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Disability and Labour Force Participation in Greece: A Microeconometric Analysis

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Αναπηρία και συμμετοχή στην αγορά εργασίας στην Ελλάδα: Μία μικροοικονομετρική ανάλυση

Νικόλαος Κ. Κανελλόπουλος

ΠΕΡΙΛΗΨΗ

Σύμφωνα με στοιχεία της ΕΛΣΤΑΤ το 2002 το 18,2% του πληθυσμού της Ελλάδας αντιμετωπίζει κάποιο χρόνιο πρόβλημα υγείας ή κάποια μορφή αναπηρίας. Αν και το ποσοστό ανεργίας για τα άτομα αυτά εμφανίζεται μικρότερο σε σχέση με το αντίστοιχο του γενικού πληθυσμού, το ποσοστό των μη οικονομικά ενεργών είναι εντυπωσιακά υψηλότερο (84%) σε σχέση με αυτό του γενικού πληθυσμού (58%). Αυτοί οι αριθμοί υπονοούν σοβαρές επιπτώσεις στην ποιότητα ζωής των ατόμων με χρόνια προβλήματα υγείας ή κάποια μορφή αναπηρίας και συγχρόνως επιβαρύνσεις στον Κρατικό Προϋπολογισμό με κονδύλια για συντάξεις ή επιδόματα αναπηρίας. Ενίσχυση της ενεργοποίησης αυτών των ατόμων και αξιοποίηση των παραγωγικών τους δυνατοτήτων στην αγορά εργασίας είναι προφανές ότι θα έχει θετική επίδραση τόσο στην οικονομική και κοινωνική τους ζωή, καθώς και στα οικονομικά του δημοσίου. Προς τούτο είναι χρήσιμο για τους σχεδιαστές της σχετικής πολιτικής, καθώς και για κάθε ενδιαφερόμενο, μεταξύ άλλων, να γνωρίζουν επακριβώς ποια και πόση είναι η επίδραση της αναπηρίας στην απόφαση για συμμετοχή στην αγορά εργασίας. Διαφορετικό μίγμα πολιτικής χρειάζεται να ακολουθηθεί αν η αναπηρία έχει έντονα αρνητικό αντίκτυπο στη συμμετοχή και διαφορετικό μίγμα αν έχει μικρή ή αμελητέα επίδραση.

Η ακριβής μέτρηση της επίδρασης της αναπηρίας στη συμμετοχή στην αγορά εργασίας δεν είναι μια απλή διαδικασία. Αρχικά, η συμμετοχή στην αγορά εργασίας είναι μια δυναμική διαδικασία και έτσι άτομα που ήδη συμμετέχουν έχουν μεγαλύτερη πιθανότητα να συμμετέχουν και στο μέλλον. Επιπλέον, πέρα από τα παρατηρούμενα και καταγραμμένα χαρακτηριστικά των ατόμων (φύλο, ηλικία, εκπαίδευση, οικογενειακή κατάσταση) υπάρχουν και ορισμένα απαρατήρητα χαρακτηριστικά, τα οποία μπορεί να επηρεάζουν θετικά ή αρνητικά την απόφασή για συμμετοχή στην αγορά εργασίας. Εάν αυτά δε ληφθούν υπόψη και δεν χρησιμοποιηθούν κατάλληλες οικονομετρικές τεχνικές , τότε θα υπάρχει σφάλμα στη μέτρηση της επίδρασης των προβλημάτων υγείας στη συμμετοχή στην αγορά εργασίας και πιθανόν λάθος προτάσεις πολιτικής.

Είναι λογικό ότι διαστρωματικά στοιχεία, όπως τα παραπάνω της ΕΛΣΤΑΤ, δεν προσεγγίζουν ούτε το δυναμικό χαρακτήρα της συμμετοχής στην αγορά εργασίας ούτε την απαρατήρητη ετερογένεια των ατόμων, μιας και αποτυπώνουν ένα στιγμιότυπο από τη ζωή τους. Σε αυτή την περίπτωση χρονικά επαναλαμβανόμενα διαστρωματικά στοιχεία (*longitudinal data*) προσφέρουν περισσότερα πλεονεκτήματα, αφού επιτρέπουν τη χρήση πιο πολύπλοκων υποδειγμάτων, τα οποία μπορούν να εξετάσουν διαδικασίες οι οποίες είναι εκ φύσεως δυναμικές αλλά και να ελέγξουν για κάθε μορφή ετερογένειας, παρατηρούμενης ή μη.

Στην παρούσα εργασία χρησιμοποιούνται τα ελληνικά στοιχεία από το ευρωπαϊκό πάνελ (*European Community Household Panel - ECHP*) για τα έτη 1994-2001.

Για την ανάλυση της πιθανότητας συμμετοχής στην αγορά εργασίας χρησιμοποιούνται διάφορα υποδείγματα *probit*. Αρχικά χρησιμοποιείται ένα διαστρωματικό στατικό *probit*. Στη συνέχεια το υπόδειγμα εμπλουτίζεται, περιλαμβάνοντας μια δυναμική σχέση μεταξύ τωρινής και προηγούμενης συμμετοχής. Ακολούθως το δυναμικό υπόδειγμα εκτιμάται λαμβάνοντας υπόψη την απαρατήρητη ετερογένεια (*unobserved heterogeneity*) με τις κατάλληλες τεχνικές τυχαίων επιδράσεων (*random effects*). Τέτοια δυναμικά υποδείγματα τυχαίων επιδράσεων, παρόλα τα πλεονεκτήματά τους στην πληρότητα μέτρησης, παρουσιάζουν το πρόβλημα των αρχικών συνθηκών (*initial conditions*). Προκειμένου να αντιμετωπιστεί αυτό και τα αποτελέσματα να μην είναι μεροληπτικά, εκτιμώνται υποδείγματα που διορθώνουν το πρόβλημα των αρχικών συνθηκών όπως αυτά προτάθηκαν από τους Heckman, Orme και Wooldridge.

Τα προβλήματα αναπηρίας ομαδοποιούνται σε τρεις κατηγορίες ανάλογα με το πόσο σοβαρή επίδραση ασκούν στις καθημερινές δραστηριότητες των ατόμων. Από τις εκτιμήσεις προκύπτει ότι η επίδρασή τους στην πιθανότητα συμμετοχής στην αγορά εργασίας είναι αρνητική και αυξάνεται (σε απόλυτα μεγέθη) όσο εντείνεται το μέγεθος των προβλημάτων. Ειδικότερα, η επίδραση των προβλημάτων αναπηρίας κυμαίνεται από -9,4% έως -47,4% στο στατικό υπόδειγμα. Όταν ωστόσο συνυπολογίζεται η δυναμική φύση της συμμετοχής η επίδραση μειώνεται σε -9,5% έως -34%. Όταν επιπλέον η απαρατήρητη ετερογένεια λαμβάνεται υπόψη η αρνητική επίπτωση μειώνεται μεταξύ -6% και -20%. Οι δε αρχικές συνθήκες είναι πάντοτε ενδογενείς, δεν επηρεάζουν ωστόσο το μέγεθος της επίδρασης των προβλημάτων υγείας στη συμμετοχή, όπως κάνουν στις υπόλοιπες ανεξάρτητες μεταβλητές. Συνοπτικά προκύπτει ότι η μη σωστή μέτρηση, με τη χρήση κατάλληλων οικονομετρικών υποδειγμάτων, φαίνεται ότι υπερεκτιμά την επίδραση των προβλημάτων αναπηρίας στην απόφαση για συμμετοχή στην αγορά εργασίας.

Σε όλα τα εκτιμημένα μοντέλα, συμμετοχή στην αγορά εργασίας κατά την προηγούμενη περίοδο αυξάνει τη πιθανότητα συμμετοχής κατά την τρέχουσα περίοδο. Η επίδραση αυτή κυμαίνεται από 36,5% έως 41,7% καταδεικνύοντας ότι υπάρχει *genuine state dependence*. Όσον αφορά τα υπόλοιπα χαρακτηριστικά των ατόμων φαίνεται ότι συστηματικά η ηλικία και η εκπαίδευση αυξάνουν την πιθανότητα συμμετοχής, ενώ το να ζει κάποιος στην Αττική και να είναι παντρεμένος τη μειώνουν. Είναι ενδιαφέρον ότι η ύπαρξη μικρών παιδιών στο νοικοκυριό είναι στατιστικά ασήμαντη, ενώ το ύψος του εισοδήματος που δεν προέρχεται από εργασία έχει αρνητική επίδραση, η οποία ωστόσο εκδηλώνεται μόνο μέσω της απαρατήρητης ετερογένειας.

Disability and Labour Force Participation in Greece: A Microeconometric Analysis†

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Abstract

This paper examines the effect of disability upon the decision to participate in the Greek labour market, using data from the European Community Household Panel 1994-2001. A range of dynamic random effects probit models accounting for unobserved heterogeneity, initial conditions and state dependence have been estimated. Results indicate that not accounting for such potential estimation issues tends to overestimate the effect of disability and previous participation upon current participation. In models ignoring dynamics, unobserved heterogeneity and initial conditions the effect of disability varies between 9.4 and 47.4 per cent depending on the level of disability, while when these issues are accounted for the effect of disability is reduced to 5.8 to 20.8 per cent.

Keywords: Disability, labour force participation, initial conditions, dynamic random effects probit models, Greece

JEL Classification: J22, C33, C35

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1. INTRODUCTION

Individuals with disabilities face numerous obstacles to enter and remain in employment. This has severe implications both on their standard of living and in their social life. On the contrary their labour market integration increases their earnings, promotes social cohesion and contributes to the improvement of the aggregate labour market indicators as well as reduces the expenditure of national budget. In this respect many countries have revised their policy priorities towards disabled and instead of providing services and benefits to disables; the encouragement of equality and full labour market participation is promoted. The success and evaluation of such policies however depends, among other things, upon the exact influence of disability on labour force participation and it would be helpful to both government officials and employers to know the magnitude of the impact of disability upon labour market participation. Other policy implications emerge when disability exerts a strong negative influence upon disables activation and other when its effect is slight or insignificant.

The exact specification and measurement of the impact of disability upon labour market participation involves certain theoretical and methodological problems. The effect of unobserved characteristics of disabled, as well as their previous employment status should be taken into consideration when such an analysis is undertaken. This paper, using panel data, attempts to identify the effect of disability upon labour market participation taking into account the effect of unobserved heterogeneity and previous participation for Greece.

These issues are rather relevant for Greece, a country with relatively low participation ratio of the productive age population, demographic ageing, perennial public budget deficits and currently unsustainable public debt. Even though the employment ratio in Greece is on the rise and from 55% in 1990 it steadily increased to 62% in 2009, it remains lower than in many other European Union countries (Kanellopoulos (2011), European Commission (2010), OECD (2009)). This upward trend reflects the increasing activation of women, especially the young, while that of men remains rather stable. These figures are far behind the targets set by Lisbon Agenda and presumably have been feeding the consistent social security deficits and hampered the economic performance of the country. The upward trend of young female participation in the recent years indicates that it would continue to increase in the coming years and would reach that of EU average. In this context, apart from the information provided from the 2002 ad hoc LFS module that 84 per cent of individuals with health problems are out of the labour market (Hellenic Statistical Authority (2003)), it is not clear what is the actual effect of health status upon labour market activation.

In this paper we use longitudinal data to measure the impact of disability on labour market participation controlling for state dependence, unobserved heterogeneity and initial conditions. These longitudinal data are unique in the sense that they are the only available source of information allowing the proper examination of the effects of health status upon participation in Greece. Their panel nature provides information regarding the same individual for a longer time allowing us to apply techniques which take into consideration unobserved heterogeneity as well as previous labour market status. In this context this paper is the first systematic attempt to measure the effect of disability upon labour market participation in Greece. It also contributes to a recently growing literature using econometric techniques which eliminate the initial conditions problem in dynamic non linear panel models. To achieve that we estimate a range of dynamic random effects probit models that allow the initial conditions to be endogenous and incorporate unobserved heterogeneity. These models also measure accurately the effect of previous participation i.e. state dependence and thus result in a more precise estimation of the true effect of disability upon labour market participation.

The structure of the paper is as follows. Section 2 describes the data used in the analysis as well as the relevant variables. The econometric techniques applied are presented in section 3, while the results are discussed in section 4. Section 5 concludes.

2. DATA

The data used in this paper are from the Greek side of the European Community Household Panel (ECHP) dataset, which is a harmonized cross-national longitudinal survey focusing on household income and living conditions for many European Union countries. ECHP was designed by Eurostat, it ran from 1994 to 2001 and the Greek data were collected by National Statistical Service of Greece following a centrally designed questionnaire [1]. We have restricted our analysis to working age individuals. The original sample contains 15,374 individuals resulting to 85,748 person-year observations. Since in the analysis the lagged dependent variable is used as a regressor, the sample was further restricted to include only consecutive information. When the sample criteria were applied and observations with missing information were excluded, the final sample consists of 44,755 person-year observations for 8,959 adults. It is worth noting that out of these individuals only 4,445 are observed in all 8 waves, thus we use an unbalanced panel dataset.

The unit of analysis is the individual and the definition used to construct the labour market participation is that of International Labour Office (ILO, 2004). Thus, an individual is in the labour market when (s)he is working or (s)he is unemployed[2]. Table 1 provides information regarding the distribution of labour force participation across waves. The average rate of participation over the eight examined years is around 65 per cent. This ratio does not change much over the years. It is informative to examine movements into and out of the labour market since the same individuals might not continuously participate in successive years. Thus, rows [2] to [4] of Table 1 report the conditional probabilities by participation status at *t-1*. Row [2] shows that the probability of participating at *t* is much higher for those who

[1] For more details on ECHP see EPUNet (2004), Eurostat (2003a), Eurostat (2003b), Eurostat (2004) and Peracchi (2002).

[2] According to the ILO/OECD unemployment definition an individual is unemployed if (s)he does not have a job, (s)he had looked for a job in the past four weeks and is available for work.

participated at *t-1*. In particular, during the examined period the raw conditional probability of remaining in the labour market is on average 92.2 per cent, i.e. raw data indicate that there is considerable state dependence in labour force participation. The average raw entry probability is 12.2 per cent suggesting that someone participating at *t-1* is 7.6 times more likely to remain in the labour market at *t* compared to someone out of the labour market at *t-1*. Two measures of state dependence, the difference and the ratio between remaining and entering the labour market are presented in rows [5] and [6] respectively. One can immediately see that the state dependence for the case of participants is 80 percentage points, while the probability of participating conditional on previous year participation is 7.6 times higher compared to someone without previously participating.

Table 1: Labour force participation by wave

Year		1995	1996	1997	1998	1999	2000	2001	All
Pr(y_t=1)	[1]	64.66	63.95	64.44	65.62	64.83	64.90	66.65	64.98
Pr(y_t=1|y_{t-1}=1)	[2]	88.91	91.14	93.16	93.16	91.27	95.66	95.28	92.22
Pr(y_t=1|y_{t-1}=0)	[3]	15.99	12.03	14.13	11.87	6.58	12.33	8.27	12.18
Pr(y_t=0|y_{t-1}=1)	[4]	11.09	8.86	6.84	6.85	8.73	4.34	4.73	7.78
State Dependence									
Difference	[2] - [3]	72.92	79.11	79.02	81.29	84.69	83.33	87.01	80.04
Ratio	[2] / [3]	5.56	7.57	6.59	7.85	13.86	7.76	11.52	7.57

Note: y_t=1 if the individual is participating in time *t* and zero otherwise.
Source: ECHP, waves 1-8.

It emerges from Table 1 that although the probability of participation remains rather the same over the examined period, the difference in the two conditional probabilities increased substantially. This stems from the fact that the participation incidence among those previously participating consistently increased, while that for those previously out of the labour force fell almost by 50 per cent. This might be interpreted as an insiders-outsiders phenomenon in the Greek labour market.

These raw probabilities ignore any observed characteristics of examined individuals. Table 2 presents labour force participation probabilities (unconditional and conditional) for various subgroups of the sample. Column [1] displays the probability of participating while columns [2] – [4] the probability of participating conditional on participation status a year ago. The difference between the unconditional and conditional probabilities is substantial within all examined subgroups, suggesting that there is considerable persistence in labour market participation.

Men have a high participation probability 82.5 per cent which is almost double to that of women (48.8 per cent). The conditional probability of remaining in the labour market is high for both men (95.7 per cent) and women (86.8 per cent). The entry probability is slightly higher for men than women, while the exit probability for women is three times higher than that of men.

Table 2 shows that there is considerable age effect on the likelihood of participating in the labour force. The probability of labour force participation increases with age until the age of 45 and after that it falls. The same age pattern holds for remaining in the labour force, while the entry probability declines as age advances. It is worth noting that exiting from the labour force is higher for younger and older individuals.

Table 2: Unconditional and conditional labour force participation probabilities

	$Pr(y_t=1)$ [1]	**$Pr(y_t=1 \mid y_{t-1}=1)$ [2]**	**$Pr(y_t=1 \mid y_{t-1}=0)$ [3]**	**$Pr(y_t=0 \mid y_{t-1}=1)$ [4]**
Gender				
Male	82.55	95.69	15.12	4.31
Female	48.85	86.81	11.33	13.19
Age				
16-25	59.06	86.09	26.45	13.91
26-35	78.51	94.29	20.04	5.71
36-45	78.28	95.41	14.95	4.59
46-55	68.65	93.60	8.91	6.40
56-65	37.56	82.57	4.47	17.43
Highest level education completed				
University	85.31	95.99	19.05	4.01
High school	68.10	92.84	15.64	7.17
Primary school	57.38	90.11	9.77	9.89
Marital status				
Married	63.60	92.22	9.49	7.79
Not married	68.77	92.42	20.41	7.58
Housing				
Owner	63.71	92.09	11.42	7.91
Renting	70.44	93.24	15.23	6.76
Region of residence				
Attica	63.50	92.79	11.39	7.21
Northern Greece	47.04	85.09	11.17	14.91
Central Greece	51.42	86.63	12.42	13.37
Islands & Crete	66.18	92.21	13.79	7.79
ALL	**64.98**	**92.22**	**12.18**	**7.78**

Note: $y_t=1$ if the individual is participating in time *t* and zero otherwise.
Source: ECHP, waves 1-8.

More educated individuals are more likely to participate. In particular, the participation probability for those holding a university degree is 85.3 per cent, for high school graduates 68.1 per cent and for primary school graduates 57.4 per cent. The conditional probabilities on previous participation status are very similar around 90 per cent, but more educated individuals have higher entry and lower exit probabilities. Table 2 also shows that never married individuals are slightly more likely to participate, however controlling for previous participation status this small advantage disappears.

House ownership was used as a proxy to non wage income. Those renting a house have 7 percentage points higher probability of participating, while house owners have smaller chances of entering the labour market and marginally higher of exiting. Finally, those living in the insular Greece or in Attica have higher chances of participating compared to those residing anywhere else in Greece. Interestingly the entry and the remaining probabilities are more or less similar in the examined regions.

The overview of the dynamics of labour force participation presented in Tables 1 and 2 ignores that labour market participation can be very much influenced by health status. ECHP provides sufficient information on health status to examine this hypothesis. It asks all surveyed individual the question “Do you have any chronic physical or mental health problem, illness or disability?” To better capture the degree that this illness or disability problem may affect the individual labour market behaviour we utilize a follow up question “Are you hampered in your daily activities by this physical or mental health problem, illness or disability?” Using the relevant information we have separated the sample into four mutually exclusive groups on the basis of their health status[3].

[1] Those reporting no chronic health problems,

[2] those reporting chronic health problems but without limitations in their daily activities,

[3] those reporting chronic health problems with some limitations in their daily activities, and

[4] those reporting chronic health problems with severe limitations in their daily activities.

Table 3 presents the unconditional and conditional probability of participating by health status and gender. It is striking how much the incidence of participating declines as the health status deteriorates. On average an individual who reports no health problems has a 69.2 per cent chance of participating. This probability is reduced by 10 percentage points if (s)he reports a chronic problem without limitations and by 15 extra percentage points when (s)he reports some limitations. It is further reduced to 21.6 per cent when severe activity limitations due to chronic health problems or disabilities are reported. The difference in the rates of participation between those reporting no health problems and those reporting severe limitations due to health problems is 60.5 percentage points for men and 36.2 percentage points for women. These substantial differences clearly suggest that it is necessary to control for the different levels of disability in the following analysis.

[3] Similar approach to the measurement of disability using ECHP data for Ireland has been used by Gannon (2005).

The second column of Table 3 displays the conditional probability of remaining in the labour force by health status. Interestingly the rate of participation of those without chronic health problems and those with chronic problems without limitations is very small for both genders. Individuals reporting some limitations have a lower participation rate, on average around 85 per cent. The conditional probability for those with severe limitations is remarkably lower, around 67 per cent, for both genders. These differences suggest that the effect of state dependence is reduced substantially as health deteriorates.

It is worth mentioning that the probability of entering participation becomes smaller as the health status worsens. Those with some limitations have half the chances of participating next year given no participation the current year, compared to those without health problems. Likewise those with severe limitation have a quarter of the probability of healthy individuals to enter participation. This seems to be more intense for men. On the other hand exit from the labour market is consistently higher for individuals with poor health. In particular on average the exit probability for a healthy individual is 6.7 per cent. It slightly increases to 6.8 for those with no limitations and is further increased to 14.4 for those with some limitations. Individuals with severe limitations face a 32.9 per cent probability of exiting the labour market. The effect of health problems with severe limitations appears identical for both men and women. However, women without health problems or without severe limitations have higher likelihood of not participating the following year.

Table 3: Labour force participation by health status

	$Pr(y_t=1)$ [1]	**$Pr(y_t=1 \mid y_{t-1}=1)$** [2]	**$Pr(y_t=1 \mid y_{t-1}=0)$** [3]	**$Pr(y_t=0 \mid y_{t-1}=1)$** [4]
All	**64.98**	**92.22**	**12.18**	**7.78**
No health problems	69.17	93.27	14.10	6.73
No limitations	59.20	93.20	11.01	6.80
Some limitations	44.43	85.63	6.80	14.37
Severe limitations	21.58	67.11	3.52	32.89
Men	**82.55**	**95.69**	**15.12**	**4.31**
No health problems	88.16	96.94	20.99	3.06
No limitations	74.66	96.33	14.10	3.67
Some limitations	63.38	89.91	7.59	10.09
Severe limitations	27.66	67.61	2.98	32.39
Women	**48.85**	**86.81**	**11.33**	**13.19**
No health problems	51.84	87.61	12.57	12.39
No limitations	39.39	85.71	9.29	14.29
Some limitations	30.88	79.42	6.52	20.58
Severe limitations	15.62	66.10	3.93	33.90

Note: y_t=1 if the individual is participating in time *t* and zero otherwise.
Source: ECHP, waves 1-8.

Useful information is also the transitions between the four different health statuses and the corresponding participation rate, as provided in Table 4. It appears that less than 5 per cent of healthy individuals report health problems the following year. Out of those reporting disabilities without limitations 27 per cent declares more health problems the following year, while 55 per cent states no health problems next year. Moreover, 60 per cent of those with some limitations either continue to have the same health status or worsen, while 37.6 percent report no health problems. Finally 53.7 per cent of the individuals with severe health problems do not improve within a year. The interesting information in Table 4 is the participation probabilities given previous and current health status. It seems that when health worsens, the participation rate falls. Likewise when health improves, the participation rate increases. For instance, those who reported some limitations due to a chronic illness if their health status does not change, have on average a participation rate of 43.8 per cent. If their health worsens, the incidence of participation falls to 20 per cent, while if their health improves their participation rate increases to 49.5 or 57.8 per cent depending on the level of improvement.

Table 4: Health status transition probabilities

Year		***t***			
t-1		**[1]**	**[2]**	**[3]**	**[4]**
No health problems	**[1]**	95.12 [70.29]	0.63 [64.14]	2.95 [49.09]	1.31 [32.61]
No limitations	**[2]**	55.40 [63.56]	17.84 [65.79]	19.95 [51.76]	6.81 [41.38]
Some limitations	**[3]**	37.57 [49.51]	3.81 [57.83]	44.37 [43.79]	14.24 [20.00]
Severe limitations	**[4]**	23.48 [38.86]	1.56 [40.91]	21.22 [25.33]	53.75 [14.21]

Note: Number in brackets is the probability of participating given present and past health status.
Source: ECHP, waves 2-8.

All probabilities presented in Tables 1-3 suggest that there is strong state dependence in labour force participation, i.e. the probability of participating at *t* is higher for those participating at *t-1*. An emerging question is how much of the observed persistence in the raw data is due to observed characteristics, to unobserved characteristics (heterogeneity) or to genuine state dependence. As Heckman (1981b, 1981c) points out aggregate probabilities do not necessarily imply true state dependence. An explanation for potential spurious state dependence is that certain unobserved characteristics might raise the probability of participating, even if this is not the case in the individual probabilities. Moreover, do initial conditions have a significant effect on this persistence? The following section provides a framework of analysis of these probabilities controlling for observed and unobserved heterogeneity as well as for the initial conditions problem.

3. ECONOMETRIC MODELS

Our analysis of the effect of chronic health problems or disability on the probability of labour force participation begins with a simple static pooled probit model and then dynamics as well as unobserved heterogeneity are incorporated. Dynamic models include as independent variable the previous labour force participation status to allow for state dependence. In these models special attention should be paid to the treatment of unobserved heterogeneity, as well as the initial conditions problem, which arises when the beginning of the examined period does not coincide with the beginning of the stochastic participation process. Unobserved heterogeneity is modelled following Mundlak (1978) and Chamberlain (1984), while for the initial conditions problem solutions suggested by Heckman (1981a), Orme (1996) and Wooldridge (2005) are adopted. It is intentionally chosen to estimate a variety of alternative models and apply more than one solution to the initial conditions problem in order to examine whether the obtained results of health impact are robust. The objective is to properly measure true (structural) state dependence as it is expressed by the coefficient of the lagged dependent variable along with the effect of health status and not to model the mechanism causing this state dependence.

3.1. A Pooled Probit

The analysis begins with a basic static pooled probit, which is nothing else than a usual cross-sectional probit where observations for all individuals and from all time periods are pooled together.

$$y_{it}^{*} = x_{it}'\beta + \xi D_{it} + v_{it} \tag{1}$$

where the subscript $i=1,2,...,N$ denotes individuals that are included in the sample and the subscript $t=2,3,...,T$ represents the time periods for which the model is estimated. y_{it} is the observed indicator of participating in the labour force and takes values one if individual participates and zero otherwise, y_{it}^{*} is the underlying construct generating y_{it}, and x_{it} is a vector of strictly exogenous explanatory variables of labour force participation. D_{it} is a vector of health status indicators and v_{it} is an error term with the usual properties. This model is very simple in terms of specification but due to data limitations it is very commonly estimated. Moreover, it provides some baseline results to compare with estimates from more sophisticated models. Obviously the first improvement will be to add dynamics in (1). Thus,

$$y_{it}^{*} = \gamma y_{it-1} + x_{it}'\beta + \xi D_{it} + v_{it} \tag{2}$$

where y_{it} is the labour force participation status of individual i in the previous year $t\text{-}1$.

3.2. A Dynamic Random Effects Probit

The models so far assume that current participation is affected by previous participation, health status and other exogenous variables. However, certain unobserved characteristics may also influence the decision to participate. In order to take into consideration this unobserved heterogeneity the error term from (1) and (2) can be factorised into two components $v_{it} = \varepsilon_i + u_{it}$, so that equation (2) becomes:

$$y_{it}^{*} = \gamma y_{it-1} + x_{it}'\beta + \xi D_{it} + \varepsilon_i + u_{it} \tag{3}$$

where the variable ε_i captures all unobserved, time invariant factors that affect y_{it} and is called unobserved heterogeneity. Moreover, $\varepsilon_i \sim N\left(0, \sigma_{\varepsilon}^2\right)$ and it is independent of u_{it}, which is the usual error term and it is assumed that $u_{it} \sim iid\ N\left(0, \sigma_u^2\right)$ and serially independent. However, even if this is the case for the u_{it} the existence of the individual-specific error term (ε_i), which is time constant, makes the composite error term $v_{it}=u_{it}+\varepsilon_i$ to be serially correlated. It is customary to assume equal correlation between v_{it} in any two different time periods: $r = Corr\left(v_{it}, v_{is}\right) = \sigma_{\varepsilon}^2 \big/ \left(\sigma_{\varepsilon}^2 + \sigma_u^2\right)$ for $t,s=2,3,...,T;\ t \neq s$[4].

The standard uncorrelated random effects model also assumes that ε_i is uncorrelated with x_{it} for all i and in every t period. However, this can lead to omitted variable bias and thus it is necessary to allow for correlation between ε_i and x_{it}. Following Mundlak (1978) and Chamberlain (1984) a relationship between the unobserved heterogeneity ε_i and the time means of all time varying explanatory variables is assumed. Thus, $\varepsilon_i = \overline{x}_i'\delta + \alpha_i$ where $\alpha_i \sim iid\ N\left(0, \sigma_{\alpha}^2\right)$ and independent of x_{it} and u_{it} for all i and in all t periods. As a result, a correlated random effects probit model emerges, with extra regressors the means of all time varying variables. Substituting into (3) we get:

$$y_{it}^{*} = \gamma y_{it-1} + x_{it}'\beta + \xi D_{it} + \overline{x}_i'\delta + \alpha_i + u_{it} \tag{4}$$

A crucial assumption one needs to make in a dynamic model is whether the initial observation of the dependent variable y_{i1} and the unobserved heterogeneity α_i $_{\alpha_i}$ are correlated or not. If y_{i1} is considered as exogenous and thus uncorrelated with α_i then (4) is estimated by using the Gauss-Hermite quadrature because the likelihood can easily be decomposed into two independent factors and their joint probability for $t \geq 2$ can be maximized without referring to that of the initial period. However, this requires that the initial period is also the beginning of the stochastic process that generates labour force participation status. Nevertheless, this is not the case, as a great number of individuals in the examined sample were in the labour force well before they enter the survey and the initial conditions problem arises. In other words y_{i1} is endogenous as it is correlated with α_i , and so the obtained estimator will be inconsistent and will tend to overestimate the coefficient of the lagged dependent variable γ and underestimate the coefficients of the x-vector variables (Chay and Hyslop, 2000).

3.3. Heckman's Estimator

To deal with the initial condition problem, following Heckman (1981a), we indicate a reduced form equation for the initial observation:

[4] Because y is a binary variable it is standard to assume that $\sigma_u^2 = 1$.

$$y_{i1}^{*} = z_{i1}'\lambda + \eta_i \tag{5}$$

where y_{i1}^{*} is a binary variable taking the value of one if the individual is participating in year 1 and zero otherwise. z_{i1}' is a vector of strictly exogenous instruments, which affect y_{i1}^{*}, $\text{var}(\eta_i) = \sigma_\eta^2$ and $corr(\alpha_i, \eta_i) = \rho$. Since we do not want ρ to be zero, a linear specification is introduced, in terms of orthogonal error components:

$$\eta_i = \theta\alpha_i + u_{i1} \tag{6}$$

By construction α_i and u_i are orthogonal to one another with $\theta=\rho\sigma_\eta/\sigma_\alpha$ and $\text{var}(u_{i1}) = \sigma_\eta^2(1-\rho^2)$. Furthermore, it is assumed that the initial observation of y is not correlated with $u_{it}{}_{u_{it}}$, i.e. $E(u_{it},y_{i1})=0$ and also it is not correlated with x_{it} for all i and in all $t=2,\dots,T$.

If equation (6) is replaced into equation (5), equation (7) emerges

$$y_{i1}^{*} = z_{i1}'\lambda + \theta\alpha_i + u_{i1} \tag{7}$$

which in combination with equation (4) constitute the following full specification of Heckman's model:

$$\begin{cases} y_{i1}^{*} = z_{i1}'\lambda + \theta\alpha_i + u_{i1}, & i = 1,\dots,N \text{ and t} = 1 \\ y_{it}^{*} = \gamma y_{it-1} + x_{it}'\beta + \xi D_{it} + \overline{x}_i'\delta + \alpha_i + u_{it}, & i = 1,\dots,N \text{ and t} = 2,\dots,T \end{cases} \tag{8}$$

According to Heckman (1981a, 1981c) under the assumption that $\alpha_i \sim IN(0,\sigma_\alpha^2)$ is independent of u_{it} and that the distribution of y_{it}^{*} conditional on y_{it-1}, x_{it}, D_{it} and α_i is independent normal this model can be estimated by maximizing the following likelihood function:

$$L = \prod_{i=1}^{N} \int_{\alpha^*} \left\{ \Phi\left[\left(z_{i1}'\lambda + \theta\sigma_\alpha\alpha^* \right)(2y_{i1}-1) \right] \prod_{t=2}^{T} \Phi\left[\left(\gamma y_{it-1} + x_{it}'\beta + \xi D_{it} + \overline{x}_i'\delta + \sigma_\alpha\alpha^* \right)(2y_{it}-1) \right] \right\} dF(\alpha^*) \tag{9}$$

where F is the distribution function of $\alpha^*=\alpha/\sigma_\alpha$ and $\sigma_\alpha = \sqrt{r/(1-r)}$. With α taken to be normally distributed, the integral over α^* can be evaluated using Gaussian-Hermite quadrature (Butler and Moffit, 1982). To check the exogeneity of the initial conditions one can perform a t-test on θ.

3.4. Orme Model

Orme (1996) follows Heckman's (1979) two-step procedure for corrections for endogenous sample selection and assumes that the model is fully specified by a system of a simple probit for the initial period, like the one in equation (5), and a dynamic random effects probit model for the remaining time periods, like the one described in equation (4). In view of the fact that the correlation between the lagged dependent variable (y_{it-1}) and unobserved heterogeneity (α_i) causes the initial conditions problem Orme suggests replacing the latter with a new error term which is uncorrelated with y_{it-1}. Orme then proposes a linear specification, in terms of orthogonal error components in such a way that again $\rho \neq 0$:

$$\alpha_i = \kappa\eta_i + w_i \tag{10}$$

Substituting (10) into (4) we get:

$$y_{it}^* = \gamma y_{it-1} + x_{it}'\beta + \xi D_{it} + \bar{x}_i'\delta + \kappa\eta_i + w_i + u_{it} \tag{11}$$

Orme points that firstly in this new random effects probit, there are two individual specific random effects, η_i and w_i, secondly by construction $E(w_i|y_{i1})=0$ and thirdly if one can control for the presence of term η_i in (11) then the initial conditions problem no longer applies. Orme suggests to construct a generalised error term derived from the initial observation equation (5) , equivalent to that used in Heckman's sample selection model. Under the assumption of bivariate normality between (η_i, a_i)

$$e_i \equiv \mathrm{E}(\eta_i \mid y_{i1}) = (2y_{i1} - 1)\phi(\lambda Z_{i1}) \big/ \Phi\left[(2y_{i1} - 1)(\lambda Z_{i1})\right] \tag{12}$$

where ϕ is the Normal density function and Φ is the Normal distribution function. Because u_{it} is assumed to be orthogonal to the regressors, w_i can be treated as the common error component in a random effects probit, as long as we take care of the unobservable η_i. Taking into consideration that e_i is derived from a probit model from equation (5), it is reasonable to substitute η_i by its conditional expectation. Thus, equation (11) becomes a random-effects probit with an extra regressor e_i . A test of the null hypothesis that ρ=0, i.e. initial conditions are exogenous, can be obtained by a simple *t*-test on the coefficient of e_i. A potential problem is that the generalised error is heteroskedastic. However, Orme after performing Monte-Carlo simulations provides evidence that the estimator performed relatively well. Arulampalam and Stewart (2009) have also reached the same conclusions.

3.5. Wooldridge Model

Wooldridge (2005) proposes a parametric method of estimation, which instead of modelling the density of $f(y_1,...,y_T \mid x, a)$, models the density of $f(y_2,...,y_T \mid y_1, x, a)$. In other words Wooldridge suggests modelling the unobserved heterogeneity conditional on the value of the initial period and other exogenous variables. ε_i is expressed in terms of y_{i1} and $\bar{x}$, following Mundlak (1978) and Chamberlain (1984) and thus get:

$$\varepsilon_i = \mathrm{a}_0 + \mathrm{a}_1 y_1 + \bar{x}_i'\delta + \alpha_i \tag{13}$$

The intuition is that the correlation between unobserved heterogeneity and the initial observation is allowed in (13), which now generates a new error term uncorrelated with the initial observation. Thus, substituting (13) into (3) the model is now fully specified as:

$$y_{it}^* = \gamma y_{it-1} + x_{it}'\beta + \xi D_{it} + \mathrm{a}_1 y_1 + \bar{x}_i'\delta + \alpha_i + u_{it} \tag{14}$$

This estimator is as a random effects probit following a different approximation for the unobservables. To test whether the initial conditions are exogenous one can perform a *t*-test on a_1.

3.6. Interpretation of coefficients

Unlike linear models, the coefficients of all estimators presented in the previous sections are not equal to the change in the conditional mean of the dependent variable when regressors change by one unit. This means that it is necessary to estimate the partial effect of the independent variables the on Pr(y_{it}=1). There are several ways to estimate the marginal effects for this kind of models. Here the predicted probabilities, used to calculate the marginal effects, are estimated based on the coefficients from the estimated models and taking the variable of interest fixed at 1 and 0 while the rest of the regressors are kept in their sample mean value. For instance the relevant probabilities regarding previous participation are[5]:

$$\begin{cases} \widehat{\Pr}(y_{it}=1 \mid y_{it-1}=1, x_{it}=\overline{x}_{it}, D_{it}=\overline{D}_{it}) = \Phi\left\{\left(\gamma + \overline{x}'_{it}\hat{\beta} + \hat{\xi}\overline{D}_{it}\right)\left(1-\hat{r}\right)^{0.5}\right\} \\ \widehat{\Pr}(y_{it}=1 \mid y_{it-1}=0, x_{it}=\overline{x}_{it}, D_{it}=\overline{D}_{it}) = \Phi\left\{\left(\quad \overline{x}'_{it}\hat{\beta} + \hat{\xi}\overline{D}_{it}\right)\left(1-\hat{r}\right)^{0.5}\right\} \end{cases} \tag{15}$$

The difference between the two probabilities gives the marginal effect at the mean of y_{it-1} on the Pr(y_{it}=1). In the same context marginal effects are estimated for all independent variables. Standard errors were estimated using the delta method.

4. EMPIRICAL RESULTS

The estimates of the alternative models of the probability of participation in the labour market are given in Table 5. In order to check whether our results are sensitive to the choice of estimator six alternative estimators are presented: a static and a dynamic pooled probit and a dynamic random effects probit model, which assume that initial conditions are exogenous, as well as the Heckman's, Orme's and Wooldridge's estimators, which assume endogenous initial conditions. All models contain the variables listed in Table A.1 plus year dummies. Random effects models also contain time averages of time varying explanatory variables to account for any correlation with unobserved heterogeneity. Moreover, pre-sample information is used to model the initial conditions in the Heckman and Orme model[6]. A Wald test of their significance is presented in the bottom of Table 5 indicating in both models that they are jointly highly significant. The main variables of interest are that of the level of limitations in daily activities due to chronic health problems or disability, as well as that of previous period participation status.

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[5] In order to keep the formulas as simple as possible the vector $\overline{x}_{it}$ contains not only the exogenous explanatory variables but also the time means of all time varying explanatory variables, as well as the auxiliary terms used to model the initial conditions and $\hat{\beta}$ contains all the estimated coefficients.

[6] The instruments used to model initial conditions are an indicator of whether the individual was unemployed five years prior of the beginning of the survey and time indicators of when (s)he started working. When instruments are not included in the initial conditions equation identification of the Heckman and Orme model relies on non-linearities in the functional form. Some researchers (Andrén,2007 and Hansen *et al*, 2006) choose not to include instruments and rely on the functional form.

The key question is what the effect of disability is on labour force participation in the Greek labour market. To answer this we have included three variables capturing the effect on daily activities due to a chronic health problem. In general the effect of these disability variables is negative and quite high. In all dynamic models a chronic health problem without limitations on daily activities has no significant effect upon the probability of participation. On the other hand having a chronic health problem with some limitations has a negative and highly significant effect, while reporting severe limitations has a more adverse effect on the probability of participation.

In the static probit presented in column 1 the effect of all disability indicators is negative and significant and increases along with health problems. When dynamics are introduced in the model (column 2), the adverse effect of disability is reduced for those with severe and some limitations and becomes insignificantly different from zero for those with no limitations. Since random effects probit and pooled probit models use a different normalization, their coefficients are not directly comparable (Arulampalam, 1999). To compare coefficients between the random effects models and the pooled probit model the former must be multiplied by $\sigma_u / \sigma_v = (1-r)^{0.5}$. The scaled coefficients for those with severe limitations in columns 3 to 6 are -0.548, -0.558, -0.562 and -0.527 respectively, while the corresponding coefficients for those with some limitations are -0.200, -0.231, -0.246 and -0.233. This suggests that the coefficients of disability are further reduced when random effects are taken into consideration. It is worth mentioning that the rescaled coefficients are less than half of those of the static model. Since it is useful to estimate the exact effect of disability, as well as to avoid the need to scale all the coefficients to make reasonable comparisons, marginal effects are also estimated following the formula presented in the previous section. Such marginal effects take into account the necessary adjustment and exactly quantify the effect of each explanatory variable. Apparently since marginal effects are closely related to the coefficients of the estimated models they follow the same pattern. Thus, marginal effects are reduced when dynamics are introduced (column 2) and are further reduced when unobserved heterogeneity and initial condition are also modelled. In particular having some limitations reduces the probability of participating by around 6 per cent, while having severe limitations has a negative effect of 17.9 to 20.8 per cent. The marginal effect of having chronic health problems with some limitations turns out rather similar for all estimated models in columns 3 to 6. The marginal effect of the incidence of severe limitations from the Heckman model is slightly smaller (2 percentage points relatively to the rest). Moreover, the marginal effects of severe disabilities is much higher (in absolute value) than that of some limitations, in all six models. This finding is in accordance with the descriptive statistics of Table 3 and suggests that more severe health problems have stronger negative effect on the likelihood of participating. This difference seems to be consistent and does not change as unobserved heterogeneity and initial conditions are also included in the analysis. The results suggest that not specifying the model correctly will tend to severely overestimate the effect of disability on the decision to participate or not. It is necessary that one must take into account the effect of previous participation as well as unobserved heterogeneity and initial conditions. Ignoring these will lead to spurious results.

In all dynamic model specifications the previous participation status came out positive and highly significant, suggesting that previous participation increases the probability of current participation. One must notice that the marginal effect of the lagged dependent variable is defined in a "similar" way to the state dependence difference measure presented in Table 1 and measures the difference between the participation probability for those previously participating with those who were not in the labour force a year ago. The raw state dependence measure is very high around 80 per cent, while the corresponding estimated marginal effects is reduced to 73.3 per cent when observed characteristics are included (pooled probit). Moreover, when unobserved heterogeneity is introduced (random effects probit) the marginal effect of the lagged dependent variable is further reduced to 65.9 per cent. Finally, when the initial conditions are modelled (Heckman, Orme and Wooldridge models) the estimated state dependence measure ranges between 36.5 and 41.7 per cent. These estimates imply that state dependence is severely overestimated when all observed and unobserved characteristics as well as initial conditions are not taken into consideration. Even though there is substantial reduction, almost by half, of the positive effect of previous participation, it remains quite large implying that there is a considerable *ceteris paribus* dependence between previous and current labour market participation.

Comparing the models that assume endogenous initial conditions it turns out that for the Wooldridge estimator the coefficient and the marginal effect of γ is smaller than that of the other two. The Orme and Heckman estimators provide as far as γ is concerned results very close each other. However, all three estimates/marginal effects are not very different and highly significant. Another worth mentioning finding is that in all three models the cross-period correlation for the composite error term as estimated by r is fairly close together and much higher than that of the simple random effects probit[7]. This suggests that the equi-correlation between two periods is much lower for the model assuming exogeneity of the initial conditions. A final remark regarding the relative performance of the three models is that in terms of log likelihood the Heckman model is slightly superior, while in terms of correct predictions all three perform equally well.

The endogeneity of the initial conditions for models of columns 4 to 6 is checked by a *t*-test on the coefficient of the generalised error from the initial period for the Orme model, on the initial observation for the Wooldridge model and on θ for the Heckman model. In all three specifications initial conditions are highly significant and we cannot reject the hypothesis that they are endogenous and thus they should be taken into consideration in the estimations. In the Orme and Wooldridge estimators the initial conditions are incorporated in the models as auxiliary regressors. In both cases they have a significant positive effect on the probability of participating, suggesting that individuals who were participating in the beginning of the

[7] r can also be interpreted as the proportion of the total variance due to the individual level variance component. In all models a likelihood-ratio test to check whether r is equal to zero, i.e. individual level variance is no significant, was performed and give a p-value=0.000 in all instances.

examined period, are more likely to participate the following years. The size of the corresponding coefficient and marginal effect differs between models as a result of the different distributional assumptions made. Finally, it is worth mentioning that in the Wooldridge estimator participating in the base year increases participation probability by 36.6 percentage points, while the corresponding number for previous year participation is slightly smaller 36.5, suggesting that the effect of first year participation implies a longer effect.

Looking at the impact of the other explanatory variables one immediately notices the high negative effect of being female on the probability of participating. This is highly significant and fluctuates between 22.5 per cent for the Wooldridge model and 30 per cent for the Orme model. Age has significant positive effect on the probability of participating, which increases until the age of 45 and then declines even still positive, suggesting an inverse U shape between age and the probability of participating. Interestingly married individuals are less likely to participate. Higher education increases the probability of participation. The effect for those with high school degree is around 3 per cent in models accounting for the initial conditions while for university graduates it varies between 11 percent in the Wooldridge model to 17 per cent in the Heckman model. In all models individuals living in Attica display lower participation rates by 3.7 to 5.1 per cent. Finally, all random effects estimators suggest that the existence of children aged less than 12 years old in the household is insignificant. The same holds for unearned income, which is either insignificant or has a negligible effect.

As indicated above, following Mundlak (1978) and Chamberlain (1984), we assume a linear relationship between the unobserved heterogeneity ε_i and the time means of all time varying explanatory variables. This allows checking whether certain characteristics are associated with unobserved heterogeneity that reduces/increases the probability of participation. Regarding the means of the disability variables this of severe limitations is significant and negative in all models while this of no limitations is significant and negative in all models except the Heckman. Interestingly, the mean of some limitations even though it turns out negative it is insignificant in all models. These results indicate that the incidence of a health problem itself has a longer negative effect on the probability of participating through unobserved heterogeneity. Even the existence of a chronic health problem with no limitations is now correlated with certain unobserved characteristics that reduce the likelihood of participation. In particular the impact of this is quite large, between -20.3 to -22.7 per cent. It is important to notice that the effect of severe limitations through unobserved factors is stronger than the incidence of a health problem with severe limitations. These in combination with previous results on the coefficients of disability suggest that health problems play an important negative role in the decision of individuals to participate in the labour market.

Two final points worth mentioning regarding time averages is that marriage has now a positive and significant effect implying that it affects unobserved characteristics in a way that increases the probability of participating. The second point is that the amount of unearned income, which turn out insignificant in the χ-vector, is highly significant in the $\overline{x}$ -vector, while the existence of children under twelve is significant only in the dynamic random effects probit. Both have a negative impact and their marginal effect is rather similar in all models[8].

A general conclusion regarding the size of the independent variables is that the estimates of γ is reduced when unobserved heterogeneity is included and is further reduced when initial conditions are incorporated in the model. On the other hand the size of β and δ is higher for the random effects probit compared to the pooled probit and even higher for the models with endogenous initial conditions. This, along with the results for disability, suggest that not modelling for unobserved heterogeneity and initial conditions overestimates the effect of previous participation and disability and underestimates the effect of the other independent variables on the probability of participating in the labour market.

[8] In models estimated separately for women these variables and their marginal effects came out significant and with expected sign. Results are not presented but are available from the author upon request.

Table 5: Dynamic random effects probit models for labour force participation probability

	Static Pooled Probit (1)		**Dynamic Pooled Probit (2)**		**Dynamic RE Probit (3)**		**Orme estimator (4)**		**Wooldridge estimator (5)**		**Heckman estimator (6)**	
LFP at t-1			2.292***	[0.733]***	2.156***	[0.659]***	1.567***	[0.417]***	1.428***	[0.365]***	1.638***	[0.411]***
			(0.019)		(0.026)		(0.033)		(0.033)		(0.032)	
Generalized error from t=1							0.717***	[0.235]***				
							(0.033)					
Initial participation									1.419***	[0.366]***		
									(0.058)			
θ											1.035***	
											(.0566)	
Disability												
Severe limitations	-1.270***	[-0.474]***	-0.903***	[-0.341]***	-0.592***	[-0.201]***	-0.727***	[-0.205]***	-0.769***	[-0.208]***	-0.715***	[-0.179]***
	(0.037)		(0.047)		(0.069)		(0.075)		(0.076)		(0.075)	
Some limitations	-0.405***	[-0.150]***	-0.272***	[-0.095]***	-0.200***	[-0.063]***	-0.231***	[-0.061]***	-0.246***	[-0.062]***	-0.233***	[-0.058]***
	(0.028)		(0.036)		(0.051)		(0.056)		(0.057)		(0.056)	
No limitations	-0.258***	[-0.094]***	-0.040	[-0.013]	0.139	[0.040]	0.123	[0.030]	0.119	[0.028]	0.098	[0.025]
	(0.063)		(0.082)		(0.104)		(0.117)		(0.120)		(0.117)	
Female	-1.190***	[-0.386]***	-0.657***	[-0.210]***	-0.795***	[-0.234]***	-1.260***	[-0.305]***	-0.951***	[-0.225]***	-1.284***	[-0.232]***
	(0.016)		(0.020)		(0.031)		(0.044)		(0.040)		(0.045)	
Age 17-25	0.697***	[0.196]***	0.646***	[0.173]***	0.758***	[0.183]***	1.119***	[0.213]***	1.058***	[0.201]***	1.163***	[0.152]***
	(0.031)		(0.039)		(0.049)		(0.067)		(0.070)		(0.071)	
Age 26-35	1.172***	[0.311]***	0.766***	[0.211]***	0.955***	[0.236]***	1.449***	[0.281]***	1.254***	[0.247]***	1.429***	[0.193]***
	(0.026)		(0.033)		(0.045)		(0.061)		(0.061)		(0.061)	
Age 36-45	1.254***	[0.334]***	0.814***	[0.225]***	0.994***	[0.248]***	1.546***	[0.302]***	1.279***	[0.256]***	1.497***	[0.204]***
	(0.023)		(0.030)		(0.041)		(0.056)		(0.054)		(0.054)	
Age 46-55	0.932***	[0.267]***	0.602***	[0.174]***	0.736***	[0.193]***	1.155***	[0.241]***	0.954***	[0.201]***	1.115***	[0.163]***
	(0.021)		(0.027)		(0.035)		(0.047)		(0.046)		(0.045)	
Married	-0.182***	[-0.061]***	-0.156***	[-0.050]***	-0.377***	[-0.107]***	-0.446***	[-0.105]***	-0.435***	[-0.100]***	-0.442***	[-0.111]***
	(0.021)		(0.027)		(0.082)		(0.091)		(0.093)		(0.092)	
University	0.669***	[0.196]***	0.323***	[0.098]***	0.416***	[0.114]***	0.720***	[0.157]***	0.524***	[0.116]***	0.668***	[0.167]***
	(0.024)		(0.030)		(0.036)		(0.049)		(0.050)		(0.050)	
High school	0.072***	[0.025]***	0.028	[0.009]	0.054**	[0.016]**	0.129***	[0.032]***	0.105***	[0.025]***	0.109***	[0.027]***
	(0.018)		(0.022)		(0.026)		(0.034)		(0.036)		(0.035)	
Attica	-0.183***	[-0.064]***	-0.093***	[-0.031]***	-0.110***	[-0.034]***	-0.200***	[-0.051]***	-0.150***	[-0.037]***	-0.215***	[-0.041]***
	(0.017)		(0.022)		(0.026)		(0.036)		(0.039)		(0.038)	
Child < 12	-0.097***	[-0.034]***	-0.025	[-0.008]	0.069	[0.021]	0.041	[0.010]	0.031	[0.007]	0.033	[0.008]
	(0.019)		(0.024)		(0.051)		(0.056)		(0.057)		(0.056)	
Unearned income	-0.005***	[-0.002]***	-0.002***	[-0.001]***	0.001	[0.000]	0.001*	[0.000]*	0.001*	[0.000]*	0.001	[0.000]
	(0.000)		(0.000)		(0.001)		(0.001)		(0.001)		(0.001)	

Time-averaged characteristics												
Severe limitations					-0.978***	[-0.319]***	-1.235***	[-0.405]***	-1.337***	[-0.443]***	-1.222***	[-0.306]***
					(0.109)		(0.139)		(0.149)		(0.137)	
Some limitation					-0.085	[-0.028]	-0.078	[-0.025]	-0.084	[-0.028]	-0.052	[-0.013]
					(0.086)		(0.111)		(0.118)		(0.109)	
No limitations					-0.578***	[-0.189]***	-0.621**	[-0.203]**	-0.686**	[-0.227]**	-0.398	[-0.100]
					(0.203)		(0.265)		(0.282)		(0.263)	
Married					0.261***	[0.085]***	0.306***	[0.100]***	0.268**	[0.089]**	0.289***	[0.072]***
					(0.090)		(0.104)		(0.107)		(0.105)	
Child < 12					-0.157***	[-0.051]***	-0.116	[-0.038]	-0.119	[-0.039]	-0.111	[-0.028]
					(0.060)		(0.072)		(0.075)		(0.072)	
Unearned income					-0.006***	[-0.002]***	-0.009***	[-0.003]***	-0.007***	[-0.002]***	-0.008***	[-0.002]***
					(0.001)		(0.001)		(0.001)		(0.001)	
Constant	0.587***	[0.201]***	-0.971***	[-0.317]***	-0.792***	[-0.258]***	-0.338***	[-0.111]***	-1.146***	[-0.380]***	-0.297***	
	(0.029)		(0.039)		(0.052)		(0.068)		(0.072)		(0.069)	
σ_u					0.409	(0.031)	0.836	(0.030)	0.935	(0.031)		
r					0.143	(0.019)	0.411	(0.018)	0.466	(0.017)	0.456	(0.017)
Log-likelihood	-25,743.51		-16,680.11		-16,571.64		-16,146.13		-16,072.41		-15,418.63	
Sample size	44,755		44,755		44,755		44,678		44,755		55,505	
PCP	75		91		91		89		89		89	
x^2(df) [p-value]	15448.9(20) [0.0000]		33575.7(21) [0.0000]		16296.4(27) [0.0000]		10156.8(28) [0.0000]		8924.6(28) [0.0000]		8998.7(27) [0.0000]	
Wald test for instruments validity $x^2(3)$ [p-value]							732.98 [0.000]				502.12 [0.000]	

Notes:
(i) Standard errors shown in parentheses.
(ii) Marginal effects shown in square brackets.
(ii) All models also contain year dummies.
(iv) Models (1) - (6) estimated using observations for t>1 only.
(v) PCP: Percentage of Correct Predictions.
(vi) df: Degrees of freedom.
(vii) Log-likelihood for models (1) - (6) combined with wave 1 standard probits.
(viii) * p<0.10, ** p<0.05, *** p<0.01.
† Different sample size due to missing values in instruments used for initial conditions.

5. SUMMARY AND CONCLUSIONS

This paper examines the impact of the level of disability upon the individual labour force participation, taking into consideration state dependence, unobserved heterogeneity as well as initial conditions. The paper uses data from the Greek side of the European Community Household Panel and applies a range of different static and dynamic pooled and random effects estimators. The key findings are as follow.

The existence of a chronic health problem has an adverse effect on the probability of participating. This effect is higher for more severe health problems and influences participation directly (incidence of a health shock) and indirectly (unobserved heterogeneity), suggesting that unobservables are an important part of the model that should not be ignored.

It is necessary to take into consideration along with the dynamic nature of participation the unobserved heterogeneity. Different estimators suggest that unobserved heterogeneity is important and ignoring it tends to overestimate the effect of disability upon participation. In a static pooled model the effect of disability varies between 9.4 and 47.4 per cent depending on the level of disability. When dynamics are incorporated this is reduced to 9.5 to 34.1 and when unobserved heterogeneity is accounted for the effect of disability is further reduced to 5.8 to 20.8 per cent.

There is considerable state dependence in labour force participation, which is a result of both (un)observed heterogeneity as well as the incidence of previous participation. To correctly measure the latter, dynamic random effects probit models that control for observed and unobserved heterogeneity as well as for the initial conditions have been used. The paper has shown that ignoring unobserved heterogeneity and initial conditions results to considerably overestimate the effect of previous participation. In all estimated models the probability of participating is evidently higher, if the individual was participating the previous year, even after controlling for observed and unobserved characteristics. This fluctuates between 36.5 to 41.7 per cent, clearly indicating that there is genuine state dependence in labour force participation.

Certain other characteristics, such as age and higher level of education increase the probability of participating. On the other hand living in Attica and being married reduce the participation rate. Interestingly, the existence of young children in the household appears insignificant, while the amount of unearned income appears either insignificant or close to zero, but has a negative impact in the way it affects unobserved characteristics.

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APPENDIX

Table A.1: Variable definitions, means and standard deviations

Variable	Description	Mean	(SD)	Mean	(SD)†
Female	Female	0.521	(0.500)	0.525	(0.499)
Disable with severe limitations	Chronic health problem with severe limitations in daily activities	0.042	(0.200)	0.042	(0.200)
Disable with some limitations	Chronic health problem with some limitations in daily activities	0.063	(0.244)	0.064	(0.244)
Disable with no limitations	Chronic health problem with no limitations in daily activities	0.012	(0.108)	0.012	(0.108)
Age 17-25	Aged between 17-25	0.122	(0.327)	0.098	(0.298)
Age 26-35	Aged between 26-35	0.210	(0.407)	0.210	(0.408)
Age 36-45	Aged between 36-55	0.229	(0.420)	0.236	(0.425)
Age 46-55	Aged between 46-55	0.221	(0.415)	0.231	(0.421)
Married	Married	0.733	(0.443)	0.751	(0.432)
University	Holds a degree from university	0.164	(0.370)	0.159	(0.366)
High school	Higher level of education is high school	0.291	(0.454)	0.291	(0.454)
Attica	Lives in Attica	0.268	(0.443)	0.253	(0.435)
Child < 12	Children under 12 in the household	0.297	(0.457)	0.299	(0.458)
Unearned income	(Net household income - net individual disposable income)/100,000	24.79	(29.48)	25.45	(30.12)

Notes: (i). Pooled data from the ECHP waves 1-8 (1994-2001) (ii). Sample size unrestricted (minimum number of observations: 44,834, maximum: 55,505). † Sample size restricted to those with non-missing in all variables (44,755 observations)

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