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A dynamic analysis of disability and labour force participation in Ireland 1995–2000

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Summary

This paper aims to analyse the effect of disability on participation in the labour force, using the Irish component of the European Community Household Panel Survey 1995–2000. A range of panel models are considered, but to allow for any unobserved influences or state dependence in labour force participation, our preferred model is a dynamic panel model. We show how the estimates of current disability are changed once we control for the effect of past disability and previous participation. We compare base estimates of disability with those controlling for unobserved heterogeneity and past participation. The results suggest that the base effect of disability is overestimated by between 40–60% for men and by 5–10% for women. Copyright © 2005 John Wiley & Sons, Ltd.

Keywords disability; labour force participation; static and dynamic panel models

Introduction

People with disabilities face many barriers to full participation in society, not least in the labour market, and the extent and nature of participation in the labour market has a multitude of direct and indirect effects on their living standards and quality of life. In studying the effect of disability on labour force participation, we are faced with a variety of analytical challenges, such as the effect of unobserved characteristics of disabled individuals and the effect of their past participation in the labour market. This paper uses panel data methods to control for these factors and we estimate the impact of disability on participation, controlling for unobserved heterogeneity and past participation.

The focus of previous policy for disabled people has been on the provision of services, whereas more recently, there is a campaign for civil rights and the provision of legislation for equality and full participation.[a] Employers and policy makers are therefore interested in whether or not disability has an effect on participation. In estimating the effect of disability on labour force participation, there are two main sources of bias that may arise, from measurement error and endogeneity. Previous research [1,2] has already set out the main issues involved and we now review these to emphasise the motivation for this paper. Firstly, there may be problems with the measurement of the disability variable and lack of comparability across individuals may lead to underestimates of the effect of disability (via classical measurement error). On the other hand, economic or psychological

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incentives may affect an individual's response to questions on disability, leading to differential measurement error within the self-reported measure of disability in the participation model. Secondly, participation and disability may be endogenously related because of direct effects of disability on participation and vice versa. In addition, there may be unobservables that influence both disability and participation outcomes, for example through an individual's time preference or previous investments in human or health capital. In this paper, we focus on controlling for the latter, referred to by Lindeboom and Kerkhofs [2] as 'classical endogeneity'.[b]

Our data offer the possibility of analysing the relationship between disability and labour force participation over a significant period rather than just at a point in time, and allow us to use panel data techniques in our estimation. Using panel data, we capture the effects of variables that are particular to an individual and are constant over time. Labour force participation may also be influenced by past participation, where non-participants in the previous year may be less likely to participate in the current year. Although this may be true for all individuals, it may also be a specific characteristic of disabled people and lead to an incorrect interpretation of the disability effect. It may be that disability reduces the probability of previous participation, and therefore indirectly influences current participation. Using panel data, we can incorporate this state dependence effect and re-estimate the effect of disability on participation.

More recently, Lindeboom and Kerkhofs [2] also include the effect of past labour market outcomes on current health in their retirement model. They find that for elderly people, working in the previous period slightly decreases the value of health. They estimate a multinomial logit model, to facilitate the three different labour market states compared to working, available to individuals nearing retirement age in the Netherlands. Although they only have two waves of panel data, by using information on previous labour market history, they specify an equation for initial participation and estimate the probability of working initially. This is included into the overall likelihood function from which unobserved effects are integrated out. They find that the effects of health are exaggerated for elderly people in a simple multinomial model, compared to their preferred model.

The contribution of this paper to the literature on disability and participation is in the way the dynamics are estimated and in particular the approach used to deal with initial conditions. Bound *et al*. [4] analysed the dynamic relationship between health and labour force transitions, and examined how the timing of health shocks affects labour force behaviour but noted that to credibly control initial conditions is a difficult task. We follow the Wooldridge [5] approach to control for unobserved effects that may be correlated with disability and we discuss this further in the section following the subsequent section. This is different from the approach to [2] mainly because we use six waves of panel data and can therefore identify the effect of past participation within a less complicated model. The main focus in this paper is to model two labour market outcomes – participation and non-participation – and hence we concentrate on a binary response variable. In contrast to [2,4], we follow an approach by Wooldridge [5] that allows us to avoid specifying a distribution for the initial participation. The likelihood function from our approach is easier to estimate and serves the same purpose in terms of looking at the effect of unobserved heterogeneity. Our findings using Irish data are similar to previous international research; reported disability status overestimates the effect of disability on participation. In addition, we show exactly how much unobserved heterogeneity contributes to variation in participation and how this changes the effect of disability. Finally, we show the effect of past disability (via its effect on previous participation), on current labour force participation.

Data

The data on disability and labour force participation in Ireland are from the Living in Ireland Survey 1995–2000.[c] The Living in Ireland Survey is the Irish component of the European Community Household Panel, conducted by the ESRI for Eurostat. Within the sample there is considerable attrition over the period with 7254 individuals responding in 1995 and only 3670 of these still present by 2000. We present the composition of the sample at each wave in Table 1 and return to the potential effects of this attrition in the section following the subsequent section. We wish to focus

Table 1. Sample size and composition at each wave, age 15–64, Living in Ireland Survey 1995–2000

	1995	1996	1997	1998	1999	2000
Men	50.4	50.5	50.4	49.8	49.9	49.1
Women	49.6	49.5	49.6	50.2	50.1	50.9
Age 15–24	24.9	24.7	24.2	23.7	22.8	23.1
24–34	20.5	20.2	20.3	20.5	20.0	18.7
35–44	20.6	20.7	21.1	20.9	21.4	21.3
45–54	19.1	19.4	19.3	19.7	19.8	19.5
55–65	14.8	14.9	15.0	15.2	15.9	17.4
Education						
Primary	26.9	26.3	26.2	24.6	23.8	21.8
Secondary	59.8	60.7	60.7	58.7	58.3	60.7
Third level	13.2	13.1	13.1	16.6	17.9	17.6
Married	59.1	58.7	59.2	58.5	58.6	56.9
N	7254	6337	5782	5273	4482	3670

on individuals of working age, hence we exclude those aged 65 and over.

In the Living in Ireland Survey, detailed information on current labour force status was obtained. For current purposes this allows us to distinguish between those who were at work, or unemployed but seeking work – who we will count as active in the labour force – and all others, whom we will count as inactive. The percentage of those unemployed but seeking work is quite low ranging from 7.5% in 1995 to 2.8% in 2000, giving a panel average of 5.1%. For this reason, we do not include them as a separate category in our dependent variable. Only 2.2% of the panel is retired before the age of 65, with more men than women taking early retirement. For those who had a disability in the previous year, 1% changes from employment to retirement in the current year, and only 0.5% go from non-participation into retirement. Of all those currently with a disability, 2% of men leave employment for retirement and 4% retire following a spell of non-participation. While it would be interesting to analyse the effect of disability on early retirement, again the sample size does not allow such investigation. A more detailed survey of disability and retirement of older workers in Ireland would provide better data for this purpose.

A measure of disability can also be constructed from the Living in Ireland Survey on the basis of individual responses to the following question:

'Do you have any chronic, physical or mental health problem, illness or disability?'

It may well be, that not only the presence of such an illness or disability but also the extent to which it limits or restricts a person, is important. To capture this, we use responses to a follow-up question concerning the impact of the disability to distinguish

(a) those reporting a chronic illness or disability and saying that it limits them severely in their daily activities;
(b) those who report a chronic illness or disability and saying it limits them to some extent, and
(c) those who report such a condition but say it does not limit them at all in their daily activities.

We should note that employers in Ireland as in many other industrialised countries are obliged by law to make 'reasonable accommodation' for those affected by disability, by changes in the work environment or in the way a job is performed to enable a person with a disability to fully do a job and enjoy equal employment opportunities. For this reason, in the survey a person may respond as not limited in daily activities, but without adaptation it is possible that they should be classified as severely limited. The extent to which respondents say they are limited relates to their daily activities rather than work, but similar measures have been shown to have significant discriminatory power in

Table 2. Labour force status by level of restriction for those with chronic illness or disability, age 15–64, Living in Ireland Survey 1995–2000

	Severe limitation	Some limitation	No limitation	No chronic illness or disability
Men				
Participation	34.92	58.02	81.45	91.59
Non-participation	65.08	41.98	18.55	8.41
N	189	655	318	6026
Women				
Participation	13.82	31.82	44.65	55.15
Non-participation	86.18	68.18	55.35	44.85
N	123	707	318	6522

terms of labour force participation in research elsewhere (e.g. [6]). Furthermore, as Table 2 shows there are different rates of participation for each sub-group, so it is important that we distinguish between the different levels of disability, in our analysis of labour force participation.

The effects of disability on labour force participation may differ among individuals, depending on other characteristics, for example age or education. Since disability may be correlated with other variables, we include measures of age, education, region, unearned income, age of youngest child and marital status. These variables are defined in detail in Table 3 and summary statistics are provided in Table 4. The youngest individuals in this sample are aged 16 and the number of observations of males and females are 7188 and 7670, respectively.

Model

A general model of labour force participation and disability may be constructed as follows:

$$y_{it}^{*} = b_0 + b_1 y_{it-1} + b_2 D_{it} + b_3 D_{it-1} + b_4 z_{it} + \alpha_i + \varepsilon_{it} \quad (1)$$

where y_{it} is the observed indicator of labour force participation, y_{it}^{*} is the underlying construct generating y_{it} and z_{it} represents a range of other variables. We also include lagged values of disability into our model, (D_{it-1}) and this allows us to distinguish the effects on participation of those who have a longer-term disability from those who have just acquired their disability. The individual time invariant unobserved effect is captured by α_i. In order to distinguish between the two effects – unobserved individual effects and past participation – we include a lagged dependent variable into the model.

The empirical model is motivated by a lifecycle model where the choice between consumption and leisure is considered as a lifetime decision, and we assume that individuals maximise their expected utility over their lifetime (following [4]). Our main aim is to concentrate on how the disability effect changes once we allow for unobserved individual effects and state dependence in labour force participation. For individuals who have different expectations about future disability depending on the duration of their disability, those with previous disability that is expected to persist are less likely to participate in the future. Since disability may be expected to reduce wages and increase disutility of work we would expect current disability to be negatively related to current participation. The effect is reinforced if disability increases access to unearned income.[d] We are also interested in the effects of lagged disability conditional on current disability. People who are persistently with a disability may be less likely to recover; therefore, we might expect differences in the behaviour of two individuals both of whom are disabled today if their previous disability status was different. Lagged effects may also be significant if transition takes time or if there is state dependence in unemployment. In this case we might expect to see different behaviour across two individuals neither of whom are disabled today if past disability had caused one of them to leave the

Table 3. Variable definitions for dependent and independent variables

Variable	Definition
LFP	= 1 if participating in the labour market, = 0 otherwise
Disabled with severe limitation	= 1 if disabled and severely limited in daily activities, = 0 otherwise
Disabled with some limitation	= 1 if disabled and limited to some extent in daily activities, = 0 otherwise
Disabled with no limitation	= 1 if disabled and not limited in daily activities, = 0 otherwise
	(Base category=no disability)
Age 15–24	= 1 if aged 15–24 years, = 0 otherwise
Age 25–34	= 1 if aged 25–34 years, = 0 otherwise
Age 35–44	= 1 if aged 35–44 years, = 0 otherwise
Age 45–54	= 1 if aged 45–54 years, = 0 otherwise
	(Base category=aged 55–64 years)
BMW	= 1 if living in border, midlands, west region, = 0 otherwise
	(Base category=rest of country)
Secondary education	= 1 if highest level of education completed is secondary, = 0 otherwise
Third level education	= 1 if highest level of education completed is third level, = 0 otherwise
	(Base category=no qualifications or highest level of education completed is primary)
Married	= 1 if married or living with a partner, = 0 otherwise
Age youngest child < 4	= 1 if age of youngest child is less than 4, = 0 otherwise
Age youngest child > = 4 and < 12	= 1 if age of youngest child is greater than or equal to 4 and less than 12, = 0 otherwise
Age youngest child > = 12 and < 18	= 1 if age of youngest child is greater than or equal to 12 and less than 18, = 0 otherwise
	(Base category=no children)
Unearned income	=Net household income − net individual disposable income
	(Net individual disposable income includes net incomes from work, social welfare payments and child benefit. Net household income aggregates individual data to household level)

Note: The regional classifications are based on the NUTS (Nomenclature of Territorial Units) classification used by Eurostat.

labour force in the past. We test this in our paper by explicitly modelling state-dependence in labour force participation and observing the resulting effect on lagged disability.

To provide some baseline estimates of disability we firstly estimate a static pooled model assuming that the errors are independent over time and uncorrelated with the explanatory variables. This model assumes that disability is exogenous (we relax this assumption later on) and provides us with base estimates, with which we can compare results from models that incorporate unobserved heterogeneity and state dependence. For notational purposes, we let x_{it} include disability, lagged disability and other variables, for the remainder of the paper. The log likelihood function for the pooled panel data is similar to that of the cross-sectional probit:

$$\log L(\beta) = \sum_{i=1}^{N}\sum_{t=1}^{T} y_{it} \log F(x'_{it}\beta) + \sum_{i=1}^{N}\sum_{t=1}^{T} (1 - y_{it}) \log(1 - F(x'_{it}\beta)) \quad (2)$$

and maximising this across all i with respect to β, we obtain the pooled probit estimator. The standard errors have been adjusted to account for clustering at the level of the individual.

Table 4. Summary statistics for all variables

Variable	Percentage of sample in each category: Men	Women
LFP	86.6	51.9
Disabled with severe limitation	2.6	1.6
Disabled with some limitation	9.1	9.2
Disabled with no limitation	4.4	4.1
No disability	83.8	85.0
Age 15–24	12.3	10.1
Age 25–34	16.4	17.2
Age 35–44	26.2	27.1
Age 45–54	24.4	25.6
Age 55–64	20.7	20.0
BMW	24.7	21.9
Secondary education	51.8	59.0
Third level education	16.7	13.3
No education or primary only	31.4	27.6
Married	68.7	73.3
Age youngest child < 4	12.5	13.3
Age youngest child $>= 4$ and <12	21.3	24.5
Age youngest child $>= 12$ and <18	15.2	17.7
Unearned income	228.64	389.5
	(240.13)	(307.7)
N	7188	7670

Note: For unearned income we present the mean and standard deviation (in parentheses).

While this model provides us with base estimates of disability, it does not allow us to answer two important questions. The first interesting question is whether or not the control variables appropriately account for any unobserved characteristics of disabled people that also influence their labour force participation decision? If this were not true, we would expect that the actual effect of current disability should be lower.

The second question is whether or not past disability affects current participation directly, or does it work through a separate channel by negatively affecting past participation? If so, we would expect to see that past participation influences current participation, and the effect of past disability should disappear. This would suggest that past disability still does have an effect on current participation, but does so by (a) directly influencing past participation and therefore, (b) indirectly affecting current participation.

To allow for these effects we estimate a dynamic model of participation that incorporates both past participation and unobserved effects. In general terms the following likelihood is derived and maximised;

$$\begin{aligned} f(y_{i0},\ldots,y_{iT} \mid x_{i0},\ldots,x_{iT},\beta) &= \int_{-\infty}^{\infty} f(y_{i0},\ldots,y_{iT} \mid x_{i1},\ldots,x_{iT},\alpha_i,\beta) \\ &\quad \times f(\alpha_i \mid x_i)\,\mathrm{d}\alpha_i \\ &= \int_{-\infty}^{\infty} \left[\prod_{t=1}^{T} f(y_{it} \mid y_{i,t-1},x_{it},\alpha_i,\beta) \right] \\ &\quad \times f(y_{i0} \mid x_{i0},\alpha_i,\beta) f(\alpha_i \mid x_i)\,\mathrm{d}\alpha_i \end{aligned} \tag{3}$$

In a dynamic model, we observe individuals at some time after the start of the process, and most likely the initial state is not randomly assigned to the individual. If the initial state and unobserved effect are correlated we need to specify $f(y_{i0} \mid x_i, \alpha_i)$ – known as the initial conditions problem. Heckman [7] suggests approximating $f(y_{i0} \mid x_i, \alpha_i)$ and then specifying $f(\alpha_i \mid x_i)$. Then $f(y_{i0}, \ldots, y_{iT} \mid x_i)$ is obtained by integrating out the unobserved effect. The main difficulty in this approach is in specifying the distribution of initial participation. We therefore follow an alternative approach suggested by Wooldridge [5] where we consider:

$$f(y_{i1}, \ldots, y_{iT} \mid y_{i0}, x_i) = \int_{-\infty}^{\infty} f(y_{i1}, \ldots, y_{iT} \mid y_{i0}, x_i, \alpha_i) \times f(\alpha_i \mid y_{i0}, x_i)\, \mathrm{d}\alpha_i \tag{4}$$

To control for correlated unobserved heterogeneity we follow [8,5] and specify the distribution of the unobserved effect conditional on the initial value y_{i0} and the time-averages of any potentially endogenous variables:

$$\alpha_i = \alpha_0 + \alpha_1' y_{i0} + \alpha_2' \bar{x}_i + a_i \tag{5}$$

The estimate of α_1 is of interest as it shows the direction of the relationship between the unobserved effect and the initial value of labour force participation. The relative importance of the unobserved effect in the error variance of the labour force participation equation is measured as $\rho = \sigma_a^2/(1+\sigma_a^2)$. This is also the correlation between the composite latent error $(\alpha_i + \varepsilon_{it})$ across any two time periods.

The likelihood function is now:

$$\int_{-\infty}^{\infty} \left[\prod_{t=1}^{5} f(y_{it} \mid y_{i,t-1}, x_{it}, \alpha_i, \beta) \right] \times f(\alpha_i \mid y_{i0}, x_{it}, \beta)\, \mathrm{d}\alpha_i \tag{6}$$

where $f(\alpha \mid y_{i0}, x_{it}, \beta) = \Phi(\alpha_1 y_{i0} + \alpha_2 \bar{x}_i, \sigma_a^2)$ if $y_{it} = 1$.

In this model of labour force participation, the individual effects are assumed to be random draws from a population, but correlated with the explanatory variables. We estimate a dynamic random effects probit model and maximise this likelihood function with respect to β and σ_α^2. This model assumes that the errors can now be correlated over time through the unobserved effect. The explanatory variables are assumed to be strictly exogenous, and are uncorrelated with the error term, ε_{it}, for each individual. The advantage of using this model over the pooled static model is that we can now estimate parameters with greater efficiency. While the pooled model would allow us to obtain consistent estimates of these parameters, it is inefficient relative to our full conditional maximum likelihood model. Furthermore, the pooled model does not allow for correlation between the unobserved effect and explanatory variables.

The means of variables are added as a set of controls for unobserved heterogeneity and we are now estimating the effects of changing explanatory variables but holding the average fixed. However, we should note that in this model, it is only possible to identify the effect of time-constant explanatory variables if we assume that the unobserved effect is partially uncorrelated with the time constant variable, where the coefficient for the correlated random effect part of that variable is zero.

In the pooled probit model we obtained estimates of β/σ_u and because the total error variance was normalised to 1, the estimated βs were population-averaged parameters by default. However, the random effects model parameter estimates will only be the same as those from the pooled model when $\sigma_\alpha^2 = 0$. Therefore, we need to rescale the βs that are estimated from the model. This is achieved by dividing the parameter estimates from the random effects model by $\sqrt{(1+\sigma_a^2)}$.

The dynamic random effects probit model relies on the assumption of strict exogeneity of the explanatory variables (x_i) conditional on α_i:

$$P(y_{it} = 1 \mid x_i, y_{it-1}, \ldots, y_{i0}, \alpha_i) = P(y_{it} = 1 \mid x_{it}, y_{it-1}, \alpha_i) \tag{7}$$

This means, that conditional on participation in the previous year and conditional on the unobserved individual effect, participation in the current year should not be related to any explanatory variable in past or future years. However, in our dynamic model, misspecification may arise from feedback effects from current labour force participation to future disability. We tested for exogeneity of the three limitation variables, by including future values of disability into the pooled probit model, (following [5]). If the current disability variables are strictly exogenous, we should find the future values to be insignificant. We found that severe and some limitations are significant, meaning that these two variables are subject to feedback effects in the model for men. In that case, we should not rely on the results of the dynamic random effects model, as the assumption of strict exogeneity has been violated. Using a pooled dynamic probit model we can obtain consistent (yet inefficient) estimates and in that sense is more reliable than the random effects model, [9]. The pooled probit model with time averages only requires contemporaneous exogeneity, i.e. it only restricts the relationship between the disturbance and explanatory variables in the same time period. The pooled probit model does not rely on the strict exogeneity assumption, and so allows us to estimate a dynamic model of participation controlling for correlated heterogeneity, providing consistent but inefficient estimates [9].

Results

Descriptive statistics

Some descriptive statistics on the labour market transitions for people with and without disabilities are a useful starting point. This data provides us with an idea of the basic effect of disability on labour force participation and give us further insight to the results of our model later on. In Table 5, we show the transitions in participation for people with disabilities. Firstly, we note that of those who have a disability at any year, 73% remain so in the following year. This means that approximately one quarter of all individuals recover from their disability, and so it is of interest to see if there is a lagged effect of disability on their current labour force participation. Similarly, the participation of those who do not recover is of interest. Furthermore, 6% of all men and women have a new disability each year, and again we would like to observe if this affects their current participation status. Within the group that do not recover there are also changes in the severity of their disability, so in our model we focus on the three categories of severe, some and no restrictions in daily activities.

Table 5. Transitions in disability status, age 15–64, living in Ireland Survey 1995–2000

	No disability t (%)	Disability t (%)
No disability $t-1$	94.0	6.0
Disability $t-1$	27.2	72.8

Static pooled probit model

Using the Living in Ireland Survey 1995–2000, we estimate a range of panel models to capture the effect of disability on participation. The main variables of interest are, disability and the associated limitations in daily activities, but we also control for other factors that may be correlated with disability, as mentioned earlier. In addition, it is likely that past disability has a direct effect on current participation, so we include lagged variables for the three types of disability. Pooling all available data for the years 1995–2000, and estimating a standard probit model, we obtain estimates from the pooled balanced sample.[e] We present results from this pooled static model in Table 6, Columns 1 and 4, for men and women, respectively. These results are presented as parameter coefficients, but we will later discuss some of the main results in terms of percentage effects.

The effects of current disability are quite high for both men and women, reducing the probability of current labour force participation significantly. At a first glance, disability has a greater negative effect on the labour force participation probability of men, compared to women. Although the effect of a severely limiting disability is less for women than men, it is still substantial. In the case of men, even those with no limitations have a slight reduction in the probability of participation. For women, we see that the probability of participation for those with no limitations, is not significantly different from women with no disability. The gap between the effects of severe and some limitations is quite large for men and even more pronounced for women, suggesting that severe disability has a more negative effect on women's participation. Past disability, in the previous year, also has a substantial effect on current participation, and is not much lower than the effect of current disability. The equality of current and past disability is also implicit in the results presented by Au *et al.* [10, Table 10]. This applies in the case of severe and some limitations, for both men and women. Similar to current disability and severe limitations, we see that individuals who previously had a severely limiting disability have a much lower probability of current participation, compared to those with no previous disability.

Table 6. Panel model results

	Men (coefficients)			Women (coefficients)		
	[1] Pooled static	[2] Random effects dynamic (re-scaled)	[3] Pooled dynamic	[4] Pooled	[5] Random effects dynamic (re-scaled)	[6] Pooled dynamic
Lag LFP		0.7511**	1.687**		0.7494**	1.7974**
		(0.1194)	(0.0918)		(0.0835)	(0.0623)
Disabled with severe limitation	−1.2368**	−0.6639**	−0.5653**	−0.9173**	−0.8256**	−1.1359**
	(0.1314)	(0.2653)	(0.2218)	(0.1736)	(0.2827)	(0.2393)
Disabled with some limitation	−0.7886**	−0.5159**	−0.4757**	−0.3296**	−0.3137**	−0.4210**
	(0.0814)	(0.1594)	(0.1285)	(0.0755)	(0.1283)	(0.1106)
Disabled with no limitation	−0.2066**	−0.3464**	−0.3397**	−0.0175	−0.1811**	−0.2732**
	(0.1042)	(0.2161)	(0.1380)	(0.0928)	(0.1497)	(0.1326)
Lagged disability						
Disabled with severe limitation	−1.0555**	−0.2534	−0.0765	−0.6203**	−0.1470	0.0102
	(0.1275)	(0.2593)	(0.2465)	(0.1626)	(0.2863)	(0.2643)
Disabled with some limitation	−0.5802**	0.0259	0.1796	−0.2742**	−0.0056	0.0514
	(0.0783)	(0.1592)	(0.1302)	(0.0714)	(0.1303)	(0.1177)
Disabled with no limitation	−0.0925	0.0887	0.1298	−0.0290	−0.0495	−0.0464
	(0.1175)	(0.2254)	(0.1461)	(0.0962)	(0.1566)	(0.1363)
Initial condition						
LFP in 1995		1.2059**	0.6399**		0.8984**	0.6315**
		(0.2096)	(0.0944)		(0.1353)	(0.0626)
Random effect (time averages)						
Disabled with severe limitation		−0.8815**	−0.9013**		−0.3077	−0.2653
		(0.5948)	(0.4588)		(0.7211)	(0.5607)
Disabled with some limitation		−0.7265**	−0.7146**		−0.1387	−0.1209
		(0.3237)	(0.2371)		(0.2744)	(0.2041)
Disabled with no limitation		0.3616	0.2146		0.4464*	0.5171*
		(0.5068)	(0.3297)		(0.3844)	(0.3087)
Constant	0.4642**	−0.8210**	−1.0449**	−0.5446**	−0.1118**	−1.5214**
	(0.1332)	(0.2167)	(0.1332)	(0.1074)	(0.1595)	(0.0945)
N	5930	5930	5930	6330	6330	6330
Pseudo R^2	0.2772		0.5371	0.1700		0.5303
Rho		0.4684**			0.3984**	

$^{**}p \leq 0.05$, $^{*}p \leq 0.10$.
(Significance in random effects models are based on *t*-stats on base coefficients, not on the rescaled coefficients reported in this table). Estimation was carried out using the xtprobit command in Stata Version 7.0.

In terms of the other explanatory variables (see Table A1), we see that labour force participation increases with age up to 34 (compared to those aged 55–64), but the effect falls slightly after the age of 44. Those with secondary or third level education have a greater probability of participating in the labour market. As expected, we see that women with children are less likely to participate, and this effect gets smaller as the youngest child is older. The opposite effect is found for men, where children increase the probability of participation, in particular when the youngest child is either aged less than 4, or in the older age group of 12–18.

Dynamic model

The results from the dynamic random effects probit model with correlated heterogeneity are presented in Table 6, columns 2 and 5 for men and women, respectively. We discuss these results in three steps, (1) state dependence, (2) the effect of current and lagged disability and (3) unobserved heterogeneity.

The coefficient on lagged participation is viewed as an indicator of state dependence, and suggests that previous participation has a significant positive effect on current participation, for both men and women. This suggests, that even after controlling for observed and unobserved differences among individuals, participation in the previous year is associated with a higher probability of participation in the current year. This effect is similar for men and women.

Current disability with severe and some limitations now has a lower effect on current participation, and this difference is more pronounced for men. Previous disability is now insignificant for men and women. By including past participation into the model, the effect of previous disability appears to have no effect on current participation. This suggests that previous disability may have influenced previous participation, and now influences current participation via the channel of past participation. This does not imply, that past disability has no effect on current participation – it simply suggests that its effect is now operating through the channel of past participation. In this respect our findings are similar to those reported in [4]. In [4] lagged health was not an important determinant of labour force exits. However, it is important to realise that the sample used in [4] was restricted to those individuals who were working at time $t-1$, which is equivalent to conditioning on past participation. Our results make this conditioning explicit.

Disney *et al.* [11] tested whether participation responses to health shocks were symmetric and found that health improvements have a weaker effect on transitions from inactivity than the reverse. One could test for these asymmetries in our model by including an interaction term between current and lagged disability. However, given the three-fold specification of disability adopted in this paper, a full set of interactions would be intractable. This is an issue that we hope to return to in future work.

The results from this dynamic model, suggest that unobserved characteristics may have been part of the effect of current disability in the pooled model for men. Indeed, if we look at the correlated part of the random effect (time averages), this would suggest that having severe or some limitations is associated with unobserved characteristics that reduce the probability of participation for men, i.e. part of the original current disability effect is due to unobserved characteristics. For women, the disability results of the random effects model are generally the same as in the static pooled model. The extent of unobserved effects is higher in the model for men, with 47% of the total variance due to unobserved heterogeneity. The corresponding result for women is 40%.

Two different patterns emerge for men and women when we use the pooled estimator of the dynamic model. The results of the dynamic pooled probit model are presented in columns 3 and 6 of Table 4. Firstly, for men the effects of all variables are generally the same, compared to the random effects model, with the exception of lagged and initial participation. Previous participation has a higher effect, and initial participation has a lower effect. This could indicate that the random effects estimate of state dependence, may be biased due to a violation of the no-feedback assumption. For women, the effects of current disability are now higher compared to those in the random effects model. The effect of young children has increased slightly. The estimate on lagged disability has increased, and the effect of initial participation is now lower.

We note that although the random effects model for women may be preferable, we would still expect reasonably similar results from the pooled dynamic model. This is not the case, as the pooled model provides more negative estimates of disability. To explore this further, we again followed [5] and tested for the exogeneity of two variables – age of youngest child and education. Third level education failed the strict exogeneity test, and it is possible that there is some interaction between disability and education for women. Since the disabled may have a lower incentive to invest in education, the effects of disability on participation are small once the endogeneity of education is taken into account, [12]. This will be explored in future work.

Average partial effects

So far, we have presented the results as parameter estimates, but it is also interesting to present some of the results as percentage effects. So we now estimate some average partial effects, using the population-averaged parameters $\beta_a = \hat{\beta}/\sqrt{(1+\sigma_a^2)}$. This allows us to get partial effects, that are averaged over the population distribution of the unobserved effect and we can then compare these to the partial effects of the pooled model. The probability of participation is $N^{-1}\sum_{i=1}^{N}\Phi(\hat{\psi}_a + x_{it}\hat{\beta}_a + \bar{x}_i\hat{\xi}_a) = N^{-1}\sum_{i=1}^{N}\Phi[(\psi + x_{it}\beta + \bar{x}_i\xi)(1+\sigma_a^2)^{-1/2}]$ and for a discrete variable we evaluate this expression at different values for x_{it}, i.e. 0 and 1, and form the difference to obtain the average partial effect. The average partial effect for a continuous variable x_j is obtained by using the average across i of $\hat{\beta}_{aj}\phi(\hat{\psi}_a + x^0\hat{\beta}_a + \bar{x}_i\hat{\xi}_a)$.

Our main variables of interest are current and lagged disability, but the parameter estimates for lagged disability in the dynamic models are insignificant. For this reason, we only discuss the average partial effects calculated for current disability and lagged participation. In Table 7, columns 1 and 4, we see that the average partial effect of current disability is similar for men and women in the pooled static model. Once we introduce unobserved heterogeneity and state dependence into the model, this effect is much lower for men. In the pooled dynamic model, disabled men who are severely limited in daily activities are approximately 8 percentage points less likely to participate compared to those with no disability. Although this effect is quite small, we also see that men who did not participate in the previous year have a lower probability of current participation by 40 percentage points. The parameter estimates of lagged disability were insignificant in this model, suggesting that part of the non-participation in the previous period is due to the effect of previous disability.[f]

The results for women are quite different, in that when we control for unobserved heterogeneity and state dependence, the effect of current disability is now slightly higher in the pooled dynamic model, compared to the pooled static model. However, the preferred dynamic model for women may be the random effects model, given that we did not reject strict exogeneity of the disability variables. Therefore, the results suggest that women who are currently severely limited have a lower probability of current participation by 25 percentage points. The effects of some and no limitations are much lower. Similar to the case of men, when we compared the static and dynamic models, we saw earlier that the effect of lagged disability is no longer significant. In Table 7, we show that the average partial effect of lagged participation is 13 percentage points – this is the magnitude of state dependence.

Within the context of similar research using data from other countries, the contribution of unobserved effects to the base disability effect is quite similar in this paper. Using data for the UK, [13] show that 50% of the difference in participation rates between disabled and non-disabled men is due to unexplained effects. Likewise, Kreider [3] uses US data and finds that the estimate of disability for men is overestimated by 17.2%. Lindeboom and Kerkhofs [2] use data from the

Table 7. Average partial effects

	Men			Women		
	[1] Pooled static	[2] Random effects dynamic (rescaled)	[3] Pooled dynamic	[4] Pooled static	[5] Random effects dynamic (rescaled)	[6] Pooled dynamic
Disabled with severe limitation	−0.3346** (0.0504)	−0.1111**	−0.0865** (0.0471)	−0.3377** (0.0502)	−0.2557**	−0.3979** (0.0598)
Disabled with some limitation	−0.1680** (0.0238)	−0.0746**	−0.0654** (0.0230)	−0.1308** (0.0295)	−0.0787**	−0.1666** (0.0428)
Disabled with no limitation	−0.0330** (0.0187)	−0.0461**	−0.0438** (0.0221)	−0.0069 (0.0369)	−0.0435**	−0.1086** (0.0524)
Lag LFP*		0.1292**	0.3927**		0.1296**	0.6286**

$^{**}p \leqslant 0.05$, $^{*}p \leqslant 0.10$.
(Significance in random effects models are based on t-stats on base coefficients, not on the rescaled coefficients reported in this table). Estimation was carried out using the xtprobit command in Stata Version 7.0.

Netherlands and show that the effect of bad health on the probability of receiving disability benefit is overestimated, but the effect on the probability of receiving unemployment benefit is underestimated. The coefficients for the base models are -4.179 and -0.826, and for the corrected models are -2.261 and -2.131, respectively. Compared to all these findings, our parameter estimates for currently disabled men with severe or some limitations, suggest that approximately 40–50% of the base effect is due to unobserved individual effects/ state dependence. For women, we find that the original estimates of severe and some limitations are overestimated by about 5–10%.

In terms of policy, the results from this paper show that unobserved effects are an important factor in the participation decision for disabled people. In this paper, we cannot determine the nature of these unobserved characteristics, but further knowledge on these effects are necessary for integration of disabled people into the labour force. We find that past participation is also an important factor in the participation decision for disabled people, and the effect of past disability on past participation is relevant in this context.[g] The results highlight the difference in effects between longer term and short-term disability. The effect of past disability may have a continued effect through state dependence in labour force participation, even after recovery from the disability. Therefore, the focus of disability policy should be on early targeting of disabled individuals into employment. Additional information on how participation affects future disability will also prove useful, in that we may be able to establish how past occupational injuries from past participation affect current disability and participation, and people with these disabilities may re-join the labour force. The incentive effects of disability benefits may also play a role here and these factors will be investigated in future research.

Conclusions

People with disabilities face many barriers to full participation in the labour market, with serious implications for living standards and quality of life. This paper has analysed the factors associated with participation or non-participation in the labour market, using data on people reporting chronic illness or disability in a large-scale Irish representative survey. The results of the panel analysis presented in this paper, bring out the scale of the impact on labour force participation, of having an illness or disability that limits the individual severely in their daily life.

We controlled for state dependence and unobserved heterogeneity by estimating a dynamic model with correlated random effects. The results show that unobserved heterogeneity contributes substantially to the base effect of disability for men, and to some extent for women. In our preferred model, (pooled dynamic) disabled men with a current severe limitation are now only 9 percentage points less likely to participate compared to non-disabled men. However, the effect of past participation is quite high, at 40 percentage points. For women, our preferred model is the dynamic model with correlated random effects. Those with a severely limiting disability have a lower probability of participation by 26 percentage points, compared to women with no disability. The effects of some and no limitations are less substantial. The effect of past participation is lower in the model for women, reducing current participation by 13 percentage points. The interaction of disability, education and participation of women, should be explored further.

In this paper, we aimed to provide more accurate estimates of the effect of disability on participation. However, we acknowledge some limitations. In particular, if the reporting of disability in the survey is prone to measurement error, we cannot estimate the true effect of disability on participation. This may help to explain the substantial contribution of unobserved individual effects, but without extending the model to allow for measurement error in reporting behaviour, our results on the effect of disability on participation are not conclusive. Again, this will form part of future research where we will model labour force participation and disability, while controlling for reporting behaviour.

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Appendix A

Other explanatory variables of panel model results are given in Table A1.

Table A1. Panel model results – other explanatory variables

	Men (coefficients)			Women (coefficients)		
	[1] Pooled static	[2] Random effects dynamic (re-scaled)	[3] Pooled dynamic	[4] Pooled	[5] Random effects dynamic (re-scaled)	[6] Pooled dynamic
Age 15–24	0.0881**	−0.8044*	−0.5994	0.9325**	−0.1242	0.0592
	(0.1631)	(0.6526)	(0.4252)	(0.1408)	(0.3934)	(0.3009)
25–34	0.9489**	−0.2594	−0.2330	1.2672**	−0.0685	−0.0317
	(0.1594)	(0.5269)	(0.3671)	(0.1118)	(0.3048)	(0.2232)
35–44	0.9263**	−0.2174	−0.2452	1.2020**	−0.0020	0.0226
	(0.1431)	(0.3834)	(0.2523)	(0.1078)	(0.2496)	(0.1789)
45–54	0.5843**	0.0922	0.0223	0.7312**	0.0905	0.0609
	(0.1066)	(0.2447)	(0.1685)	(0.0935)	(0.1784)	(0.1269)
Secondary education	0.3396**	−0.0350	−0.0513	0.4454**	−0.0354	−0.0590
	(0.0941)	(0.1923)	(0.1365)	(0.0687)	(0.1422)	(0.0902)
Third level education	0.4645**	0.6479**	0.5838**	1.2310**	0.2164*	0.2114
	(0.1275)	(0.2693)	(0.2174)	(0.1041)	(0.2059)	(0.1574)
Married	0.2918**	0.6706	0.5780	−0.3147**	−0.3427**	−0.3765**
	(0.1309)	(0.6458)	(0.4449)	(0.0894)	(0.2915)	(0.1842)
Age youngest child <4	0.3949**	0.2806	0.2240	−0.6454**	−0.6096**	−0.7032**
	(0.1913)	(0.4664)	(0.2715)	(0.1051)	(0.2177)	(0.1754)
>= 4 and <12	0.1202	0.2101	0.0871	−0.3852**	−0.3356**	−0.3934**
	(0.1435)	(0.3776)	(0.2241)	(0.0917)	(0.1987)	(0.1563)
>= 12 and <18	0.3626**	0.2887	0.1881	−0.1006	−0.2261**	−0.2767**
	(0.1177)	(0.2512)	(0.1491)	(0.0885)	(0.1566)	(0.1227)
Unearned income/100	−0.0021	0.0077	−0.0043	−0.0228**	0.0026	−0.0031
	(0.0142)	(0.0274)	(0.0244)	(0.0092)	(0.0145)	(0.0106)
BMW	0.1935**	0.1534	0.1836*	−0.0942	−0.0253	−0.0200
	(0.0846)	(0.1787)	(0.1026)	(0.0664)	(0.1222)	(0.1067)
Random effect (*time averages*)						
Age 15–24		1.1475**	0.8998*		0.9388**	0.8116**
		(0.7107)	(0.4639)		(0.4491)	(0.3238)
25–34		0.8831**	0.8192**		0.7351**	0.7433**
		(0.5869)	(0.4005)		(0.3594)	(0.2513)
35–44		0.9544**	0.9506**		0.8458**	0.8774**
		(0.4605)	(0.2951)		(0.3078)	(0.2084)
45–54		0.3444	0.3871*		0.4373**	0.5064**
		(0.3034)	(0.2013)		(0.2386)	(0.1579)
Married		−0.6698	−0.5980		0.1869	0.1999
		(0.6708)	(0.4587)		(0.3168)	(0.2058)
Secondary education		0.4802**	0.4405**		0.2498**	0.2794**
		(0.2467)	(0.1637)		(0.1731)	(0.1113)
Third level education		−0.3652	−0.3497		0.3795**	0.4228**
		(0.3198)	(0.2347)		(0.2567)	(0.1877)
Age youngest child <4		0.2600	0.2245		0.1913	0.2489
		(0.5784)	(0.3555)		(0.2803)	(0.2116)
>= 4 and <12		−0.1027	−0.0108		0.2234	0.2855
		(0.4472)	(0.2590)		(0.2405)	(0.1802)
>= 12 and <18		0.1202	0.1151		0.2012	0.2574*
		(0.3339)	(0.2052)		(0.2158)	(0.1555)
Unearned income/100		−0.0137	−0.0018		−0.0310**	−0.0248*
		(0.0393)	(0.0311)		(0.0225)	(0.0146)
BMW		0.1183	0.0743		−0.0233	−0.0291
		(0.2250)	(0.1343)		(0.1552)	(0.1166)

$^{**}p \leqslant 0.05$, $^{*}p \leqslant 0.10$.
(Significance in random effects models are based on *t*-stats on base coefficients, not on the rescaled coefficients reported in this table). Estimation was carried out using the xtprobit command in Stata Version 7.0.

Notes

a. The employment policies (Employment Equality Act 1998 and Equality Act 2004) in Ireland define disability as 'including total or partial absence of bodily or mental facilities, chronic disease, whether manifest or not, learning and personality disorders'. These policies are directed at all individuals with a disability, even if not registered as disabled.
b. The issue of reporting behaviour has been dealt with in the context of the retirement literature. Bound [1] gives an excellent exposition of the issues involved with reporting behaviour and the resulting bias, and concludes that without external information about reporting behaviour it is not possible to identify the extent of the reporting bias. Later studies follow this approach, (e.g. [2,3]) and make assumptions regarding systematic reporting behaviour to identify the extent of reporting bias. This paper does not focus on reporting issues.
c. Another data source is a special module on disability included with the Quarterly National Household Survey (QNHS) in the second quarter of 2002, which focused on the extent and nature of restriction of activities for people with disabilities and their labour force status. Similar analyses of disability labour force participation in a cross-sectional context, were carried out using QNHS data and we arrive at similar conclusions obtained from the Living in Ireland 2000 data.
d. Our specification includes a measure of unearned income but does not include a control for wages. Correctly accounting for the relationship between disability and wages is a topic for future research.
e. We tested for non-random attrition using the procedure suggested by Wooldridge [5]. The results find no evidence to suggest that our reported results are affected by non-random attrition.
f. It is important to realise that the insignificance of the lagged disability effect arises from the modelling of participation dynamics and not unobserved heterogeneity. Lagged disability remains significant in random effects models that do not model labour force dynamics.
g. The overall result in this paper is the same as in [4] – people with lagged disabilities have the same participation rate as those without previous disabilities. In the retirement literature, Au *et al*. [10] find that past disability decreases the probability of participation, but they do not control for state dependence. An additional finding in our paper is that the effect on participation is influenced by previous disability via previous participation.

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