What Explains Cross Section Variation in British Disability Benefit Rolls?

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Abstract

Although we have a pretty good idea of the types of factors that have driven rapid growth in disability benefit rolls in Britain, the US and other countries over the last 30 years or so, their relative roles have not yet been quantified satisfactorily in the literature. This paper's contribution is to exploit within-country cross sectional variation to provide estimates of the long run relationships between disability benefit rolls and factors thought to play a key role in explaining them. Specifically, a simultaneous equations model is estimated, separately for male and female local area disability benefit rolls, on British Local Authority level data. For both sexes, disability benefit rolls are shown to vary positively with disability incidence and unemployment rates, and negatively with average earnings.

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1. Introduction

The number of people of working age claiming income replacement disability benefits (known as Incapacity Benefit, or IB) in Britain currently stands at 2.6 million. This figure has grown by over 300% in 30 years (see Figure 1). Similar growth has been experienced by the US and some other OECD countries. Bound and Burkhauser (1999) review the empirical literature aiming to explain this growth, primarily but not exclusively for the US, and conclude that a combination of falling demand for low skilled workers and the characteristics of the benefits themselves is to blame. Increasing prevalence or severity of disability is generally not thought to have played a significant role. Subsequent studies for the US have drawn similar conclusions (e.g. Autor and Duggan, 2003, 2006) as have studies for other countries (e.g. Aarts and de Jong, 1992; Bowitz, 1997; Riphahn, 1999; Huddleston, 2000).

McVicar (2007), however, argues that the literature – particularly for Britain but also to some extent internationally – has so far failed to provide convincing quantitative estimates of the roles played by these different factors in the growth of disability benefit rolls. The (uncontroversial) argument is that where a number of (possibly inter-related) factors have plausibly driven benefit roll growth, a properly specified multiple regression framework is required to quantify their roles. In the British case, most existing studies have been descriptive or qualitative rather than quantitative in nature (e.g. Beatty and Fothergill, 1996; Huddleston, 2000; Walker and Howard, 2000). The exceptions use old data, have omitted key variables from their empirical models, have failed to account for the possible endogeneity of right hand side variables, or some combination of the above (e.g. Disney and Webb, 1991; Bell and Smith, 2004). Widely cited US studies have also omitted key variables or at best have used poor proxies for omitted variables, e.g. omission of disability prevalence and labour market conditions by Black et al. (2002), omission of disability prevalence by Stapleton et al. (1998) and Autor and Duggan (2003). Such studies tell us little about potential relationships between the omitted variables and benefit rolls, as well as risking omitted variable bias. In the case of disability prevalence, for example, there is little in the way of compelling quantitative evidence for any country to support Bound and Burkhauser's conclusion that it has not played a significant role in driving benefit roll growth.

One constraint holding back quantitative research on the growth of disability benefit rolls is data availability, e.g. in terms of obtaining consistent series on disability prevalence. An alternative to time series estimation, however, is to exploit cross sectional variation in disability benefit rolls within countries in order to get at the long run relationships of interest. In the US, for example, disability benefit rolls are generally higher in the South than in the rest of the country. In Britain, they are highest in Wales and the North. For the US, such cross section variation has recently been exploited by Black et al. (2002) and Autor and Duggan (2003) in estimating the response of disability benefit rolls to the replacement rate. Nolan and Fitzroy (2003) do so using British Local Authority (LA) level data, but omit labour market indicators and disability prevalence from their regressions. This paper builds on these earlier contributions to estimate a more complete empirical model for local area IB rolls, separately for men and women, using 2003 British LA level data.

The remainder of this paper is set out as follows. The following section provides a motivating search theory framework for the reduced form empirical model. Section 3 briefly describes the data and sets out the empirical model. Section 4 discusses the estimation results and Section 5 concludes.

2. Motivating Framework

The motivating model for the paper is a three state search model such as that set out by Cahuc and Zylberberg (2004, p109-115). The model can be used straight 'off the shelf' if we make the simplifying assumption that a fixed proportion of the working age population are disabled (medically eligible for IB) and that it is to individuals in this fixed proportion that the model refers.[1]

We borrow Cahuc and Zylberberg's notation for the purposes of describing the model, which assumes no on-the-job search, a rate of job loss q, an offer arrival rate λ, a rate of interest r, a rate of unemployment benefit z, income from non-participation of R_I, and a wage offer distribution H(w), and can be described by the following expressions, where x denotes the reservation wage and Ω denotes the characteristics of the labour market :

$$\left\{\begin{array}{l} x(\Omega) \geq R_I \Rightarrow particpant \\ x(\Omega) \leq R_I \Rightarrow non-particpant \end{array}\right\} \tag{1}$$

$$\left\{\begin{array}{l} w > x(\Omega) \Rightarrow employed \\ w \leq x(\Omega) \Rightarrow unemployed \end{array}\right\} \tag{2}$$

$$\Omega = \Omega(H(w), z, q, \lambda, r) \tag{3}$$

It can be shown that the individual's reservation wage x is increasing in z and λ and decreasing in r and q. The choice between job searching and non-participation will also be influenced by the income flow from non-participation R_I, which is interpreted here as the IB payment rate.

Aggregating across individuals, the proportion of the local area working age population that claim IB will equal the proportion of the working age population for whom $x(\Omega) \leq R_I$, which it follows is increasing in z, λ and R_I, and decreasing in r and q. Variation between local labour markets in terms of q and λ will therefore lead to variation in local IB rolls. Although benefits are paid at a national rate, variation between local labour markets in terms of the distribution of wages will also lead to variation in local IB rolls, with individuals in high wage labour markets more likely to participate than those in low wage labour markets. Finally, because eligibility for IB depends on disability status, local IB rolls will also increase with the number of working age disabled individuals in the local area.

The following section sets out a reduced form model based on the above motivating framework, where the number of disabled individuals in each LA is proxied by the number of self-reported disabled, the wage distribution is proxied by the median full time wage, and job destruction and offer arrival rates are proxied by local unemployment rates. Our interest here is not in the parameters of the formal search model *per se*, but in the broad empirical relationships between IB rolls and the above factors. This is partly data driven, but it is also that these reduced form

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[1] We can ignore the non-disabled in this partial equilibrium set up because they will never be eligible for IB.

empirical relationships are the subject of most of the existing empirical literature on disability benefits, and it is to this literature that this paper contributes.

3. Data and Model

There are 408 LA areas in Britain. For these areas, separately by sex, the paper uses administrative data on the proportion of the working age population claiming IB, Labour Force Survey data on the proportion reporting a limiting long-standing illness or disability, the proportion without qualifications, and on economic activity by disability status. Median full time weekly earnings data are taken from the Annual Survey of Hours and Earnings. Table 1 provides descriptive statistics.

The equation of interest is Equation (4) below[2], where IB rolls are determined by the incidence of disability, the state of the labour market as proxied by the non-disabled unemployment rate, economic incentives (average wages), and an additional control for the qualifications level of the local population.[3]

Relaxing the assumption that disability incidence is fixed and endogenous, IB rolls can be allowed to be jointly determined with disability incidence as given by Equation (5). Amongst other things, this may reflect reverse causality because of justification bias where, for a given

[2] Although the log linear form of the model is the most convenient in terms of interpretation, it is possible for predicted values to fall outside the range from zero to one. Sensitivity to this is examined by estimating the model in levels with logistically transformed LHS variables. Results are available on request.
[3] This is included because of correlation between qualifications and disability status, which puts many disabled people at a multiple disadvantage in the labour market (e.g. Berthoud, 2007).

type and severity of condition, those on disability benefits are more likely to report a disability in surveys than others (e.g. see Parsons, 1982).

$$\log IB_i = \beta_0 + \beta_1 \log Dis_i + \beta_2 \log U_i + \beta_3 \log W_i + \beta_4 \log Q_i + v1_i \tag{4}$$

$$\log Dis_i = \alpha_0 + \alpha_1 \log IB_i + \alpha_2 \log U_i + \alpha_3 \log W_i + \alpha_4 \log Q_i + \alpha_5 \log Dis_{i2002} + \alpha_6 \log Dis_{i2003}^{opposite} + v2_i \tag{5}$$

IB_i denotes the 2003 proportion of the working age population claiming IB in area i, Dis_i denotes the proportion of the working age population reporting a disability and Dis_{i2002} the 2002 value, U_i is the non-disabled unemployment rate, W_i is median weekly earnings and Q_i is the proportion of the working age population without qualifications. The 'opposite' superscript denotes the value for the opposite sex, used to instrument same sex disability incidence. This instrumenting strategy has not previously been used in the literature on disability benefits[4], but is intuitively attractive , and passes the usual tests (see Tables 2 and 3 for results).

We can also allow for the non-disabled unemployment rate to be endogenous to the model – perhaps because of common omitted factors – according to Equation (6). Again we use the opposite sex value and same sex lagged value instrumenting strategy. Some previous studies (e.g. Autor and Duggan, 2003) have alternatively constructed plausibly exogenous proxies for

[4] One reason for this may be that existing studies have tended to focus on data for males only or, where data covering males and females has been available, to not distinguish claimants by sex.

labour demand. Studies using British data have tended to do neither (e.g. Disney and Webb, 1991; Faggio and Nickell, 2005).

$$\log U_i = \gamma_0 + \gamma_1 \log W_i + \gamma_2 \log Q_i + \gamma_3 \log Dis_i + \gamma_4 \log U_{i2002} + \gamma_5 \log U_i^{opposite} + v3_i \qquad (6)$$

Finally, we allow local area average earnings to be endogenously determined – following evidence presented by Black et al. (2002), Autor and Duggan (2003), and Faggio and Nickell (2005) – according to Equation (7).

$$\log W_i = \delta_0 + \delta_1 \log U_i + \delta_2 \log Q_i + \delta_3 \log Dis_i + \delta_4 \log W_{i2002} + \delta_5 \log W_{i2003}^{opposite} + v4_i \qquad (7)$$

Tables 2 and 3 report OLS and IV estimates for Equation (4), with appropriate specification tests, for men and women respectively. Two Stage Least Squares (2SLS) results are presented for Equation (4) paired with each of the other equations in turn, and Three Stage Least Squares (3SLS) results are presented for the whole system (females) or appropriate sub-system (males). In each case, Equation (4) is identified by exclusion of the lagged and opposite sex terms in the other equations. The full system is similarly identified.

4. Results and Discussion

OLS estimation of Equation (4) suggests that local area IB rolls are positively correlated with local area disability incidence, unemployment rates, and the proportion of the population with no qualifications for both sexes. They are negatively correlated with local area average earnings.

For men, testing the exogeneity of disability incidence gives ambiguous results, with the relevant 2SLS estimates (IV1) qualitatively similar to the OLS estimates with the exception of a larger elasticity with respect to disability incidence (the OLS estimate is downward biased). For women, disability incidence is unambiguously endogenous, and the 2SLS estimates are somewhat different to the OLS estimates, with a considerably higher elasticity with respect to disability incidence, and lower elasticities with respect to the other observed variables. The instruments are shown to be relevant and validly excluded from Equation (4) for both sexes, although for women this requires the use of 2002 male disability incidence rather than 2003 male disability incidence as an instrument.

The unemployment rate is unambiguiously endogenous for men but ambiguously endogenous for women. For both, the 2SLS estimates are close to the OLS estimates, although with higher elasticity with respect to the unemployment rate (again the OLS estimate is downward biased).

Test results suggest that average earnings can be treated as exogenous for men – reflected in very similar OLS and 2SLS estimates – but not for women, for whom the lagged same sex instrument is replaced by a lagged opposite sex instrument because of a failed validity test. The resulting 2SLS estimates are similar to the OLS estimates, with the exception of a higher elasticity on average earnings (the OLS estimates are biased towards zero).

Finally, we estimate the full systems for men (Equations 4-6) and women (Equations 4-7) by 3SLS.[5] These are the preferred specifications, with estimates qualitatively similar to those obtained by OLS in each case, and it is to these estimates that the remainder of this section refers.

For both sexes the estimated elasticity of IB rolls to disability incidence is close to one. Ignoring differences in disability incidence, or only controlling for them with mortality rates, other health proxies, or fixed effects (e.g. Autor and Duggan, 2003; Nolan and Fitzroy, 2003; Disney and Webb, 1991; Black et al., 2002) is likely to obscure what appears to be an important driver of IB roll variation across space, and by extension over time, and may lead to omitted variable bias. Even if included, ignoring the potential endogeneity of disability incidence (e.g. Faggio and Nickell, 2005) is likely to lead to a biased estimate of its role in determining disability benefit rolls. For these data the bias is towards zero, which is consistent with subjective disability status measuring 'true disability' with error (see e.g. Bound, 1991).

For both sexes the estimated elasticity of IB rolls to the local unemployment rate is around half. Again, omitting such a control for the state of the labour market (e.g. Nolan and Fitzroy, 2003; Faggio and Nickel, 2005) obscures a relationship of interest and is likely to lead to omitted variable bias in other variables present in the model. If included, ignoring the potential endogeneity of such a measure (e.g. Disney and Webb, 1991) is again likely to lead to biased elasticity estimates. Disney and Webb's (1991) main conclusion, however, is correct: disability benefit rolls are influenced by the state of the labour market.

[5] For women, we drop the lagged same sex instrument for the unemployment rate because it is weak in the full system.

For both sexes the estimated elasticity of IB rolls to local area average earnings is around minus one. This is larger in absolute size than Black et al.'s estimate of around minus one half for the US but close to Faggio and Nickell's (2005) IV estimate for British men. It is also broadly consistent with the estimated replacement rate elasticity of Aarts and de Jong (1992) for the Netherlands. Evidently, economic incentives – what IB claimants can potentially obtain in the local labour market – play a role in determining disability benefit rolls in Britain just as they do in the US and elsewhere.[6]

5. Conclusions

This paper contributes to the literature by providing quantitative estimates of factors that influence the cross sectional variation in disability benefit rolls within Britain. Few papers have previously (quantitatively) examined such cross sectional differences, whether for Britain or elsewhere, and none have previously considered male and female benefit rolls separately. The treatment of disability incidence – shown here to be a key determinant of cross sectional differences in disability benefit rolls – has also been somewhat patchy in the empirical disability benefits literature. The results presented here, consistent with earlier evidence for the US and the Netherlands, suggest that economic incentives, in the form of available wages, also matter. The paper also confirms the main conclusion of Disney and Webb's (1991) paper: that local unemployment rates are a significant determinant of disability benefit rolls.

[6] Note that IB is paid at the same national rate to all recipients (although there is some variation by duration of claim) regardless of local area average earnings or an individual's previous earnings.

Of course, one should not claim too much from a single cross-section of local area level data. Further work might conduct a similar exercise for other countries or, within the UK, for the much smaller spatial units – not subject to the same sample reporting restrictions – available from the 2001 Census. Perhaps more important will be analyses of detailed longitudinal data to more concretely pin down the causal relationships of interest.

Even from the evidence presented here, however, some support can be found for recent reforms of IB by the British Government, e.g. changes to benefit rates and ‘back to work’ bonuses, under the ‘Pathways to Work’ package. Inasmuch as reducing the numbers of IB claimants is itself a direct policy aim, as is clearly suggested by the existence of a target to reduce the number of IB claimants by 1 million over the next 10 years (see Freud, 2007), such measures to influence the economic incentives of claimants appear *ex ante* to be on the right track.

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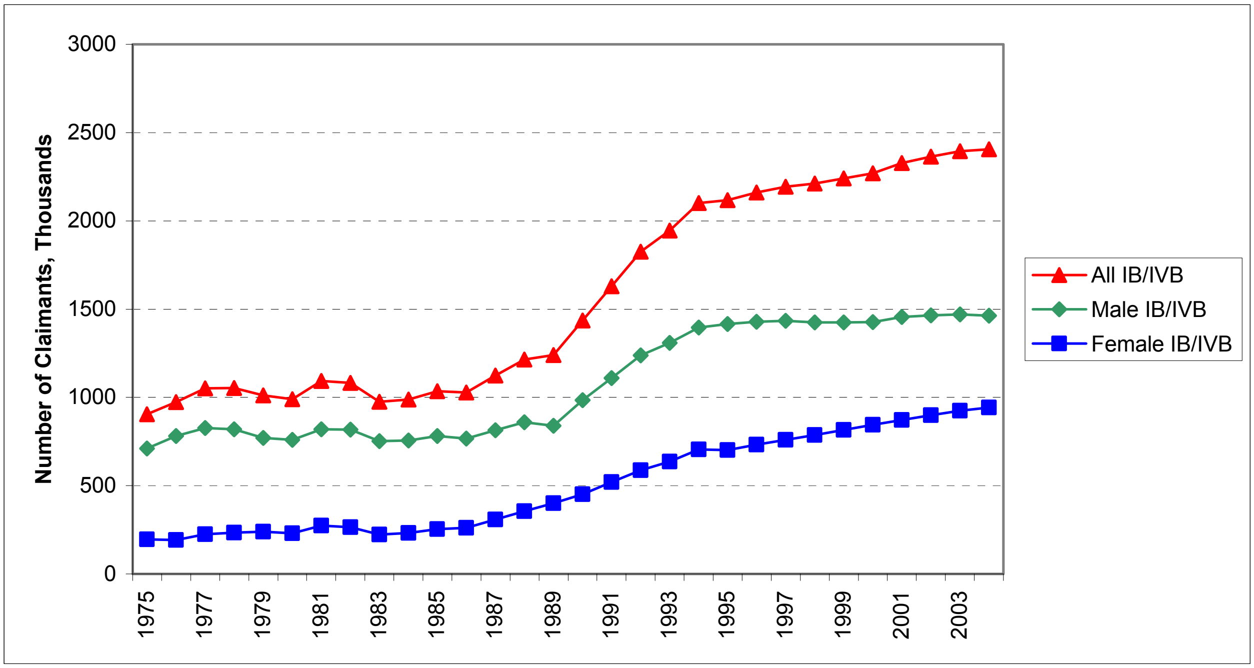
Figure 1: Number of Claimants of IB (IVB), Working Age, 1975-2004



Source: Social Security Statistics (recently Department for Work and Pensions), various years. Notes: 'Working age' is taken as 18-64 for males and 18-59 for females. Figures refer to recipients of Sickness, Invalidity (pre 1995) and Incapacity (post 1995) Benefit, Severe Disablement Allowance, National Insurance credits and Income Support for reasons of disability, collectively known as IB claimants, at end of statistical year.

Table 1: Descriptive Statistics

	Mean, Male (Female)	Max, Male (Female)	Min, Male (Female)	St. Dev, Male (Female)
IB Claimant Rate	.079 (.054)	.297 (.192)	.010 (.007)	.041 (.027)
Disabled Rate	.191 (.187)	.427 (.491)	.052 (.067)	.050 (.049)
Non-disabled Unemployment Rate	.043 (.039)	.139 (.138)	0 (0)	.027 (.024)
Median Weekly Earnings, £	455 (338)	1008 (568)	324 (203)	78.6 (59.1)
No Qualifications Rate	.135 (.158)	.329 (.380)	.033 (.027)	.051 (.060)

Table 2: OLS and IV Estimates of Equation (4), Males

	OLS	IV1 (2SLS, endog Dis_i)	IV2 (2SLS, endog U_i)	IV3 (2SLS, endog W_i)	IV4 (3SLS, endog Dis_i, U_i)
Disabled Rate	.753* (.073)	1.09* (.130)	.692* (.125)	.741* (.073)	.931* (.190)
Non-disabled Unemployment Rate	.170* (.034)	.158* (.035)	.432* (.086)	.175* (.035)	.395* (.085)
Median Weekly Earnings, £	-1.02* (.174)	-.765* (.188)	-1.03* (.212)	-1.07* (.194)	-.850* (.204)
No Qualification Rate	.209* (.054)	.177* (.054)	.097 (.077)	.186* (.054)	.090 (.067)
Constant	5.82* (.966)	4.75* (1.00)	6.42* (1.15)	6.11* (1.08)	5.58* (1.07)
R^2	.623	.601	.552	.623	.566
Prob>F (model significance)	.000	.000	.000	.000	.000
Observations	370	370	285	370	285
Prob>F (instrument relevance)		.000	.000	.000	
Hansen J-statistic p-value (instrument validity)		.987	.867	.510	
Hausman test p-value (exogeneity)		.270	.000	.972	
Davidson-MacKinnon Prob>F (exogeneity)		.000	.000	.616	

Notes: Robust standard errors in parentheses. * statistically significant at 5%. The number of observations drops for IV2 because of missing unemployment rate data for 2002.

Table 3: OLS and IV Estimates of Equation (4), Females

	OLS	IV1 (2SLS, endog Dis_i)	IV2 (2SLS, endog U_i)	IV3 (2SLS, endog W_i)	IV4 (3SLS, endog Dis_i, U_i, W_i)
Disabled Rate	.505* (.084)	1.44* (.181)	.428* (.124)	.394* (.091)	.877* (.325)
Non-disabled Unemployment Rate	.139* (.031)	.091* (.040)	.309* (.138)	.166* (.032)	.578* (.169)
Median Weekly Earnings, £	-.620* (.154)	-.227 (.178)	-.575* (.171)	-1.23* (.231)	-1.07* (.358)
No Qualification Rate	.421* (.067)	.203* (.086)	.405* (.085)	.341* (.068)	.075 (.091)
Constant	2.71* (.852)	1.46* (.968)	2.85* (1.08)	5.98* (1.26)	6.73* (2.02)
R^2	.515	.340	.481	.486	.217
Prob>F (model significance)	.000	.000	.000	.000	.000
Observations	360	359	256	358	336
Prob>F (instrument relevance)		.000	.000	.000	
Hansen J-statistic p-value (instrument validity)		.091	.379	.341	
Hausman test p-value (exogeneity)		.000	.004	.003	
Davidson-MacKinnon Prob>F (exogeneity)		.000	.250	.004	

Notes: Robust standard errors in parentheses. * statistically significant at 5%. The number of observations drops for IV2 because of missing unemployment rate data for 2002.