THE WELFARE COST OF MEANS-TESTING: PENSIONER PARTICIPATION IN INCOME SUPPORT *

Stephen Pudney
Department of Economics, University of Leicester

Monica Hernandez
Department of Economics, University of Leicester

Ruth Hancock
Nuffield Community Care Studies Unit, University of Leicester

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ABSTRACT: We estimate parametric and semi-parametric binary choice models of benefit take-up by British pensioners and develop methods, based on the compensating variation, of inferring the cash-equivalent value of disutility arising from stigma or complexity of the claims process, taking account of the self-selection induced by take-up. We extend this to incorporate the claim costs borne by welfare-recipients into the measurement of pensioner poverty. We find that allowing for implicit claim costs in this way has a non-negligible impact on poverty measurement.

KEYWORDS: benefit take-up, program participation, pensions, welfare, poverty

JEL CLASSIFICATION: C25, D63, H55, I32, I38

ADDRESS FOR CORRESPONDENCE: Stephen Pudney, Department of Economics, University of Leicester, University Road, Leicester LE1 7RH, UK; e-mail: sep2@le.ac.uk.

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 Altmann (1981), Moffitt (1983), Fry and Stark (1987, 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[1] and the implicit risk premium associated with the unpredictability of the claim outcome (see, for example, Halpern and Hausman, 1986). 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

[1] 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.

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. Almost 40% of British pensioners receive at least one of three means-tested benefits available to pensioners (Department of Social Security, 2000). The scope of means-tested pensioner benefits will be 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 (IS) which provides general income maintenance. Official estimates are that in the financial year 1999/2000 about 30% of pensioners who appeared to be entitled to IS did not receive it (Department for Work and Pensions, 2001).

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. Latest figures show that the average total state pension payment (basic pension, SERPS and other minor components) in September 2001 was marginally above the full basic state pension (Department for Work and Pensions, 2002)' but below the means-tested benefit level.

2.1 Means-tested benefit rules

There are 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.

IS is assessed and paid to pensioner units – single pensioners or pensioner couples. Entitlement to IS is zero if the pensioner unit's financial holdings are above an upper threshold (£8,000 during our sample period). 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). 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. Certain kinds of income, such as AA and DLA, are excluded in part or in full from assessable income. Actual income from capital is also excluded. Instead a notional income from capital between a lower threshold (£3,000 during our sample period) 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. 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). In the period we are concerned with, this entailed completing and taking a 40-page form to their local social security office. 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) HB, and IS is almost never received in isolation.[2]

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 will subsume 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 late 1990s, the period considered here, few pensioners would have expected the experience of claiming IS to be enjoyable or hassle-free.

[2] In our sample period only around 0.56% of IS recipients received neither HB nor CTB. Only 0.89% of those receiving IS+HB did not also receive CTB; among renters, only 1.36% of those receiving IS+CTB did not also receive HB.

3 Take-up: evidence from the 1997-2000 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 2000, 71,378 households containing 169,053 individuals participated in the survey. All adult respondents were asked whether they were receiving each of a comprehensive set of social security benefits and if so, the amount they last received. Details of private sources of income, capital holdings, personal and other characteristics relevant to calculating entitlement to IS are also recorded. In principle, therefore, we can assess each FRS pensioner unit's entitlement to IS, establish whether they are receiving IS and develop a model of take-up behaviour. In practice the data are likely to be subject to measurement error.

The pensioner rates of Income Support apply to single people aged 60 years or more or couples where either partner is aged at least 60. In the combined 1997-2000 FRS there were 26,229 such pensioner units. To simplify the analysis we restricted the sample to a final selection of 12,801 pensioner units in the following way:

(1) Only those at least five years above state pension age (i.e. men aged 70+ and women aged 65+) were included and those with employment or self-employment were excluded. Restricting the sample to those at least five years above state pension age excluded 10,339 pensioner units. Excluding those with income from employment or self-employment reduced the sample by a further 525.

(2) Only households containing a single pensioner aged at least 65 (female) or 70 (male) or a couple where both partners were five or more years above state pension age were included. The presence of other household members considerably complicates the calculation of IS entitlement and increases the scope for measurement error. This restriction excluded another 2,140 pensioner units. Of these, 15 contained children but no other adults; the remainder contained other adults.

(3) Households still re-paying a mortgage were excluded. Such payments affect the calculation of IS entitlement, are a potential source of measurement

error but exist for only a small minority of the age group of interest; 413 cases were excluded for this reason.
(4) Eleven cases in receipt of allowances from an absent spouse were also excluded.

3.2 Data cleaning

Any errors in recorded income (including state benefits) or capital will lead to errors in assessed entitlement. If income and capital are used as explanatory variables in an econometric model of take-up behaviour, such measurement errors have the potential to result in serious bias in model estimates with corresponding implications for their use in policy analysis (Pudney 2001). To minimise the potential bias from measurement error, we applied a process of error detection and correction to the data before using them to model take-up behaviour.

The first step was to reverse data edits and imputations made by DWP, affecting benefit receipts, private pension income and capital holdings. This was because we detected some inconsistencies in edits to benefit data and because some of their procedures (e.g. substitution of sample means for missing values) are not appropriate when attempting to calculate means-tested benefit entitlement. 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 where a correct value could be identified. For example, some pensioners in the FRS are able to supply a breakdown of their state pension payments which helps to 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 (2002).

For this paper, we excluded observations where recorded benefit is positive, IS-assessable income exceeds the guaranteed minimum but no IS documentation was consulted by the survey respondent. A further assumption was made to reduce the number of cases dropped due to missing values. In cases with missing assessable income but with non-missing IS and capital, assessable income was set to the amount consistent with the size of the recorded IS and the appropriate IS guaranteed minimum.

Two different variants of the dataset are considered throughout the paper for both the pre-cleaned and cleaned data. These samples differ in the entitlement measure used and give an idea of the robustness of our results. Sample 1, uses simulated IS entitlement based on assessable income, capital and the guaranteed minimum. Sample 2 substitutes recorded IS, where available, for simulated entitlement provided it does not exceed the guaranteed minimum for the benefit unit and provided the respondent consulted IS documentation. In such cases and in cases where IS receipt is missing, simulated IS entitlement is used instead. In addition, cases where recorded IS benefit or assessable income was higher than the relevant guaranteed income level and cases where the guaranteed level was smaller than recorded IS benefit plus notional income from capital were excluded from sample 2. In general, we regard the cleaned version of sample 2 as the most likely to give an accurate picture of IS entitlement.[3]

3.3 Take-up rates

Table 1 shows estimated IS take-up rates for the two samples, before and after cleaning the data. These rates are largely unaffected by the choice of sample and data cleaning. There is some variation by category of pensioner, but the typical rate of non-participation is roughly one third, ranging for different samples from 33.5% to 35.4%. Although not directly comparable, these are close to the official estimate of 30% reported by DWP for 1999/2000. In all the samples considered here, both before and after cleaning the data, single females appear to have higher take-up rates than single males and couples. Take-up rates are also higher in all the samples for pensioners in the younger age groups. Pensioners who left full-time education

[3]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.

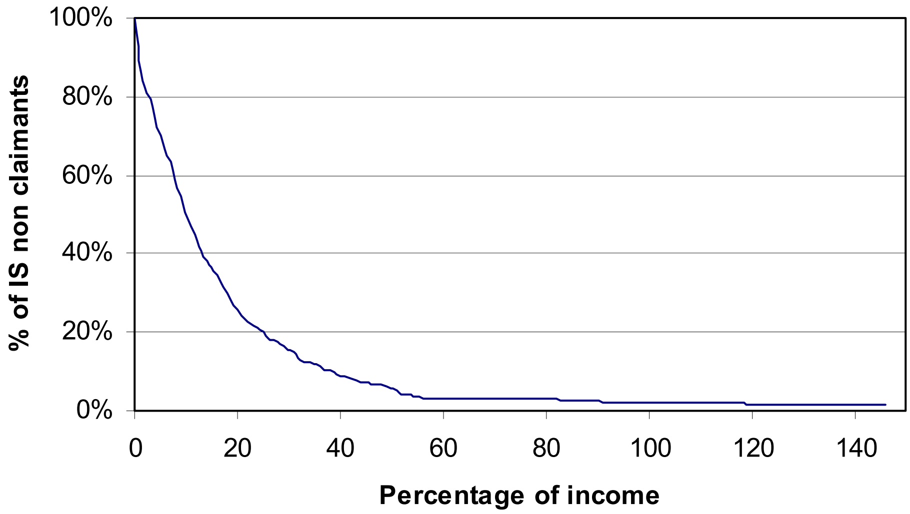
after the age of fourteen have lower take-up rates. Take-up rates are higher for pensioners with disabilities but the reverse is true for those in receipt of disability benefit. Take-up varies considerably with housing tenure, renters having much higher rates than home owners. Pensioners living in Scotland and Wales have relatively low take-up rates in all samples.

Table 1 Percentage take-up rates by demographic groups
(1997/8-1999/00 FRS estimation sample; weighted sample frequencies)

Sample group	**Take-up rate (std. error)**			
	original sample		**cleaned sample**	
	1	2	1	2
Single male	57.3	58.4	55.4	56.5
	(2.4)	(2.4)	(2.7)	(2.6)
Single female	67.6	68.3	69.2	69.6
	(1.0)	(1.0)	(1.1)	(1.1)
Couple	52.3	56.9	54.6	58.3
	(3.0)	(2.8)	(3.2)	(3.0)
Head under 70	79.3	79.9	79.2	79.5
	(2.4)	(2.3)	(2.5)	(2.5)
Head 70-79	66.2	67.6	67.5	68.5
	(1.4)	(1.3)	(1.5)	(1.4)
Head 80-89	60.1	60.9	61.3	61.9
	(1.4)	(1.4)	(1.5)	(1.5)
Head 90+	61.4	61.8	59.7	60.6
	(3.6)	(3.6)	(4.2)	(4.2)
Education < 14	72.3	73.0	69.5	70.1
	(3.6)	(3.5)	(4.0)	(4.0)
Education equal to 14	65.8	67.0	67.2	68.1
	(1.0)	(1.0)	(1.1)	(1.1)
Education > 14	58.2	58.7	59.2	59.3
	(2.0)	(2.0)	(2.2)	(2.2)
Not receiving disability benefit	65.3	66.5	66.1	66.9
	(1.1)	(1.1)	(1.2)	(1.2)
Receiving disability benefit	63.6	64.4	64.9	65.8
	(1.4)	(1.4)	(1.7)	(1.6)
Not disabled	63.8	64.8	64.9	65.7
	(1.0)	(1.0)	(1.1)	(1.1)
Either disabled	68.3	69.2	69.6	70.3
	(2.0)	(1.9)	(2.2)	(2.2)
Owner occupier	48.6	50.4	49.9	51.5
	(1.8)	(1.7)	(2.0)	(1.9)
Rent	71.6	72.4	72.3	72.9
	(1.0)	(1.0)	(1.1)	(1.1)
Rent free	54.4	54.8	54.2	54.2
	(4.9)	(4.9)	(5.5)	(5.5)
Scotland or Wales	60.3	61.7	61.2	62.2
	(2.1)	(2.1)	(2.4)	(2.3)
Rest of Britain	65.6	66.5	66.6	67.4
	(1.0)	(0.9)	(1.1)	(1.0)
Full sample	64.6	65.7	65.7	66.5
	(0.9)	(0.9)	(1.0)	(0.9)
Sample size	2929	3019	2417	2475

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. 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%. 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 revealed preference 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 (for example feelings of inadequacy or shame induced by dependency). Formally, the analysis is based on an assumption of full awareness of the benefit system. However, total lack of awareness of the existence of social security benefits is probably very rare indeed. We argue that awareness is itself partly a matter of choice, since information can be acquired with a certain expenditure of effort. The effort required to become informed can then be treated as a component of claim costs.

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 take-up model

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 can be interpreted as a compensating variation: the cash equivalent of any stigma or claim costs acting as a disincentive to take-up.

Empirically, a better fit is obtained by working with the logarithm of benefit entitlement. 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. Note that this approach can also be used in the case where the outcome of a claim is uncertain. In that case expected utility can be written:

$$EU = PU_1(Y+B;\mathbf{X},V) + (1-P)U_2(Y;\mathbf{X},V) \tag{3}$$

where P is the subjective probability of success and where success and failure can both entail stigma or cost, through the forms U_1, U_2. In this case, (2) becomes[4]:

$$B > U_1^{-1}\left(\frac{U_0(Y;\mathbf{X},V)-(1-P)U_2(Y;\mathbf{X},V)}{P};\mathbf{X},V\right) - Y \tag{4}$$

As an alternative to the approximation of (2) or (4) by $\mathbf{Z}\boldsymbol{\alpha}+V$, we can follow Moffitt (1983) and use an explicit specification of the preference shift induced by stigma or claim costs. For example, in the certainty case, we might assume:

$$U_1(Y;\mathbf{X},V) = U_0(Y-e^{\mathbf{Z}\boldsymbol{\alpha}+V};\mathbf{X},V) \tag{5}$$

[4]Condition (4) is equivalent to the certainty case if U_2 is assumed identical to U_0: in other words, if stigma only arises from benefit receipt rather than the claim process.

Thus the claim cost is treated as a fixed cash-equivalent amount varying with observable individual characteristics (including original income) and with unobservable individual-specific factors. In this case:

$$U_1^{-1}\left(U_0(Y;\mathbf{X},V);\mathbf{X},V\right) - Y = e^{\mathbf{Z}\boldsymbol{\alpha}+V} \tag{6}$$

and the condition for take-up is:

$$\ln B > \mathbf{Z}\boldsymbol{\alpha} + V \tag{7}$$

The conditional take-up probability is then:

$$\begin{aligned} \Pr(\text{take-up}|B,\mathbf{Z}) &= \Pr(V < \ln B - \mathbf{Z}\boldsymbol{\alpha}) \\ &= F\left(\frac{\ln B - \mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right) \end{aligned} \tag{8}$$

where $\sigma^2 = var(V)$ and $F(.)$ is the distribution function of the random variable V/σ. The probability (8) amounts to a standard binary response model of discrete choice, using $\ln B$ and $\mathbf{Z}$ as explanatory variables. In such a model, 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 $\boldsymbol{\alpha}$, an estimate 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.

4.2 Identification

In general terms, the model is of the form:

$$\Pr(\text{take-up}|Y,\mathbf{X}) = G(B(Y,\mathbf{X}),Y,\mathbf{X}) \tag{9}$$

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 $B(Y,\mathbf{X})$ 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.[5] There are various restrictions that might be used to resolve

[5] For example, replace B by $\lambda B + (1-\lambda)B$ where λ is an arbitrary function of $Y,\mathbf{X}$. Now rewrite the model $G(\lambda B, \mathbf{Z}^*)$, where $\mathbf{Z}^* = [(1-\lambda)B(Y,\mathbf{X}) : Y : \mathbf{X}]$. This is observationally equivalent to the original model. This identification problem is characteristic of most attempts to capture the incentive effects of tax and social security systems from cross-section surveys.

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, however, a number of 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 £3000). 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. The amount of capital itself is excluded.

4.3 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 an IS recipient:

$$\begin{aligned} \widehat{C}_1 &= E(\exp(\mathbf{Z}\boldsymbol{\alpha} + V)|V < \ln B - \mathbf{Z}\boldsymbol{\alpha}) \\ &= e^{\mathbf{Z}\boldsymbol{\alpha}} \left[\int_{-\infty}^{\ln B - \mathbf{Z}\boldsymbol{\alpha}} e^V dF(V) \right] \left[\int_{-\infty}^{\ln B - \mathbf{Z}\boldsymbol{\alpha}} dF(V) \right]^{-1} \end{aligned} \qquad (10)$$

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

$$\widehat{C}_1 = \exp\left(\mathbf{Z}\boldsymbol{\alpha}+\frac{\boldsymbol{\sigma}^2}{2}\right)\Phi\left(\frac{\ln B-\mathbf{Z}\boldsymbol{\alpha}-\boldsymbol{\sigma}^2}{\boldsymbol{\sigma}}\right)\Big/\Phi\left(\frac{\ln B-\mathbf{Z}\boldsymbol{\alpha}}{\boldsymbol{\sigma}}\right) \tag{11}$$

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

$$\begin{aligned}\widehat{C}_1 &= E(\exp(\mathbf{Z}\boldsymbol{\alpha}+V)|V\geq \ln B-\mathbf{Z}\boldsymbol{\alpha}) \\ &= \exp\left(\mathbf{Z}\boldsymbol{\alpha}+\frac{\boldsymbol{\sigma}^2}{2}\right)\left[1-\Phi\left(\frac{\ln B-\mathbf{Z}\boldsymbol{\alpha}-\boldsymbol{\sigma}^2}{\boldsymbol{\sigma}}\right)\right]\Big/\left[1-\Phi\left(\frac{\ln B-\mathbf{Z}\boldsymbol{\alpha}}{\boldsymbol{\sigma}}\right)\right]\end{aligned} \tag{12}$$

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

$$F\left(\frac{\ln\widehat{C}_2-\mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right)\Big/F\left(\frac{\ln B-\mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right) = 0.5 \tag{13}$$

and thus:

$$\widehat{C}_2 = \exp\left\{\sigma\left[\mathbf{Z}(\boldsymbol{\alpha}/\sigma)+F^{-1}\left(\frac{1}{2}F\left(\frac{\ln B-\mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right)\right)\right]\right\} \tag{14}$$

For non-claimants, the condition $\Pr(\ln B<\mathbf{Z}\boldsymbol{\alpha}+V<\ln\widehat{C}_2\,|\mathbf{Z}\boldsymbol{\alpha}+V>\ln B) = 0.5$ gives:

$$\widehat{C}_2 = \exp\left\{\sigma\left[\mathbf{Z}(\boldsymbol{\alpha}/\sigma)+F^{-1}\left(\frac{1}{2}\left[1+F\left(\frac{\ln B-\mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right)\right]\right)\right]\right\} \tag{15}$$

Note that $\widehat{C}_1$ and $\widehat{C}_2$ always lie below the unconditional mean and median of $\exp(\mathbf{Z}\boldsymbol{\alpha}+V)$ for participants and above for non-participants. Claimants will, on average, tend to be those who suffer lower than average levels of stigma or claim costs and conversely for non-claimants. The relationship between implicit claim costs and the coefficient of $\ln B$ 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 $\lim_{\sigma\to\infty} \mathbf{Z}(\boldsymbol{\alpha}/\sigma) = -F^{-1}(P)$. Consider the median (14). 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 $C < B$ is greater, the larger is σ. Conversely, $\lim_{\sigma\to\infty} \widehat{C}_2 = +\infty$ for non-claimants: as we increase σ, $\mathbf{Z}\boldsymbol{\alpha}$ must increase towards $\ln B$ 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:

$$\max_{\gamma} \ln L(\boldsymbol{\gamma}) = \sum_{i=1}^{n} \left\{ y_i \ln\left(\widehat{F}(\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha})\right) + (1-y_i)\ln\left(1 - \widehat{F}(\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha})\right)\right\} \tag{16}$$

where $\widehat{F}(.)$ is a nonparametric kernel estimate of the regression function of y_i on $\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}$.[6] We use the Gaussian kernel:

$$\widehat{F}(\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}) = \frac{\sum_{j\neq i} \phi\left(h^{-1}[(\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}) - (\ln B_j - \mathbf{Z}_j\boldsymbol{\alpha})]\right) y_j}{\sum_{j\neq i} \phi\left(h^{-1}[(\ln B_i - \mathbf{Z}_i\boldsymbol{\alpha}) - (\ln B_j - \mathbf{Z}_j\boldsymbol{\alpha})]\right)} \tag{17}$$

where $\phi(.)$ is the standard normal density function. We experimented with a variety of fixed and adaptive bandwidths (the latter using the Breiman *et.*

[6] Note that $\widehat{F}$ is not normalised to have zero mean and unit variance. Scale and location are normalised by fixing the coefficient of b at unity and excluding the intercept term from the linear form $\mathbf{Z}\boldsymbol{\alpha}$. This does not affect the construction of implicit cost estimates.

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.

Tables 2, and 3 give estimates of the stigma/claim cost coefficients $\boldsymbol{\alpha}$. The variables appearing in the model are defined and summarised in Appendix Table A2. For the probit model the estimates are calculated as minus the coefficients of the relevant variables divided by the coefficient of $\ln B_i$. The estimates are the outcome of an extensive process of specification search. The chosen form is superior to other models with alternative functional forms for income and entitlement, fuller location effects (at Standard Region level), demographic structure and a more general specification involving the ages and education levels of both members for 2-person households. To guard against pre-test bias, we have used throughout a conservative criterion, retaining explanatory variables with asymptotic t-ratios in excess of 1.0. 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 location.

The estimated effect of income is always significant (at the 5% level with a 1-tailed test) but varies considerably over the three samples. For the probit model estimated on the pre-cleaned sample 1 data, the coefficient implies a large 12% increase in expected claim costs for each additional £1 of original income. This falls to under 2% when the (presumably more accurate) cleaned data from sample 2 is used. For the Klein-Spady estimates the range is even greater: a 16% impact on the pre-cleaned sample 1 data but under 2% for the cleaned sample 2 data.

Education has a very large effect. Having schooling past age 14 is estimated to almost triple expected claim costs (on the basis of models estimated from the cleaned sample 2 data). 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, the former larger and more significant in our preferred sample 2. 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 threefold for our preferred cleaned sample 2). 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 status and thus have lower expected claim costs (by around 55%). 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 home-owner greatly increases the barriers to IS take-up, increasing estimated mean claim costs six-fold.[7]

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 rather than neighbourhoods and there are, consequently, no very strong locational effects detectable. However, residence in Wales or Scotland rather than England is estimated to raise claim costs and thus reduce take-up: a result that runs counter to what one might expect, since Wales and Scotland have higher average rates of poverty than England. They also have proportionately larger rural areas, so the lower population density than in England may account for this finding. In any case, the effect is relatively modest.

In general, the probit and the Klein-Spady estimates have similar qualitative implications in all the samples considered here. However there is an

[7]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. Although the coefficient of owner-occupiers became smaller in both samples, the model with entitlement to IS fitted better.

important difference for age, which plays a significant role in the probit model for all samples, with and without data cleaning. 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. Adjustment models based on random durations of periods of need (see Anderson and Meyer, 1997) seem inappropriate here and imply 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. No clear trends in IS take-up among pensioners have been observed at the macro level (DWP 2001 and earlier issues). However, once the more flexible semiparametric approach is used, age becomes insignificant for all samples.

Table 2 Parametric and semi-parametric coefficient estimates (Scaled coefficients $\widehat{\boldsymbol{\alpha}}$; sample 1)

	Pre-cleaning[1]		Cleaned[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	-1.687	-2.627	-0.889	-1.947
	(1.917)	(1.872)	(0.944)	(1.358)
Single female household	-3.294	-4.074	-3.582	-4.009
	(3.173)	(2.539)	(2.780)	(2.206)
Age/10	16.494	-2.084	13.894	-1.948
	(2.858)	(0.324)	(2.061)	(0.275)
$(\text{Age}/10)^2$	-0.980	0.180	-0.811	0.184
	(2.731)	(0.443)	(1.928)	(0.409)
Income per person	0.115	0.146	0.104	0.121
	(3.503)	(2.670)	(2.806)	(2.201)
Head educated past 14	1.856	1.840	2.207	1.923
	(3.380)	(2.435)	(3.101)	(2.212)
Disability benefit	1.226	0.777	0.440	0.359
	(3.021)	(1.180)	(0.728)	(0.445)
Registered disabled	-1.171	-1.239	-1.539	-1.248
	(2.206)	(1.559)	(2.215)	(1.434)
Owner occupier	3.635	3.653	3.954	3.409
	(4.549)	(3.426)	(3.920)	(2.922)
Rent free	2.602	2.392	2.738	2.524
	(2.562)	(1.804)	(2.191)	(1.611)
Wales/Scotland	0.872	0.887	1.136	1.032
	(1.891)	(1.350)	(1.943)	(1.398)

[1] $n = 2929$. [2] $n = 2417$.

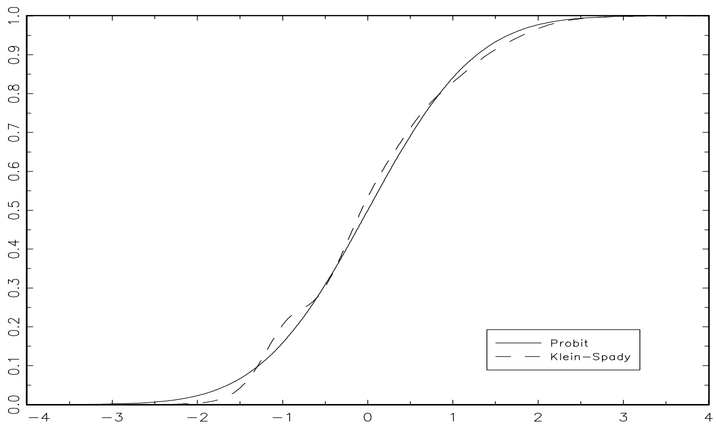
Table 3 Parametric and semi-parametric coefficient estimates
(Scaled coefficients $\widehat{\boldsymbol{\alpha}}$; sample 2)

	Pre-cleaning[1]		Cleaned[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	0.071	-0.061	0.499	0.379
	(0.179)	(0.122)	(1.341)	(0.793)
Single female household	-0.859	-0.975	-0.669	-0.775
	(2.375)	(2.222)	(1.972)	(1.769)
Age/10	10.054	-0.718	7.899	-0.484
	(3.387)	(0.298)	(2.716)	(0.142)
$(\text{Age}/10)^2$	-0.585	0.088	-0.453	0.073
	(3.152)	(0.577)	(2.477)	(0.340)
Income per person	0.030	0.033	0.015	0.016
	(3.802)	(3.379)	(2.181)	(1.748)
Head educated past 14	1.177	1.109	1.087	1.027
	(4.473)	(3.438)	(4.293)	(3.208)
Disability benefit	1.539	1.321	1.286	1.177
	(6.958)	(4.425)	(5.546)	(3.729)
Registered disabled	-0.636	-0.702	-0.595	-0.603
	(2.289)	(1.794)	(2.123)	(1.576)
Owner occupier	2.123	2.087	1.897	1.876
	(7.157)	(6.127)	(7.037)	(5.573)
Rent free	1.432	1.267	1.247	1.265
	(2.742)	(2.003)	(2.436)	(1.947)
Wales/Scotland	0.559	0.573	0.634	0.628
	(2.233)	(1.733)	(2.541)	(1.962)

[1] n =3019 . [2] n = 2475.

Figure 2 shows the distribution functions $\hat{F}(.)$ for the probit and the Klein-Spady models in sample 2 using the cleaned data. 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 4 shows some summary measures of the estimated claim costs for the subsample of pensioners receiving Income Support. These estimates are constructed using expressions (10) and (14), which give quite different results because of the skewness in the lognormal distribution for C. The Klein-Spady 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. Data cleaning makes relatively little difference to the results. Even using the preferred semi-parametric estimates the average estimated claim cost for IS recipients is small, averaging around £3.59 per week in the preferred sample 2.

Figure 3 shows the empirical distribution of these estimated claim costs for the subset of pensioners within sample 2 (cleaned data) 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.

Table 4 Summary measures of estimated stigma/claim costs for Income Support recipients (£ per week)

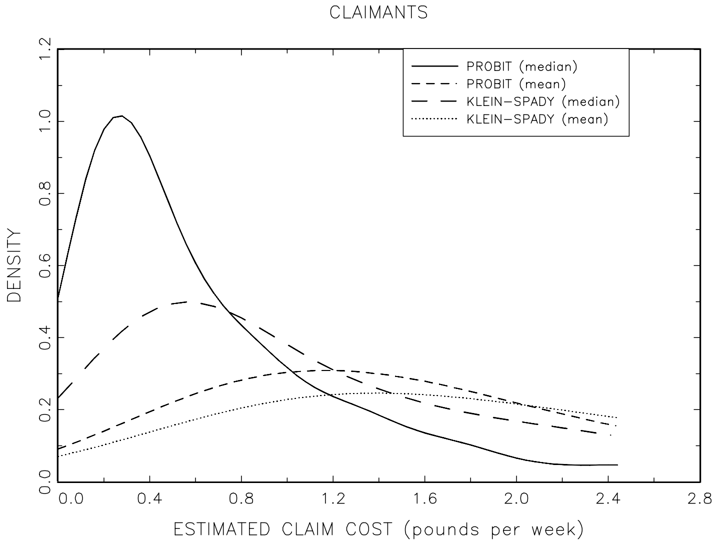
		Pre-cleaning		
		Mean	Median	Std. dev.
Probit: conditional mean method ($\widehat{C}_1$)	Sample 1[1]	1.64	1.01	1.53
	Sample 2[2]	2.89	1.75	2.65
Probit: conditional median method ($\widehat{C}_2$)	Sample 1[1]	0.08	0.02	0.15
	Sample 2[2]	0.67	0.35	0.86
Klein-Spady: conditional mean method ($\widehat{C}_1$)	Sample 1[1]	4.02	2.09	5.37
	Sample 2[2]	4.05	2.58	4.12
Klein-Spady: conditional median method ($\widehat{C}_2$)	Sample 1[1]	2.59	0.78	4.92
	Sample 2[2]	2.22	1.07	3.12

[1] $n = 1893$; [2] $n = 1983$;

		Cleaned data		
		Mean	Median	Std. dev.
Probit: conditional mean method ($\widehat{C}_1$)	Sample 1[4]	1.25	0.82	1.18
	Sample 2[5]	2.93	1.80	2.72
Probit: conditional median method ($\widehat{C}_2$)	Sample 1[4]	0.04	0.01	0.07
	Sample 2[5]	0.82	0.47	1.00
Klein-Spady: conditional mean method ($\widehat{C}_1$)	Sample 1[4]	2.07	1.25	3.40
	Sample 2[5]	3.59	2.31	3.41
Klein-Spady: conditional median method ($\widehat{C}_2$)	Sample 1[4]	0.50	0.25	0.70
	Sample 2[5]	1.73	0.98	2.05

[4] $n = 1588$; [5] $n = 1646$;

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



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 (14) and (15). 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 (11) and (14) to estimate implicit claim costs for the average 1984 pensioner claimant. Respectively, these come to £1.70 and £1.02 (updated to 1998 prices) using the conditional mean and median methods. These are comparable with our 1997-2000 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.

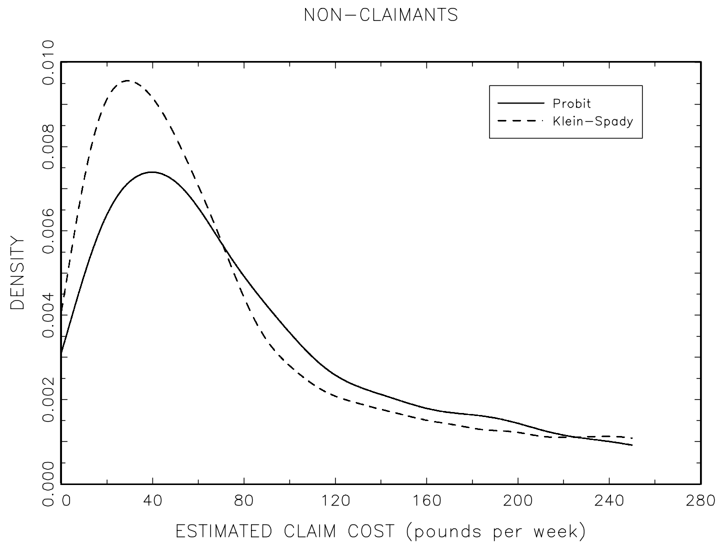
Table 5 shows the median of the estimated costs of claiming for IS non-participants. The estimates of claim costs are found to be of a smaller magnitude with the Klein-Spady model as opposed to the probit model, when the cleaned data is used and also in sample 2. These median claim costs are much larger than those found for IS recipients. 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, and for the pre-cleaned data.

Table 5 Median of estimated stigma/claim costs for Income Support non-recipients (£ per week)

		Pre-cleaning	Cleaned
Probit: $\widehat{C}_2$	Sample 1[1]	746.73	799.03
	Sample 2[2]	111.25	71.83
Klein-Spady: $\widehat{C}_2$	Sample 1[1]	107.06	102.34
	Sample 2[2]	72.14	55.09

[1] n=1036 (pre-cleaned) 1036(cleaned); [2] n=829 (pre-cleaned) 829(cleaned);

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



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:

$$FGT = \frac{E\left[S\,Q\left(Y, B, \mathbf{X}\right)\right]}{E\left(S\right)} \tag{18}$$

where:

$$Q(Y,B,\mathbf{X}) = \begin{cases} \left(1 - \frac{Y+B}{T(\mathbf{X})}\right)^2 & \text{if } Y+B \le T(\mathbf{X}) \\ 0 & \text{otherwise} \end{cases} \tag{19}$$

A baseline estimate of this measure can be computed by replacing the expectations in (18) with sample averages:

$$\widehat{FGT} = \sum_{i=1}^{n} S_i Q(Y_i, B_i, \mathbf{X}_i) \, / \sum_{i=1}^{n} S_i \tag{20}$$

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 (20) is substituted by

$$Q^*(Y,B,\widehat{C},\mathbf{X}) = \begin{cases} \left(\frac{T(\mathbf{X})-Y}{T(\mathbf{X})}\right)^2 & \text{if } Y \le T(\mathbf{X}) \\ & \text{ and } B = 0 \\ \left(\frac{T(\mathbf{X})-Y-B+\widehat{C}}{T(\mathbf{X})}\right)^2 & \text{if } Y+B-\widehat{C} \le T(\mathbf{X}) \\ & \text{ and } B > 0 \\ 0 & \text{otherwise} \end{cases} \tag{21}$$

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 (18) 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:

$$Q^{**} = \begin{cases} \left(\frac{T(\mathbf{X})-Y}{T(\mathbf{X})}\right)^2 & \text{if } Y \le T(\mathbf{X}) \\ & \text{ and } B = 0 \\ \frac{\int_{-\infty}^{\ln B} \left(\frac{T(\mathbf{X})-Y-B+C}{T(\mathbf{X})}\right)^2 dF\left(\frac{\ln C - \mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right)}{F\left(\frac{\ln(B)-\mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right)} & \text{if } Y+B \le T(\mathbf{X}) \\ & \text{ and } B > 0 \\ \frac{\int_{\ln(Y+B-T(\mathbf{X}))}^{\ln B} \left(\frac{T(\mathbf{X})-Y-B+C}{T(\mathbf{X})}\right)^2 f\left(\frac{\ln C - \mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right) d\ln C}{F\left(\frac{\ln(B)-\mathbf{Z}\boldsymbol{\alpha}}{\sigma}\right)} & \text{if } Y+B > T(\mathbf{X}), \\ & \text{ and } Y < T(\mathbf{X}) \\ 0 & \text{otherwise} \end{cases} \tag{22}$$

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 6 and 7. The effects of adjusting for claim costs are small. Depending on the sample, threshold and estimator used, measured poverty is some 4-16% higher when claim costs are taken into account. This is not negligible, but hardly a dramatic impact.

Table 6 Foster-Greer-Thorbecke poverty measures (probit model).

			Poverty line			
			1.2 M	1.1 M	M	0.9 M
Sample 1	Pre-cleaning $n = 11,524$	$\widehat{FGT}$	1.64	0.93	0.53	0.34
		$\widehat{FGT}^*$ (mean)	1.70	0.97	0.53	0.34
		$\widehat{FGT}^*$ (median)	1.64	0.93	0.53	0.34
		$\widehat{FGT}^{**}$	1.72	0.99	0.56	0.36
	Cleaned $n = 10,268$	$\widehat{FGT}$	1.42	0.76	0.40	0.25
		$\widehat{FGT}^*$ (mean)	1.47	0.79	0.41	0.25
		$\widehat{FGT}^*$ (median)	1.42	0.77	0.40	0.25
		$\widehat{FGT}^{**}$	1.48	0.81	0.43	0.26
Sample 2	Pre-cleaning $n = 11,477$	$\widehat{FGT}$	1.73	1.03	0.61	0.39
		$\widehat{FGT}^*$ (mean)	1.84	1.10	0.63	0.40
		$\widehat{FGT}^*$ (median)	1.75	1.04	0.61	0.39
		$\widehat{FGT}^{**}$	1.87	1.14	0.67	0.43
	Cleaned $n = 10,236$	$\widehat{FGT}$	1.46	0.82	0.45	0.28
		$\widehat{FGT}^*$ (mean)	1.56	0.89	0.47	0.29
		$\widehat{FGT}^*$ (median)	1.48	0.83	0.46	0.28
		$\widehat{FGT}^{**}$	1.59	0.92	0.51	0.31

Table 7 Foster-Greer-Thorbecke poverty measures (Klein-Spady model).

			Poverty line			
			1.2 M	1.1 M	M	0.9 M
Sample 1	Pre-cleaning $n = 11,524$	$\widehat{FGT}$	1.64	0.93	0.53	0.34
		$\widehat{FGT}^*$ (mean)	1.81	1.05	0.57	0.35
		$\widehat{FGT}^*$ (median)	1.75	1.01	0.55	0.35
		$\widehat{FGT}^{**}$	1.83	1.07	0.60	0.37
	Cleaned $n = 10,268$	$\widehat{FGT}$	1.42	0.76	0.40	0.25
		$\widehat{FGT}^*$ (mean)	1.50	0.81	0.41	0.25
		$\widehat{FGT}^*$ (median)	1.44	0.78	0.40	0.25
		$\widehat{FGT}^{**}$	1.50	0.83	0.44	0.26
Sample 2	Pre-cleaning $n = 11,477$	$\widehat{FGT}$	1.73	1.03	0.61	0.39
		$\widehat{FGT}^*$ (mean)	1.89	1.14	0.65	0.41
		$\widehat{FGT}^*$ (median)	1.81	1.08	0.63	0.40
		$\widehat{FGT}^{**}$	1.92	1.18	0.69	0.43
	Cleaned $n = 10,236$	$\widehat{FGT}$	1.46	0.82	0.45	0.28
		$\widehat{FGT}^*$ (mean)	1.59	0.90	0.48	0.29
		$\widehat{FGT}^*$ (median)	1.52	0.85	0.46	0.29
		$\widehat{FGT}^{**}$	1.62	0.94	0.52	0.32

7 Conclusions

This paper studies the take-up of Income Support by UK pensioners using data on the financial years 1997/8-1999/2000 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 location. 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 small for most IS recipients, typically around £2-4 per week for the average benefit recipient, and consequently the degree of measured poverty among pensioners increases by only a modest amount (up to 16% 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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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 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
Wales/ Scotland	Dummy variable = 1 if the household lives in Wales/Scotland

Appendix: Additional tables

The following tables are included to assist referees. They are not intended for publication

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

Variable	Sample mean	Probit coeff[1]	Sample mean	Probit coeff[2]
$\ln(B_1)$	2.550	0.151		
	(0.023)	(0.030)		
$\ln(B_2)$			2.545	0.261
			(0.022)	(0.027)
$\ln(B_3)$				
Single male household	0.143	0.255	0.142	-0.018
	(0.006)	(0.110)	(0.006)	(0.104)
Single female household	0.760	0.498	0.754	0.224
	(0.008)	(0.093)	(0.008)	(0.086)
Age/10	7.907	-2.493	7.899	-2.622
	(0.013)	(0.739)	(0.013)	(0.727)
$(\text{Age}/10)^2$	63.009	0.148	62.881	0.153
	(0.205)	(0.046)	(0.202)	(0.046)
Income per head	65.464	-0.017	66.705	-0.008
	(0.331)	(0.002)	(0.340)	(0.002)
Head educated past 14	0.200	-0.280	0.196	-0.307
	(0.007)	(0.064)	(0.007)	(0.063)
Disability benefit	0.380	-0.185	0.378	-0.401
	(0.009)	(0.079)	(0.009)	(0.070)
Registered Disabled	0.193	0.177	0.192	0.166
	(0.007)	(0.071)	(0.007)	(0.070)
Owner occupier	0.276	-0.549	0.278	-0.554
	(0.008)	(0.056)	(0.008)	(0.055)
Rent free	0.035	-0.393	0.034	-0.374
	(0.003)	(0.132)	(0.003)	(0.130)
Wales/ Scotland	0.178	-0.132	0.179	-0.146
	(0.007)	(0.065)	(0.007)	(0.064)
n	2929		3019	

[1]Intercept = 11.429 (std err = 2.929);
[2]Intercept = 11.633 (std err = 2.883)

Table A4 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]
$\ln(B_1)$	2.403 (0.025)	0.138 (0.032)		
$\ln(B_2)$			2.442 (0.024)	0.296 (0.030)
$\ln(B_3)$				
Single male household	0.142 (0.007)	0.123 (0.119)	0.142 (0.007)	-0.148 (0.112)
Single female household	0.755 (0.009)	0.495 (0.101)	0.748 (0.009)	0.198 (0.094)
Age/10	7.876 (0.014)	-1.919 (0.838)	7.871 (0.014)	-2.338 (0.833)
$(\text{Age}/10)^2$	62.511 (0.223)	0.112 (0.053)	62.439 (0.220)	0.134 (0.053)
Income per head	66.802 (0.394)	-0.014 (0.002)	68.097 (0.407)	-0.004 (0.002)
Head educated past 14	0.203 (0.008)	-0.305 (0.070)	0.199 (0.008)	-0.322 (0.069)
Disability benefit	0.345 (0.010)	-0.061 (0.092)	0.346 (0.010)	-0.381 (0.083)
Registered Disabled	0.175 (0.008)	0.213 (0.082)	0.175 (0.008)	0.176 (0.081)
Owner occupier	0.268 (0.009)	-0.546 (0.062)	0.270 (0.009)	-0.562 (0.061)
Rent free	0.034 (0.004)	-0.378 (0.147)	0.034 (0.004)	-0.369 (0.147)
Wales/ Scotland	0.173 (0.008)	-0.157 (0.073)	0.173 (0.008)	-0.188 (0.073)
n	2417		2475	

[1]Intercept = 9.041 (std err = 3.305);
[2]Intercept = 10.321 (std err = 3.289)