consistent and more efficient estimate. For those receiving benefit, log claim costs are given by $\ln C = \mathbf{Z}\boldsymbol{\alpha} + V$ and are conditional on the event $\mathbf{Z}\boldsymbol{\alpha} + V < \ln B$. Thus we can estimate the expectation in the numerator of (14) as the sample average of $S_i E\left[Q(Y_i, B_i, \mathbf{X}_i) | B_i, \mathbf{X}_i\right] = S_i Q_i^{**}$ where Q_i^{**} is constructed as follows:

$$Q^{**} = \begin{cases} \left(\frac{T(\mathbf{X})-Y}{T(\mathbf{X})}\right)^2 & \text{if } Y \leq T(\mathbf{X}) \\ & \text{and } B = 0 \\ \frac{\int_{-\infty}^{\ln B} \left(\frac{T(\mathbf{X})-Y-B+C}{T(\mathbf{X})}\right)^2 dF\left(\frac{\ln C - \mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right)}{F\left(\frac{\ln(B)-\mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right)} & \text{if } Y+B \leq T(\mathbf{X}) \\ & \text{and } B > 0 \\ \frac{\int_{\ln(Y+B-T(\mathbf{X}))}^{\ln B} \left(\frac{T(\mathbf{X})-Y-B+C}{T(\mathbf{X})}\right)^2 f\left(\frac{\ln C - \mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right) d\ln C}{F\left(\frac{\ln(B)-\mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right)} & \text{if } Y+B > T(\mathbf{X}), \\ & \text{and } Y < T(\mathbf{X}) \\ 0 & \text{otherwise} \end{cases} \tag{18}$$

where Y is the net income of the benefit unit excluding benefits and the poverty line, $T(\mathbf{X})$, is a percentage of the IS guaranteed minimum for the benefit unit M.

The results are given in Tables 4 and 5. The effects of adjusting for claim costs are moderate. Depending on the sample, threshold and estimator used, measured poverty is some 4-13% higher when claim costs are taken into account. This is not a dramatic impact, but it is non-negligible.

Table 4 Foster-Greer-Thorbecke poverty measures (probit model).

			Poverty line			
			1.2 M	1.1 M	M	0.9 M
Sample 1	$n = 17,089$	$\widehat{FGT}$	0.0148	0.0082	0.0045	0.0028
		$\widehat{FGT}^*$ (mean)	0.0154	0.0086	0.0045	0.0028
		$\widehat{FGT}^*$ (median)	0.0149	0.0082	0.0045	0.0028
		$\widehat{FGT}^{**}$	0.0156	0.0088	0.0047	0.0029
Sample 2	$n = 17,081$	$\widehat{FGT}$	0.0154	0.0088	0.0049	0.0031
		$\widehat{FGT}^*$ (mean)	0.0162	0.0093	0.0051	0.0031
		$\widehat{FGT}^*$ (median)	0.0155	0.0089	0.0050	0.0031
		$\widehat{FGT}^{**}$	0.0165	0.0096	0.0054	0.0033

Table 5 Foster-Greer-Thorbecke poverty measures (Klein-Spady model).

			Poverty line			
			1.2 M	1.1 M	M	0.9 M
Sample 1	$n = 17,089$	$\widehat{FGT}$	0.0148	0.0082	0.0045	0.0028
		$\widehat{FGT}^*$ (mean)	0.0160	0.0090	0.0047	0.0028
		$\widehat{FGT}^*$ (median)	0.0154	0.0086	0.0046	0.0028
		$\widehat{FGT}^{**}$	0.0162	0.0092	0.0049	0.0029
Sample 2	$n = 17,081$	$\widehat{FGT}$	0.0154	0.0088	0.0049	0.0031
		$\widehat{FGT}^*$ (mean)	0.0166	0.0096	0.0052	0.0031
		$\widehat{FGT}^*$ (median)	0.0160	0.0092	0.0050	0.0031
		$\widehat{FGT}^{**}$	0.0169	0.0099	0.0055	0.0033

7 Conclusions

This paper studies the take-up of Income Support by UK pensioners using data on the financial years 1997/8-2001/2002 from the British Family Resources Survey. Two binary choice models of IS take-up are estimated: a probit model and a more flexible semiparametric model. In addition to the (log) level of entitlement, the main factors contributing to high claim costs are income per head, education, status as a recipient of disability benefit, owner-occupation and newly entitled status. Using a revealed preference approach we consider the implicit costs of claiming Income Support. These costs might arise from the onerous nature of the claims process, from social stigma associated with being on welfare and from the difficulty of acquiring information about the benefit system. We develop a new technique of constructing individual-specific estimates of claim costs, allowing for the self-selection effect of the take-up process. Implicit costs are found to be moderate for most IS recipients, typically around £3-4 per week (or about 13-15% of entitlement) for the average benefit recipient, and consequently the degree of measured poverty among pensioners increases by a modest but non-negligible amount (up to 13% for the Foster-Greer-Thorbecke index) when these claim costs are taken into account.

The revealed preference approach argues that non-participants judge themselves to be better off foregoing than claiming their entitlements because of these costs. It does not follow from our results, however, that non-participation is no cause for concern. The fact that some eligible individuals choose not to participate in means-tested programmes simply indicates that they find living below the poverty line preferable to living on welfare. If governments want to use means-tested welfare programmes to prevent poverty, they need to find ways to reduce the size of the costs involved relative to the size of the benefits paid out.

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Appendix: Additional tables

Table A1 Weekly rates of principal social security benefits applicable to pensioners in the 1997-8, 1998-9 and 1999/0 FRS

		£ per week		
		1997/8	1998/9	1999/0
Basic state pension	Full rate	62.45	64.70	66.75
	'Married woman's' rate	37.35	38.70	39.35
Age 80+ addition to state pension		0.25	0.25	0.25
Attendance Allowance	Higher rate	49.50	51.30	52.95
	Lower rate	33.10	34.30	35.40
Disability Living Allowance (care component)	Highest rate	49.50	51.30	52.95
	Middle rate	33.10	34.30	35.40
	Lowest rate	13.15	13.60	14.05
Disability Living Allowance (mobility component)	Higher rate	34.60	35.85	37.00
	Lower rate	13.15	13.60	14.05
Income Support for pensioners	single pensioner under 75	68.80	70.45	75.00
	single pensioner 75-79	71.00	72.65	77.30
	single pensioner 80+	75.70	77.55	82.25
	single pensioner with SDP	112.85	116.05	122.00
	couple, both under 75	106.80	109.35	116.60
	couple, one or both 75-79	109.90	112.55	119.85
	couple, one or both 80+	115.15	117.90	125.30
	couple, one or both 75-79, one with CP	123.25	126.20	133.80
	couple, one or both 80+, with CP	128.50	131.55	139.25
	couple, both with SDP	189.45	194.90	204.80
	upper capital threshold	8000	8000	8000
	lower capital threshold	3000	3000	3000

Notes: It is not possible to receive both Attendance Allowance and the care component of Disability Living Allowance. Disability Allowance (care and mobility component) is payable to people aged 65+ only if they started to receive it before reaching 65.
[1] CP = Carer Premium; [2] SDP = Severe Disability Premium

Table A1 (cont) Weekly rates of principal social security benefits applicable to pensioners in the 2000-1 and 2001-2 FRS

		£ per week 2000/1	2001/2
Basic state pension	Full rate	67.50	72.50
	'Married woman's' rate	40.40	43.40
Age 80+ addition to state pension		0.25	0.25
Attendance Allowance	Higher rate	53.55	55.30
	Lower rate	35.80	37.00
Disability Living Allowance (care component)	Highest rate	53.55	55.30
	Middle rate	35.80	37.00
	Lowest rate	14.20	14.65
Disability Living Allowance (mobility component)	Higher rate	37.40	38.65
	Lower rate	14.20	14.65
Income Support for pensioners	single pensioner under 75	78.45	92.15
	single pensioner 75-79	80.85	92.15
	single pensioner 80+	86.05	92.15
	single pensioner with SDP	126.25	133.70
	couple, both under 75	121.95	140.55
	couple, one or both 75-79	125.35	140.55
	couple, one or both 80+	131.05	140.55
	couple, one or both 75-79,	171.25	182.10
	one with CP	136.10	164.95
	couple, one or both 80+,	139.50	164.95
	with CP	145.20	164.95
	couple, both with SDP	211.45	223.65
	upper capital threshold	8000	12000
	lower capital threshold	3000	6000

Notes: It is not possible to receive both Attendance Allowance and the care component of Disability Living Allowance. Disability Allowance (care and mobility component) is payable to people aged 65+ only if they started to receive it before reaching 65.
[1] CP = Carer Premium; [2] SDP = Severe Disability Premium

Table A2 Variable definitions

Variable	Definition
$\ln(B_i)$	Log of IS entitlement as calculated in Sample i (£ per week)
Single male household	Dummy variable = 1 for single-man household, 0 otherwise
Single female household	Dummy variable = 1 for single-woman household, 0 otherwise
Age	Age of the head of the household
Income per head	= Net income (£ per week) excluding IS per person in the household
Head educated past 14	Dummy variable = 1 if household head left school aged 15 or more, 0 otherwise
Disability benefit	Dummy variable=1 if any person in the household receives AA , DLA self care and/or Mobility component of DLA
Registered Disabled	Dummy variable = 1 if any person in the household is registered as disabled with the LA
Owner occupier	Dummy variable = 1 if the household owns the house
Rent free	Dummy variable = 1 if the household is non-owner-occupier and lives rent-free
Newly entitled 01/02 survey	Dummy variable = 1 if the household was not entitled before the reforms (01/02 survey)

Table A3 Sample means of explanatory variables and probit coefficients (cleaned data; standard errors in parentheses)

Variable	Sample mean	Probit coeff[1]	Sample mean	Probit coeff[2]
$\ln(B_1)$	2.547	0.161		
	(0.019)	(0.024)		
$\ln(B_2)$			2.599	0.244
			(0.019)	(0.023)
Single male household	0.142	0.364	0.141	0.259
	(0.006)	(0.091)	(0.005)	(0.088)
Single female household	0.743	0.566	0.746	0.475
	(0.007)	(0.077)	(0.007)	(0.074)
Age/10	7.849	-1.405	7.856	-1.471
	(0.011)	(0.660)	(0.011)	(0.648)
$(\text{Age}/10)^2$	62.084	0.077	62.185	0.081
	(0.171)	(0.042)	(0.169)	(0.041)
Income per head	70.192	-0.016	71.014	-0.010
	(0.306)	(0.002)	(0.315)	(0.002)
Head educated past 14	0.226	-0.223	0.223	-0.227
	(0.007)	(0.053)	(0.006)	(0.052)
Disability benefit	0.299	-0.134	0.320	-0.174
	(0.007)	(0.068)	(0.007)	(0.065)
Registered Disabled	0.152	0.202	0.161	0.205
	(0.006)	(0.067)	(0.006)	(0.065)
Owner occupier	0.286	-0.552	0.285	-0.550
	(0.007)	(0.047)	(0.007)	(0.047)
Rent free	0.027	-0.369	0.028	-0.306
	(0.003)	(0.129)	(0.003)	(0.125)
Newly entitled 01/02 survey	0.099	-0.366	0.085	-0.386
	(0.005)	(0.075)	(0.004)	(0.078)
n	4003		4129	

[1]Intercept = 7.155 (std err = 2.597);
[2]Intercept =6.911 (std err = 2.553)