Essays on the Labor Supply Dynamics of Older Workers

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Abstract

The European Union laid out the Lisbon agenda in order to become "the most competitive and dynamic knowledge-based economy in the world, capable of sustainable economic growth with more and better jobs and greater social cohesion".

The fulfillment of this ambition should not discard the introduction of policies enhancing the economic participation of the elderly. In fact, lower fertility rates and increasing life expectancy have triggered the aging process of European population and are deeply altering its demographic structure.

An aging society may exert larger pressure on the financial sustainability of pay-as-you-go pension systems owing to the decreasing age-profile of employment rates. On the one hand, the revenues coming from workers contributions to the scheme are expected to fall because a smaller fraction of individuals will be at work. On the other hand, the overall pension burden is expected to increase because a rising proportion of the population will be eligible for pension benefits.

In view of these considerations, European Union fosters Social Security reforms which assure long-run financial equilibrium and fully take into account the ongoing demographic modifications. At the same time, such institutional changes should be accompanied by public programs aimed at extending the working life of the elderly. Indeed, stimulating the labor market participation of older workers is an ulterior strategy to alleviate the financial burden of Social Security because many of them who no longer

find profitable the permanence in the labor market are likely to apply for retirement benefits.

The development of policies aimed at achieving these purposes requires the support of analyses examining the labor supply dynamics of older workers. Throughout the thesis we will focus on transitions out of employment by considering elderly individuals at work in a given time period and describing their probability of becoming not employed in the future.

In line with Blau (1994), we opt to estimate the dynamics of labor market position rather than imposing some arbitrary definition of retirement because the exit from the labor force may occur following alternative paths, such as becoming labor pensioners at the end of a full time job or starting an unemployment spell due to firm-downsizings that turn out to be a bridge towards *ad-hoc* early retirement arrangements. In order to preserve such heterogeneity, in this study workers are allowed to leave employment for whatever reason and to perform any possible employment trajectory.

Chapter 1 studies the impact of Information and Communication Technology (ICT) skills on the probability of ceasing from work. The point at issue is of particular interest because the Lisbon agreements place the improvement and the diffusion of ICT knowledge among the priorities of the Union. Our work draws data from the waves 2000-2004 of the Bank of Italy Survey on Household's Income and Wealth. Remarkably, the questionnaire of the 2000 survey provides unique information to measure ICT knowledge at the individual level. In fact, respondents are asked to report the utilization of a PC at work and to rank their computer literacy according to a predefined scale. This distinction is useful to disentangle the effects of actual individual investments in ICT training from those due to implicit job requirements reflecting firm organization. The econometric specifications adopted allow for the potential endogeneity of technological knowledge in a labor supply framework. In fact, older individuals who plan to work longer may be more willing to enhance their skills as they face a longer period over which training costs can be recouped. Improvements in ICT skills are estimated to produce a positive effect on employment chances only for males with high education. This pattern corroborates the hypothesis of complementarity between educational attainments and technological knowledge (Weinberg, 2004).

The literature provides a huge amount of evidence suggesting a positive relationship between individuals health and their own employment chances. Instead, few studies consider the role of the health conditions of other household members in the determination of the opportunity cost of keeping on working. Chapter 2 explores this research area by investigating how the labor supply dynamics of married workers is influenced by their own health conditions and by those of their cohabiting partners. Exploiting the European Community Household Panel (1995-2001), the overall psychophysical well-being of individuals is described by means of alternative indicators based on either self-assessments or more objective health indexes. Partners employment conditions are included among the explanatory factors because of their close relationship with health and the well-documented coordination between the labor market positions of couple members (Michaud, 2003). We also control for different sources of heterogeneity that may arise in our set-up and prevent us to estimate consistently the causal effects of interest. Our results confirm that healthier individuals are more prone to keep on working and, on the contrary, suggest that those living with a spouse in poor health are characterized by a higher propensity towards leaving employment.

Finally, Chapter 3 deals with the study of labor market transitions when the use of panel datasets involves a number of major complications, for instance due to severe attrition, small sample size, short time-span covered or lack of relevant variables. Following the approach proposed in Güell and Hu (2006), we estimate the individual hazard rate of stopping working only resorting to repeated independent cross-sections representative of the same population. Although workers are not observed over time, we are still able to evaluate their own likelihood of becoming not employed by combining the labor market outcomes experimented in different time periods by individuals of the same birth-cohort. Drawing data from the ISTAT survey Aspetti della Vita Quotidiana (1993-2003), we recover the evolution of the risk of leaving employment in a period characterized by several pension system reforms. We propose evidence that while the Amato (1995) and Prodi (1997) reforms passed, the cohorts of older workers at time show the highest hazard rate of leaving employment. On the contrary, their counterparts in 2001 and 2002 are more likely to prolong their working life. These findings suggest that, in the short run, repeated institutional changes designed to lower the Social Security generosity actually induce an increase in the exit rate from the labor force.

Abstract (in Italian)

L'Unione Europea ha sviluppato l'Agenda di Lisbona progettando di diventare l'economia basata sulla conoscenza più competitiva e dinamica, in grado di coniugare crescita sostenibile con occupazione di qualità e maggiore coesione sociale.

Realizzare questa ambizione non può tralasciare l'introduzione di politiche capaci di aumentare la partecipazione economica degli anziani. Infatti, tassi di fertilità più bassi che in passato e sempre migliori aspettative di vita hanno innescato un processo di invecchiamento della popolazione Europea modificandone profondamente le caratteristiche demografiche.

Una società che invecchia può esercitare una pressione più forte sulla sostenibilità finanziaria dei sistemi pensionistici *pay-as-you-go* a causa della relazione negativa che intercorre tra età e tassi di occupazione. Da un lato, ci si attende che i contributi pagati dai lavoratori diminuiscano perché minore sarà la frazione di persone occupate. Dall'altro, la complessiva spesa pensionistica è prevista aumentare come risultato della maggiore quota di popolazione idonea a richiedere l'erogazione del trattamento di quiescenza.

Alla luce di queste considerazioni, l'Unione Europea incoraggia l'avvio di riforme pensionistiche che assicurino l'equilibrio finanziario di lungo periodo e prendano seriamente in considerazione le trasformazioni demografiche in corso. Allo stesso tempo questi cambiamenti devono essere accompagnati da programmi pubblici mirati ad allungare la vita lavorativa degli anziani. Infatti, incentivare la partecipazione al lavoro in questa fascia di età è un

ulteriore modo di consolidare finanziariamente il sistema dato che una larga parte dei lavoratori anziani che esce dal mercato del lavoro richiede i benefici pensionistici.

Lo sviluppo di politiche finalizzate a raggiungere questi obiettivi richiede il supporto di analisi miranti a descrivere la dinamica dell'offerta di lavoro degli anziani. La mia tesi si concentra sulle transizioni verso la non occupazione considerando persone di almeno mezza età al lavoro in un certo istante di tempo ed esaminando la loro probabilità di smettere di lavorare in futuro.

In linea con Blau (1994) decidiamo di stimare l'evoluzione della condizione lavorativa effettiva piuttosto che adottare uno specifico modello di pensionamento perchè la transizione verso l'inattività economica avviene attraverso molteplici modalità, come il cessare un rapporto di lavoro a tempo pieno per diventare pensionato o l'incominciare un periodo di mobilità che in realtà traghetta verso il pensionamento anticipato attraverso appositi dispositivi legislativi. Al fine di conservare tale eterogeneità, questa tesi considera transizioni verso la non occupazione avvenute per qualunque motivazione ed al termine di qualsiasi percorso occupazionale.

Il primo capitolo studia l'impatto della conoscenza delle nuove tecnologie dell'informazione e della comunicazione (ICT) sulla probabilità di smettere di lavorare. Il punto in questione è di grande importanza perché gli accordi di Lisbona pongono lo sviluppo e la diffusione di queste conoscenze tra le priorità dell'Unione. La nostra analisi utilizza i dati 2000-2004 dell'Indagine sui Bilanci delle Famiglie Italiane realizzate dalla Banca d'Italia. E' importante notare che il questionario della *wave* 2000 fornisce interessanti informazioni per misurare la conoscenza delle ICT a livello individuale. Infatti, agli intervistati viene chiesto di riportare l'eventuale utilizzo del PC al lavoro e di classificare la loro abilità nell'utilizzo del PC secondo una scala predefinita. Questa distinzione è utile per separare gli effetti dovuti agli investimenti nella conoscenza fatti da una persona da quelli invece prodotti da assetti organizzativi impliciti del posto di lavoro. La strategia econometrica utilizzata ammette che ci possa essere endogeneità delle abilità rispetto alla situazione occupazionale. Infatti, coloro che progettano di lavorare più a lungo, possono essere più incentivati a riqualificare le loro conoscenze ICT perché hanno di fronte un orizzonte temporale più ampio per recuperare i costi derivanti da questa scelta. Le nostre stime documentano che per gli uomini con istruzione più elevata, la capacità di utilizzare le ICT si traduce in una maggiore probabilità di rimanere occupato. Questa evidenza empirica avvalora l'ipotesi di complementarietà tra istruzione e conoscenze tecnologiche (Weinberg, 2004).

La letteratura economica conta svariati studi che dimostrano l'esistenza di una relazione positiva tra la condizione di salute di una persona e le sue possibilità nel mercato del lavoro. Al contrario, poche analisi considerano il ruolo svolto dalla salute degli altri componenti della famiglia nella determinazione del costo-opportunità di continuare a lavorare. Il secondo capitolo esplora questa area di ricerca esaminando come la dinamica occupazionale dei lavoratori sposati sia influenzata dalle proprie condizioni di salute e da quelle dei loro coniugi. Sfruttando lo *European Community Household Panel* (1995-2001), il complessivo stato psicofisico delle persone è descritto da indicatori basati su autovalutazioni o informazioni più particolareggiate sulla loro salute. La condizione lavorativa dei coniugi è inclusa tra le variabili esplicative per via dello stretto legame con la loro stessa salute e il noto coordinamento esistente tra le scelte lavorative dei componenti di una coppia (Michaud, 2003). Inoltre, le specificazioni adottate tengono conto delle varie fonti di eterogeneità non osservata che possono sorgere nel nostro contesto ed impedirci di stimare in maniera consistente gli effetti causali di interesse. I risultati confermano che le persone più in salute risultano più propense a continuare a lavorare e, al contrario, mostrano che coloro che vivono con un coniuge in cattiva salute presentano una maggiore probabilità di uscita dallo stato di occupato.

Infine, il terzo capitolo esamina lo studio delle transizioni tra diversi stati occupazionali quando l'uso di *dataset* longitudinali è sconsigliabile a causa di *attrition*, bassa numerosità campionaria, breve periodo temporale coperto e mancanza di variabili rilevanti. Seguendo l'approccio proposto in Güell and Hu (2006), stimiamo a livello individuale il rischio di uscire dallo stato di occupato ricorrendo solo a *dataset* trasversali ripetuti e rappresentativi della stessa popolazione. Nonostante i lavoratori non siano seguiti nel tempo, il loro rischio di diventare non occupati è calcolato combinando i risultati occupazionali caratterizzanti le persone della stessa coorte in diversi istanti temporali. Utilizzando i dati ISTAT dell'indagine Aspetti della Vita Quotidiana (1993-2003), ricostruiamo l'evoluzione della probabilità di smettere di lavorare lungo un arco di tempo nel quale hanno avuto luogo molti cambiamenti nel sistema pensionistico. I nostri risultati mostrano che mentre le riforme Amato (1995) e Prodi (1997) venivano introdotte, le coorti di lavoratori dell'epoca hanno il rischio più alto di transitare verso la non occupazione. Al contrario, i loro omologhi negli anni 2001 e 2002 sono associati ad una maggiore probabilità di continuare a lavorare. Questa dinamica porta a pensare che, nel breve periodo, ripetute trasformazioni del sistema pensionistico finalizzate a ridurne la generosità portino ad un incremento delle uscite dalla forza lavoro.

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Chapter 1

Technology, Skills and Retirement

1.1 Introduction

Our main research question is to investigate empirically whether individuals who are more "technologically endowed", *ceteris paribus*, tend to retire later[1]. After a decade or more of intense research on skill-biased technological change, we believe that it is possible to conclude that observed and unobserved skills are among the most important determinants of workers' wages and employment status. While they can be modelled quite easily from a theoretical standpoint, skills are hard to measure empirically. Workers' productivity within occupations is likely to depend upon many factors, some observables and some not. Talent, which is only very partially observable, is a major determinant, but we should not overlook the importance of training, experience, firm organization and technological progress.

Skills are likely to influence retirement choices as well. In particular, given the fast diffusion of ICTs across sectors and professions, workers with

[1] This is a joint work with Federico Biagi (University of Padova) and Raffaele Miniaci (University of Brescia).

poor “technological endowments” tend to become less and less productive, particularly in industries and professions characterized by rapid technological progress. This might lead to lower expected wages and worse expected job conditions, making (early) retirement preferable. At the same time, if human capital and technology are complementary, we cannot disregard the possibility that a skill-biased technological change may favor more experienced (and older) workers because of their higher level of human capital accumulated (see Weinberg, 2004) and that this effect might vary across education levels. Hence, the prediction of the sign and the size of the effect of “technological endowments” on the probability of retiring earlier is mainly an empirical issue. In this analysis we should be aware that the ability to cope with technology is structurally different from the actual use of technology on the job. In fact, ability in dealing with technology is a valuable asset by itself, being an indicator of a more general ability to cope with changes affecting job tasks. On the other hand, the use of technology at work might just be an implicit job requirement, which does not necessarily create extra value added once we control for job characteristics.

In this paper we are able to disentangle the effects of technological skills from those arising from the use of technology on the job. In our work we focus on the retirement decisions of Italian employees aged 45-70[2]. Data are drawn from the 2000-2004 panel section of the Bank of Italy Survey on Household Income and Wealth (SHIW)[3]. SHIW turns out to be a unique source of information for the Italian case because of its panel component and since the wave 2000 contains questions about computer literacy and computer use at work of each household member.

Our results indicate that Italian male employees with higher education

[2] In our context, *retirement* actually labels the transitions out of employment. Since older individuals not at work in Italy have a small probability of finding a new job (see Table B.4), the end of an employment spell in this population group is likely to entail the retirement from the labor force.

[3] Data utilized in this chapter are publicly available at www.bancaditalia.it.

who use a computer at work tend to retire later, and the magnitude of this effect is remarkably larger than the one observed in US and Germany in previous studies. Moreover, we provide clear evidence that -for males employees with high education- the ability in the use of computers is a factor affecting long run retirement outcomes even if a PC is not used at work. In other words, if ability in the use of PCs can be considered a good proxy for individuals' technological capabilities, we find that this factor affects the retirement decision of males with higher education only. On the contrary, there is no evidence of any significant effect for female employees. These results are robust to the econometric strategy adopted.

The chapter proceeds as follows. Section 1.2 briefly reviews the literature on retirement choices and skill-biased technological change; Section 1.3 provides *prima facie* evidence of the relation between PC utilization and retirement based on our SHIW dataset; Section 1.4 discusses the results of our alternative econometric strategies; Section 1.5 draws the conclusions.

1.2 Skills and retirement: literature review

The classical economic approach to modelling retirement decisions is based on the assumption that individuals choose whether or not to retire by comparing the present value of the streams of benefits and costs occurring in the two cases (see Lazaear, 1986 and Lumsdaine and Mitchell, 1999 for a complete review of the literature). Within this framework, preferences for leisure, actual and expected levels of labor and non labor income, pension benefits, pension tax contributions and health conditions play a crucial role. We can approximately group the empirical works on retirement in two sets: the first is inspired by the work of Gustman and Steinmeier (1986), Rust (1989), Rust and Phelan (1997) and Stock and Wise (1990) who propose structural dynamic stochastic models; the second one relies upon reduced form models of the conditional probability of retirement, and does not impose any restriction on individual preferences (see Meghir and Whitehouse, 1997). Models belonging to this class typically focus on variables such as the accrual rate, the implicit tax rate or the option value of working an additional year.

The analysis of retirement behavior shows that individuals tend to retire at around the standard age of eligibility, with some retiring happening at an earlier age when early retirement is allowed (see for instance Blundell et al., 2002 and Tanner, 1998), and that institutional details of the Social Security system do matter (see Coile and Gruber, 2000 and 2001)[4].

Few empirical papers focus on the role of technological skills in the retirement decision. In this case, the econometric analysis is complicated by the fact that the relation between skills and retirement is affected also by the business cycle and by the training policy adopted by firms[5].

Bartel and Sicherman (1993) study the effect of technological change on the career of older workers. They notice that technological change can affect retirement, influencing both the training decisions and the depreciation of the stock of human capital. In particular, they test two complementary hypotheses. According to the former, everything else constant, individuals retire later in industries in which technological change is particularly rapid. According to the latter, an unexpected rise in the depreciation rate of human capital, for instance following an unexpected rise in the rate of technological change, should lead to earlier retirement. Both hypotheses are confirmed by their analysis, whose main limitation is in the use of sector data to measure technological change.

Friedberg (2003) tries to provide an answer to a question very similar to the one analyzed in our paper. In fact, she investigates whether there

[4] For recent results on the Italian case see the special issue of Labour, August 2003, 17.

[5] Notice that rapid technological change has two effects of opposite sign on the training of older workers. On the one hand, it makes training more profitable, but, on the other one, the speed of technological change increases the depreciation rate of human capital and hence reduces the incentives to train, especially for older workers, since a long period at work is required to make training beneficial for both parties.

exists evidence of a significant relationship between computer use and retirement. The basic intuition for her analysis is that computers have affected the demand for labor in various ways. First, they tend to be a substitute for unskilled labor and routine tasks. Second, they have altered the performance of non-routine tasks, mainly held by skilled workers. Finally, computerization alters the "bundle of skills and tasks that define a job". These changes can affect the retirement choice of older individuals, given that older generations tend to be less educated and hence more likely to be assigned to routine jobs. For these workers training may be generally less profitable given the higher investment costs and the reduced time horizon over which they can be recouped. Friedberg uses the US Health and Retirement Study (from 1992 to 1996) to study how the frequency of computer use at work affects the transitions towards retirement of workers aged 50-62 in 1992. She takes into account the possible correlation between the use of the computer and the unobserved propensity to retire later by estimating a linear probability model using an instrumental variable approach. In particular, she opts to instrument the use of a PC by an individual with the percentage of computer users among prime-age workers in the same occupation and industry. Her findings show that in the long run case (that is, over a four-year horizon), even controlling for many individual, firm and sector characteristics, computer use tends to induce delayed retirement. She concludes that "holding everything else constant, the median retirement age if everyone had used a computer would have occurred 12 months later".

In a very similar framework Schleife (2006) uses the German Socio Economic Panel to investigate the effect of computer use at work on the retirement outcomes of employed males aged 50-60 in 1997. As Friedberg (2003), she models retirement by means of a linear probability model, distinguishing between different timing of retirement (transitions occurring within 1999 and those occurring within 2001 are studied separately). The potential endogeneity of computer use on the job is addressed by means of an instrumental variables approach that considers computer use at home as the additional instrument. Her IV results provide no evidence that Germans who use a computer at work tend to postpone their retirement.

Our paper is strictly related to Friedberg (2003) and Schleife (2006) in the use of microdata and the focus on the effect of the use of PCs on workers' retirement decisions. However, we depart from their analyses by distinguishing between the use of a computer at work and the level of computer literacy. On the one hand, if the use of a PC at work is a requirement of a certain job profile, we incur the risk of imputing to the use of PCs effects that are actually due to structural organizational changes. On the other hand, the degree of computer literacy may effectively proxy ability to react to (past and future) innovations. Hence, we expect individuals with computer skills to be more valuable for the firm and *ceteris paribus* less likely to retire.

However, both computer utilization and computer literacy are likely to be endogenous in an equation describing retirement decisions: workers who plan to retire later, have *ceteris paribus* more incentives in investing in an upgrade of their skills. We address the endogeneity issue by using an IV approach, exploiting the information on the computer literacy of the members of the workers' household. The validity of these instruments relies on the evidence that workers living in a family where the other members are computer literate are themselves more likely to be able to use a computer and to use it at work (see Miniaci and Parisi, 2006) and on the assumption that living with people with a good knowledge of ICTs affects the individual retirement decisions only through the level of her own technological skills.

1.3 Data description

Data are drawn from the Bank of Italy Survey of Household Income and Wealth (SHIW), which, every two years, provides a sample of about 8,000 households, representative of the Italian population. It contains detailed information on demographics, income and wealth at the individual and household level[6]. The 2000 wave provides us with information on the ability of individuals in the use of computers and, for those who are working, on their use of a PC at work. Specifically, for each household member, it records self-rated computer skills on an increasing five-step scale. We rearranged this scale in order to define a dichotomous variable that takes a value of one if the individual declares to have at least some ability in PC utilization and zero otherwise[7]. For people at work, the 2000 survey also collects information concerning whether they use a PC at work. Since half of the households participating to the sample belong to a longitudinal survey, we exploit the panel section coming from the 2000, 2002 and 2004 waves. These data provide unique and valuable information for the question analyzed in this paper, since they permit us to estimate whether computer skills and computer use at work are positively influencing the probability of remaining employed.

In our work attention is focused on employees. In particular, we consider household's heads and their spouses who are employees and aged 45-70 in 2000. All the following tables refer to this group, if not otherwise specified. The percentage of "PC skilled" workers[8] goes from 46.6% for the 45-49 age group to 25.4% for those 60+ (see Table A.1). The survey documents a remarkable gap between the North and South of Italy, and strong differences between workers with at most compulsory education and those with higher education.

Given our aim, it is crucial to distinguish between workers who are com-

[6] See Banca d'Italia (2002) for further details.

[7] While this reduces the variability in our explanatory variable, it also has the effect of reducing the measurement error due to the fact that individuals self-evaluate their skills. In addition, adopting finer partitions produces severe multicollinearity problems in the econometric specifications presented in the following sections.

[8] These are workers for which the dichotomous variable previously defined takes a value equal to one.

puter literate and those who do use a computer at work. As reported in Table (A.2), according to SHIW[9] only 31% of workers in our reference group were using a computer on their job in 2000 and only 73% of skilled individuals use their computer skills at work.

Before starting to analyze how computer skills and computer use affect individuals' retirement, it is necessary to recognize that such a process is heterogeneous and this makes it difficult to develop a tight definition of retirement. For some individuals the labor supply decision is well represented by a dichotomous choice between working full time and abandoning the labor market altogether. This is the case of a typical retirement pattern by which an individual, once eligible for Social Security benefits, chooses to exit from the labor force. Alternatively, individuals may experience a smoother retirement process: despite the fulfilment of some eligibility criteria, they may choose to remain employed reducing progressively the amount of hours worked. In such a case, although a pension benefit might be formally withdrawn, these individuals should not be considered as retired from the labor market. Finally, retirement may be a by-product of a firm downsizing. In fact, this process may include an institutional arrangement that provides unemployment benefits to workers aged beyond a chosen threshold until they become eligible for Social Security. Although these workers result to be formally unemployed, they do not have any incentive to look for another job and hence they are substantially out of the labor force.

In our empirical analysis we look at two definitions of retirement. Both are based on self-reporting and hence exposed to the risk that the same expression means different things to different individuals. According to the

[9] Miniaci and Parisi (2004) document that the estimate for the percentage of computer users in the population based on SHIW is consistent with the one that can be obtained using the much larger ISTAT Multiscope Survey: according to the latter, at least 13.5% of the overall population uses a computer at work, while, according to the SHIW, this percentage amounts to 12.2%. The Bank of Italy survey tends to slightly underestimate the proportion of skilled individuals among the young and to overestimate it among people over 45 years old.

first definition (strict definition) we focus only on individuals employed in year 2000, who in the following years declare themselves as job pensioner. According to the second definition (broader definition), we consider retired those who leave the initial state of employment for whatever reason.

For those workers who are retired according to the strict definition, the 2002 SHIW wave contains the self-reported information on the retirement age. Its distribution peaks at the typical ages of 55 and 60 for, respectively, females and males, confirming the evidence obtained on the basis of the Social Security administrative dataset by Brugiavini and Peracchi (2003).

In Table (A.3) we consider the two definitions of retirement and distinguish between transitions occurring within 2002 and those occurring within 2004. The total number of observations is reduced with respect to Table (A.1) and Table (A.2) because here we focus on individuals belonging to the panel section of the survey. We point out that the number of individuals followed up to 2002 is larger than that of those followed up to 2004. This is due to the fact that the survey has the sampling design of a rotating panel and that some attrition might also be at work.

Overall, if we consider the broader definition of retirement, 18.7% (15.1%) of male (female) employees aged over 45 in 2000 retired by the end of 2002 and a further 11.7% (14.9%) of employees[10] still working in 2002 retired between 2003 and 2004 (assuming that participation to the survey is independent of the retirement behavior). There are remarkable differences between workers with ICT skills and their colleagues. In the 2000-2002 period while for females the fraction of unskilled workers retiring is 50% higher than that of their skilled counterparts, it is more than 100% higher when we consider male workers. A gap of similar magnitude is observed for females also in the longer time interval, while for male workers the gap is slightly reduced. Analogous patterns are found when we consider PC utilization at

[10]For males, this percentage comes from (28.2-18.7)/(100-18.7).

work.

When we focus upon the strict definition of retirement, which implies a potentially endogenous selection of the individuals in the sample, only for female workers we record noteworthy differences with respect to the previous case. In particular, the differences between skilled and unskilled females are almost negligible in the 2000-2002 interval.

The relationship between retirement and computer skills is then explored by means of non-parametric estimates of the survivor functions obtained stratifying by the variable of interest. We test the hypothesis of equality of the survivor functions between groups using the suitable log-rank tests.

In Figure (A.1) we consider the broader definition of retirement and, for the sample of male workers with at least secondary education, we compare the survivor functions of those with and without computer skills. Individuals with some ability in the use of PCs exhibit a higher probability of remaining employed up to age 64, and the log-rank test confirms that the difference is statistically different from zero. The previous results are not confirmed when we focus on males with lower education, as reported in Figure (A.2).

Here the curves cross around age 58 and no clear pattern emerges. Consistently, the log-rank test does not reject the null hypothesis of equality. Similar results hold when we consider computer use at work.

As for females, Figure (A.3) shows that computer literate women tend to leave their job later. In spite of this evidence, the log-rank test does not reject the null hypothesis of equality, even stratifying the sample by education. This result is due to the small size of this sample and it is confirmed also when we focus on the PC utilization at work.

When the strict definition of retirement is taken into account, all these results are confirmed.

1.4 Multivariate analysis

The evidence discussed so far does not take into account that PC skills and PC use on the job may be correlated with other potentially relevant determinants of the retirement choice, such as the type of job, the sector of employment, firm dimension, labor income, income from other sources and the demographic characteristics of the worker's household. Therefore, we run two types of multivariate analysis. In the first, as Friedberg (2003) and Schleife (2006), we use discrete choice models in order to describe how the probability of transition towards retirement is affected by computer utilization at work and/or computer skills. In the second, we investigate the same research question by estimating a semiparametric duration model.

1.4.1 Discrete choice models

In our first exercise, following Friedberg (2003) and Schleife (2006), we analyze the effect of the use of a PC at work on the probability of retirement, distinguishing between changes occurring within 2002 and those occurring within 2004. In all the specifications we control for age, age squared, education, number of days spent at home for illness during 2000, region of residence, number of household components, labor income and other household's income. As proxies for experience and pension wealth, we use respectively the age at the time of the first job and the number of years of contribution to Social Security up to 2000. Moreover, we control for job characteristics, sector of employment and firm dimensions in order to allow for variation in the rate of diffusion of new technologies in the economy.

When we estimate these models, we should take into account that OLS results might be biased due to a potential endogeneity of the use of a computer at work as well as computer literacy among workers. Endogeneity may arise from the fact that individuals who use a computer have unobserved abilities which make them more likely to continue working. Moreover, individuals who plan to retire later might decide to use a computer in order to increase their skills and hence their future employment probability. Hence, they would have a stronger incentive to improve their ICT skills because they expect a longer period of permanence in the labor market, in which the benefits from their training investments can be recouped.

Aware of this, we estimate a linear probability model also by means of an instrumental variable approach. The set of instruments we use in order to achieve the identification of the parameters of interest consists of the number of other family members with some computer skills, its interaction with the number of household components, its interaction with education and, finally, the number of children at school in the household.

Our exclusion restrictions can be summarized as follows: the skills of other household members may affect the skills of the worker, but, once controlled for her actual computer abilities, other household members skills do not affect her retirement decision. Miniaci and Parisi (2006) show that within-household peer effects are indeed relevant for the diffusion of computer skills: considering the 2000 SHIW wave, 70% of individuals co-habitating with somebody skilled are skilled and this percentage falls to less than 10% if nobody else in the family is able to use a computer.

Table (A.4) reports the estimates of the effects of using a computer at work on the transition towards retirement in the two-year period 2000-2002 and in the four-year period 2000-2004 (for both definitions of retirement). For sake of brevity, in the text we show only the estimates of the main parameters of interest. The effects of the control variables present the expected sign. They confirm that the likelihood of retirement increases with age and with cumulated pension wealth. Moreover, public sector employees tend to retire later. Furthermore, once we control for pension wealth, job characteristics and computer utilization, our estimates show that the educational attainment does not play a prominent role.

The OLS results for the two year period confirm the descriptive evidence of Table (A.3). In fact, male workers who use a computer at work retire significantly later than their companions, while for females the difference is negligible. In particular, the reduction of the retirement probability is equal to 8.6 percentage points if we consider the broader definition of retirement in the short period. This amounts to half the difference in the average probability of retirement between users and non users documented before (see Table A.3). A similar proportion is found for the four-year interval.

Instrumental variable point estimates are almost equal to the OLS ones in the two-year interval, but they more than double in the 2000-2004 period, where the difference between users and non users is significant even for females if we refer to the broader definition of retirement.

We test for the joint significance of the additional instruments in the first stage equations and we always reject the null hypothesis of insignificance. Further, we test the validity of the additional instruments and always accept the null of validity of the exclusion restrictions. The Hausman exogeneity test[11] marginally rejects the null in three of the eight cases we consider. To summarize, we have weak evidence that using an instrumental variable approach is necessary in our case.

Our results are in line with Friedberg (2003), who refers to a sample of US workers (either employee or self-employed, males and females together) aged 50-62 in 1992. In fact, although in her study the raw difference in retirement probability between users and non users is about one third of the one we estimated for our sample, her OLS estimates of the effect of using a computer on the retirement probability amount to about half of it. We are also consistent with her results when comparing IV and OLS estimates since her IV estimates are almost three times larger than the OLS ones. Instead, we find remarkable differences with Schleife (2006), who, based on a sample

[11] The validity of both overidentifying restrictions and exogeneity assumption is tested allowing for heteroskedasticity.

of German male workers (either employee or self-employed) aged 50-60 in 1997, finds no evidence that computer use induces to postpone retirement.

At the moment we are not in the position to disentangle a potential skill effect from the effect of actual computer use. In what follows we enrich the previous specification by introducing a dummy variable equal to one if the worker has some computer skills as well as its interaction with the low-education dummy.

Since PC users at work are by definition PC skilled, we then define the overall effect of using a PC at work as the sum of the effect of having PC skills and the one of using a PC at work taken *per se*. We also point out that now the parameter on the dummy for computer use measures the additional effect of using a PC at work for a PC skilled individual. Finally, it is worth noting that plugging in the specification the interaction between PC skills and the low-education dummy allows the impact of being PC literate and, in turn, the overall effect of using a PC at work to vary across education groups[12].

As in the previous case, both OLS and IV estimates are considered. Table (A.5) shows that in the two-year period (2000-2002) the OLS estimates of the overall effect of PC utilization at work are significant and negative for males, irrespective of their education. On the contrary, being computer skilled produces significant and negative variations in the probability of retirement only for low education individuals who quit employment to become job pensioners. Further, it should be noticed that once we condition on PC skills, the additional effect of using a PC on the job is significant and strikingly positive only for females who retire according to the strict definition. As illustrated by Table (A.6), in the longer period (2000-2004) we find significant results only for high education males. In particular, both the overall effect of using a PC on the job and the one of being computer literate are

[12] On the contrary, the additional effect of using a PC at work is assumed not to depend on education levels.

negative and significant. Again, it is worth noting that once PC skills are allowed for, the additional effect of using a PC at work is never significant.

These results are weakened when we switch to the instrumental variable approach. This lack of coherence is probably due to the considerable increase of standard errors, which causes a loss of precision in the estimates. However, we tested the validity of both the additional instruments and the exclusion restrictions finding no evidence of misspecification. Finally, the Hausman tests always accepted the null hypothesis of exogeneity, suggesting that there is little advantage in using an IV approach and that we can rely on OLS estimates.

If exogeneity is not an issue, more efficient estimates are obtainable by adopting alternative discrete choice approaches. Table (A.7) and Table (A.8) report the ML estimates of the logit model for a specification with the same covariates as the previous ones and the results of the Rivers-Vuong (1988) tests of exogeneity. Even in this case we never reject the hypothesis of exogeneity. Therefore, we can compare these ML estimates with the previous OLS results to see that almost all the findings of the linear probability models are confirmed. In the two-year period the overall effect of using a PC at work is negative and significant for males, irrespective of their educational attainment and the definition of retirement considered. On the contrary, being PC literate produces a decrease of the probability of leaving the initial state of employed only for high education males in the strict definition case. Like before, once PC skills are allowed for, the additional effect of using a PC at work is always statistically negligible. In the four year period most significant results are found only for high education males. As it is shown, both the overall effect of using a PC on the job and the impact of being PC literate are negative and significant. Finally, we point out how the additional effect of using a PC at work, once we condition on PC skills level, is not significant.

1.4.2 Semiparametric duration analysis

So far we have modelled the retirement process as a dichotomous choice to be made in a given time horizon (two or four years). However, retirement can be more properly studied exploiting the tools of the survival analysis since it is a decision process mainly concerned with the choice of the optimal timing of exit from the labor force. Therefore, we estimate a semi-parametric Cox model conditioning on the same set of control factors used in the previous subsection. As our dataset has a panel structure, we exploit the possibility provided by SHIW of updating the control variables as time elapses. The covariates measured as of year 2000 are used for studying transitions between 2000 and 2002, while the information conveyed by their updates to 2002 values is exploited for transitions between 2002 and 2004. In other words, we use a duration model with time-varying covariates. Unfortunately, the information on PC skills and PC use at work is available only in year 2000.

We decide to model the duration in the initial state of employment using Cox specifications and handling tied events of exit by means of the exact partial likelihood method[13]. Basically, the outcome of interest is the discrete-time hazard rate of retirement, which is defined as the probability of retiring at age a conditional on arriving employed at age $a-1$. Further, we assume the proportionality for the odds of the hazard rate. This amounts to saying that the odds of retiring result from the product between a baseline odds common to all individuals and a function summarizing individual specific characteristics. Hence, the parameters we intend to identify measure the effect of the explicative variables on the odds and, in turn, on the hazard. Cox models are particularly suitable because they can be rewritten as logit regressions, similar to those considered in the previous section. This

[13] See Thernau and Grambsch (2000) for further details concerning the methods for handling tied events.

allows us to maintain the validity of the Rivers and Vuong (1988) exogeneity test conducted for the previous discrete choice analysis.

As reported in Table (A.9), most significant results are found for high education males. While being computer literate decreases the likelihood of retirement only in the strict definition case, the overall effect of using a PC at work is found to be always significantly lower than zero. It should be noticed that, conditional on PC skills, the additional effect of using a PC at work is significant and negative for all males who retire to become job pensioners.

Looking at the shape of the baseline hazard rate reported in Figure (A.4) it is easy to recognize peaks at the typical retirement ages. As expected, male employees tend to retire around the age of 57 and 65 because of the entitlement rules for the seniority pension (*pensioni di anzianità*) and old age pension (*pensioni di vecchiaia*). These results confirm the evidence given for Italian workers by Brugiavini and Peracchi (2003) using the Social Security administrative dataset.

1.5 Conclusions

We empirically investigated the relation between technological skills and retirement choices of Italian employees aged 45-70 in 2000. We exploited the longitudinal structure of the Bank of Italy Survey on Household Income and Wealth, which provides information on a wide set of both individual, household and job characteristics. Among these we have variables capturing at the individual level the ability in the use of a computer and the actual use of a PC at work.

Two definitions of retirement are taken into account. According to the former, we consider retired individuals who leave the initial state of employment to become job pensioners (strict definition). According to the latter, retirement occurs if employment status is left for whatever reason (broader definition).

The main findings of our analysis refer to specifications which disentangle the effect of using a PC at work from that of being PC literate. In other words, keeping in mind that users of a PC on the job are by definition PC skilled, we are able to distinguish the additional effect of using a PC at work conditional on PC skills level from the overall effect of using a PC on the job, which is given by the sum of the impact of being PC literate and the one of using a PC at work taken *per se*.

All our results highlight that, overall, using a PC on the job entails a lower probability of retirement for male employees with high education. This evidence is invariant to the time horizon considered (two or four years), to the definition of retirement adopted, and to the choice of the modelling strategy (linear probability models, logit models and semiparametric survival analysis). They are weakened only in the case of instrumental variable estimates of the linear probability models. However, we document that, both in the linear probability model and in the logit model, endogeneity is not a major concern.

This result is consistent with the hypothesis that computer skills and education are complements and that the diffusion of personal computers in the economy strengthens the demand for highly skilled labor. According to this view, only high-education workers are able to fully exploit ICT technologies and hence more likely to find the conditions, for instance in terms of expected wages, that make preferable postponing retirement. We also show that at least for the broader definition of retirement, once computer skills are allowed for, the additional effect of using a computer at work is no longer significant. Furthermore, our estimates show that computer skills and computer utilization at work do not seem to play a crucial role in the retirement process of women.

In conclusion, we obtain two main results. First, there is a relevant heterogeneity on how ICT adoption affects different groups of older workers. It is only for the well educated male employees that the lack of computer skills may result in early retirement. For low-education males and for females, who tend to be concentrated either in manual job positions (employees with low education) or in selling, caring or teaching activity (women with high education), there is no effect of ICT skills on the retirement decision. Viceversa, given that more than 50% of males with high education use a computer at work and more than 65% are able to use it, the lack of computer skills among this group is such that not being able to use a PC is a strong signal of the obsolescence of human capital. Second, it is important to recognize that for males, when the relation between computer utilization and retirement is investigated but skills are not taken into account, the effect of computer utilization on the retirement process is likely to be overestimated. In fact, when we do not control for PC skills, the variable that captures PC use at work soaks the effects of both "technological skills" and their use on the job. Consistently, when we control for the level of the skills, the overall effect of the PC utilization on the job is found to be significant, both statistically and economically, but in almost all cases the additional effect of the use of a PC at work is no longer relevant.

These results resemble the findings of the empirical literature on the computer wage premium. The first evidence (Krueger, 1993) was strongly supportive of a remarkable wage premium, but as soon as more detailed data became available (see for instance DiNardo and Pischke, 1997) it was realized that heterogeneity was an issue and that it was necessary to disentangle the skill premium and the one related to the utilization of the computer. As in the case of the literature on the estimates of earnings return to computer use, more detailed data are necessary to further investigate to what extent both ability and utilization of new technologies can affect the labor supply of individuals.

Chapter 2

Health and Labor Supply Dynamics of Older Married Workers

2.1 Introduction

Population ageing is one of the greatest challenges faced by Europe. Currently, the proportion of EU citizens aged 65 or over is 16 percent and it is expected to approach 30 percent in 2050. The support ratio, which relates working age persons (20 to 64 years old) to those aged over 65, will halve in the next decades, declining from about 3.7 in 2000 to 1.9 in 2050. The combination between the decline of working age population and the low employment rates exhibited by the elderly questions the financial sustainability of public pension systems[1]. Aware of this, the Lisbon and Stockholm agreements plan to raise the participation to the labor market at all ages and, in particular, focus their attention upon the purpose of prolonging the working life of individuals around retirement. In this respect, a deep understanding of the factors determining the labor market outcomes of older workers is

[1] See Economic Policy Committee (2000) and OECD (1998).

needed to effectively design strategies aimed at enhancing their employment chances.

Our analysis considers married employed individuals aged 46-65 and draws data from the waves 1995-2001 of the European Community Household Panel (ECHP) [2] to assess empirically how their transitions towards not employment within the next year are affected by their own health conditions and by those of their spouses. In line with Blau (1994), we prefer to focus on labor supply dynamics rather than assume some predefined definition of retirement in order to (i) not exclude, for instance, episodes of re-entering the status of employed and (ii) consider information which would be otherwise dropped because of sample selection requirements[3]. In addition, as retirement process exhibits cross-country variability due to different institutional arrangements, imposing a unique rule of exit from the labor force may turn out to be unrealistic and produce misleading results.

The existing empirical literature shows a general consensus in considering health as an important determinant of older workers labor supply. In particular, several studies, such as Bound et al. (1999), Disney et al. (2006), Lindeboom and Kerkhofs (2002) and Zucchelli et al. (2007), show how healthier workers are more likely to remain employed. These findings confirm the theoretical predictions of a positive relation between psychophysical well-being and employment chances. As highlighted by Lumsdaine and Mitchell (1999), a bad shock on health may affect labor supply by altering both the budget constraint and the system of preferences. For instance, an ill employee is likely (i) to face worse compensation opportunities due to her reduced productivity and (ii) to value less the time spent in the labor market because of an enhanced disutility of work and the need of medical assistance.

[2] This chapter uses European Community Household Panel data from the December 2003 release according to the contract 14/99 with the Department of Economics and Management, University of Padova.

[3] Inclusion in the sample may depend on the years of contribution to the Social Security system or the frequency of unemployment spells during the working career.

Further, according to Grossman (1972), health conditions in a given time period result from a production function depending on the entire stream of health care investments previously made. In this framework, the empirical results quoted above suggest that adopting healthier life styles since early childhood may make individuals more likely to be healthier throughout their lives and to be more valuable for the labor market even at older ages.

Focusing uniquely on workers own health status translates in discarding the potential effects on the labor market position exerted by the psychophysical well-being of their relatives. In literature we find several contributions exploring the negative effects produced by caregiving demand of family members on employment patterns[4]. However, these investigations look mainly at the impact of looking after parents and analyse samples representative of the overall working age population.

Our analysis pursues a different goal. We restrict the attention to married older workers and examine how their probability of ceasing from work is influenced by the health status of their cohabiting partners. Finding a positive effect of spouse health conditions on the labor market attachment could uphold the view according to which workers in this age group not involved in looking after sick partners tend to prolong their working careers. As a result, the implementation of policies aimed at (i) exempting individuals from the provision of informal care and (ii) fostering the development of professional care market might turn out to be valuable in order to increase the low employment rates exhibited by the middle-aged and elderly population in Europe. Nevertheless, we do not focus only on the effect of caregiving but widen our interest to the more complex role played by spouse health in the determination of the labor supply of a married individual.

On the one hand, a healthy spouse may represent a source of help for managing nonmarket tasks, such as housework, and then result in an in-

[4] See Ettner (1996) and Heitmueller and Michaud (2006).

centive to keep on working. On the other hand, the propensity of a worker towards leaving the labor market may be enhanced by the possibility of spending her spare time with a healthy partner.

In addition, when a bad shock occurs on the health of a married individual, the labor earnings of the spouse could be considered as an additional source of income that can be exploited in order to afford the negative socio-economic effects engendered by the poor health episode, such as increased medical expenditures[5]. Alternatively, the worker may prefer to opt out of the labor market in order (i) to replace her spouse effort in home production activities and (ii) to provide the care eventually needed. Keeping on working may entail to consider the market as the provider of both (i) the goods previously home produced and (ii) the medical assistance for the partner. If the resulting costs are higher than the rewards of remaining employed, for instance in terms of labor income and pension wealth accrual, the individual might choose to leave her job.

Several works, such as Blau (1998), Blau et al. (1999), Jimènez et al. (1999) and Pozzebon and Mitchell (1989), document the presence of coordination between couple members labor supply. As pointed out by Michaud (2003), this relationship is likely to be established by spouses preference towards spending their time in the same occupational state. Additionally, in our set-up considering partner employment position is suitable in order to control for its close link with her own health. Not allowing for one of these factors could give rise to misleading estimates of the parameters of interest and imputing to a variable effects that are actually due to the other. As an example, the effect of spouse health on the labor supply of an individual described by a specification not allowing for spouse labor market outcomes could reflect the impact it is intended to capture along with the one pro-

[5] See Berkowitz and Qiu (2006), Rosen and Wu (2004) and Wu (2003) for a discussion concerning the impact of couple members health conditions on household wealth accumulation and saving decisions.

duced by the omitted variable, since those who are healthier experiment higher chances of being employed, *ceteris paribus*[6].

The chapter is organized as follows. In Section 2.2 we briefly review the issues about the measurement of health conditions in an econometric framework. Section 2.3 describes data and sample selection. Section 2.4 contains the results of descriptive nonparametric duration analysis. Section 2.5 is aimed at presenting the specifications adopted to produce the main results of this paper, which are reported in Section 2.6. Finally, Section 2.7 concludes.

2.2 Measuring health conditions

Health conditions are difficult to gauge correctly at an empirical level. Several datasets collect the opinion of individuals concerning their own health status, the so-called self-defined health status. Respondents are asked to rank their own physical and mental conditions according to a predefined scale, which usually spans from very good to very poor. This information is shown to be correlated with mortality indexes, as documented by Currie and Madrian (1999), and widely used in empirical works. Moreover, it presents the advantage of summarizing in a single variable all the information that contributes to determine the overall physical and mental conditions.

However, the opinion of an individual concerning her own health may be driven by factors unlikely to be available to the researcher, such as her concept of fair health status. As a result, people in the same health conditions may rank them differently because of unobserved heterogeneity causing the incomparability of self-assessments in the population and, consequently,

[6] Zucchelli et al. (2007) investigates how the labor supply of older workers is affected by the health and employment status of the other couple member. Remarkably, they consider a sample consisting of both married and unmarried individuals. This different sample selection prevents us from comparing our results with those produced by their analysis.

their unreliability. Nevertheless, the same conclusion applies even when the unobserved heterogeneity is assumed to be person-specific as well as time-varying. If it is the case, an individual experiencing the same health conditions in different time periods may self-rate them differently because its concept of fair health status has changed over time.

How to measure health properly is an open question. Bound et al. (1999) and Disney et al. (2006), exploit the information conveyed by more circumstantial health indicators, like indexes of functional activity or the diagnosis of severe health conditions, in order to purge self-assessments from the effects of individual unobserved heterogeneity. The validity of this approach crucially relies on the assumption that these detailed health indexes are measured according to a scale common to all population and not altered by individual characteristics.

Alternatively, other studies, like Berkowitz and Qiu (2006) and Wu (2003), gauge health condition just on the basis of the more objective health information mentioned above. However, although such indicators are less likely to be error-ridden, this advantage comes at the cost of potentially discarding other aspects relevant to define appropriately health status[7].

A further approach is to use anchoring vignettes. Its main drawback lies in the fact that its implementation crucially depends on the questionnaire design of the survey. In fact, this method consists of asking respondents to rate short hypothetical vignettes, or scenarios, describing different severity levels of health. Hence, vignettes could provide an estimate of the effects produced by unobserved person-specific heterogeneity in the determination of self-assessments. By relating individuals self-described health to the way they describe the vignettes, it is hypothetically possible to translate responses from different individuals to a scale that is comparable across

[7] An ulterior strategy is that of considering self-assessments jointly with more detailed indicators, as proposed in Jimènez et al. (1999). We think that summarizing health conditions in an unique variable is more appropriate to simplify the interpretation of the parameters and avoid redundancy problems.

different population groups and over time[8].

The unavailability of anchoring vignettes in the ECHP questionnaire and the necessity of exploiting all the relevant information to cast a reliable estimate of the health conditions of an individual lead us to tackle the likely measurement error of health self-assessments following the strategy proposed by Bound et al. (1999).

2.3 Data and sample selection

The ECHP survey[9] is designed by Eurostat and carried out yearly between 1994 and 2001 collecting information about European Union (EU) citizens aged 16 or over and living in Germany, Denmark, Netherlands, Belgium, Luxembourg, France, United Kingdom (UK), Ireland, Italy, Greece, Spain, Portugal, Austria, Finland and Sweden. ECHP is featured by a multipurpose questionnaire dealing with demographics, income, labor participation, education, training, health, social relations and migration.

The actual availability of variables relevant for the scopes of this work forces our sample to consist of the information conveyed by the 1995-2001 waves of ECHP about married employed individuals[10] aged 46-65 and resident in Denmark, Netherlands, Belgium, Ireland, Italy, Greece, Spain, Portugal, Austria and UK.

Table (B.1) contains the number of observations and individuals in the sample used for the following analyses. As it is displayed, the sample is stratified by gender and splits employees from self-employed. Table (B.2) shows how in all the groups of interest more than 85% of workers and their

[8] In the Survey of Health, Ageing and Retirement (SHARE) the description of anchoring vignettes is asked to a subsample of respondents for each country. By doing this, it is possible to purge self-assessments from unobserved country-specific heterogeneity.

[9] See Peracchi (2002) for an introduction.

[10] Note that the unit of interest is the married individual and not the couple. This amounts to say that if the spouse of a worker in the sample does not satify the eligibility criteria in terms of age and employment, her labor supply dynamics is not considered.

spouses report to be in at least fair health[11]. This is not striking, at least for workers, because, in general, their health should be good enough to allow them to carry out a job. In addition, it is shown how most spouses are themselves employed, conveying raw empirical evidence in favor of the hypothesis asserting that couple members prefer to spend their time in the same labor market position.

The focus of this analysis lies on the transition towards not employment. More precisely, we consider individuals employed at time $t = 1995, ..., 2000$ and look at their employment status in $t + 1$ by means of a dichotomous variable taking on value 1 if the individual moves to a not-employment state for whatever reason and 0 otherwise.

Table (B.3) points out that males at work experiment a lower probability of being not-employed in the next time period than females. Notably, this gap appears to be more sizeable for self-employed and widely goes up with age. Further, future labor market attachment is shown to be positively related to health measured at time t according to individual self-assessments. For instance, 8% of female employees in good health conditions at time $t = 1995, ..., 2000$ do not carry out a job in $t + 1$, but this percentage more than doubles when we look at their counterparts in bad health. Moreover, workers with a spouse in bad health conditions face a higher probability of exiting the state of employed[12]. While 13% of female employees with a partner in poor health quit their job, this percentage falls to 9% when those with a partner in good health are considered. This is in line with the view that workers cope with a bad shock on their spouses health by leaving their job and enhancing their role in home activities. Finally, it should be noted that the occupational patterns of couple members are shown to be correlated. In particular, for employees, the probability of becoming not employed drops

[11] Health self-assessments are defined according to a scale which makes up of 5 modalities, very good, good, fair, bad, very bad.

[12] We are conditioning on spouses health at time t. The same timing choice holds for their labor market position, which will be examined later.

by about one half when the cohabiting partner is not working.

We argue that examining the transitions out of employment in our sample may actually describe retirement from labor market because of the small probability of finding a new job exhibited by the population not at work in this age interval. More precisely, as documented in Table (B.4), among married individuals aged 46-65 only 5% (7%) of women (men) not employed at time t are at work at time $t+1$. It is worth noting how these re-employment rates exhibit cross-country variability. While in Denmark about 10% of individuals not at work find a job within the next year, in Italy this percentage diminishes to 3% and 6% respectively for women and men.

2.4 Nonparametric duration analysis

In this section labor supply dynamics is explored by means of a nonparametric duration analysis. The goal is to provide an estimate of the association between the probability of keeping on working in the future and the explanatory variables of interest. Our approach consists of (i) stratifying the sample according to such variables and (ii) comparing the Kaplan-Meier estimates of the survival functions obtained for the so-formed groups.

We first consider the effect on the employment patterns of a married worker exerted by changes in her own health. As before, health is measured according to the self-assessments provided by ECHP and ranked according to three distinct levels: good, fair and bad[13].

As documented in Figure (B.1), healthier female employees have higher chances of remaining at work in the future. For instance, while the probability of keeping on working amounts on average to about 50% for women aged 55 in at least fair health, it drops by more than one half for their counterparts in poor health. Remarkably, the health differentials are not

[13] An individual is assumed to be in good (bad) health if she declares to be in at least good (at most bad) health conditions.

negligible even for individuals aged 60 or over. Figure (B.2) points out that although, as expected, male employees present higher chances of remaining employed than women, the positive relationship between employment chances and health is magnified. While at age 55 those in either good or fair health experience at least 70% of chances of remaining at work next year, this probability falls dramatically to 25% for workers in bad health conditions. As before, the gap between individuals in bad health conditions and those healthier is still evident at older ages. Analogous considerations hold when self-employed are looked at.

Next, we conduct a similar analysis taking into account the association between the probability of remaining employed and partner health conditions. Figure (B.3) displays how female employees with a spouse in at least fair health steadily face higher probabilities of keeping on working. Figure (B.4) confirms this pattern for male employees. Notably, the gap between the three health-groups is sharper than in the case of women.

Turning to self-employed, the health conditions of the partner do not seem to play an active role in women labor supply dynamics. The opposite is found for men. In this case, the estimates of the survival functions are clearly separated and point to a negative relationship between the probability of ceasing from work and spouse health conditions.

To summarize, apart from the self-employed women case, the results of this descriptive analysis are consistent with the view according to which living with unhealthy spouses causes a decline in the time spent in the labor market because workers may find more profitable either providing the medical assistance or replacing partner effort in the home production of goods.

Finally, we look at the relationship between the probability of remaining employed of a married individual and the labor market position of her partner. Figures (B.5) and (B.6) points out that employees with a spouse at work are more likely to carry on working. This evidence confirms the hypothesis that couple members prefer to spend their time endowment in the same occupational state. On the contrary, when self-employed are looked at, this evidence holds for males but not for females. For this group the probability of keeping on working seems not to be affected by the labor participation of their spouses.

All these findings are confirmed by the log-rank test, which formally checks the equality between the Kaplan-Meier estimates of the survivor functions. The complete set of results is reported in Table (B.5).

The appeal of this nonparametric approach lies in the fact that it does not impose any functional assumption on the probability of remaining employed. This comes at the cost of allowing for explanatory variables only by means of sample stratifications. As a result, conditioning on a wider set of control factors may lead to small sample size problems and, consequently, to unreliable and imprecise estimates. The next section intends to set up an econometric framework aimed at overcoming these limitations.

2.5 Linear probability model

The transition towards not employment is defined by a dichotomous variable, y_{it} that takes on value 1 if an individual i employed at time t moves to a state of not employment within $t+1$ and 0 otherwise. We assume that y_{it} is determined by a linear probability model specification, such that

$$y_{it} = \beta_0 + \boldsymbol{\beta}_1' \mathbf{x}_{it} + \boldsymbol{\beta}_2 \mathbf{h}_{it} + \boldsymbol{\beta}_3 \mathbf{sh}_{it} + \beta_4 se_{it} + c_i + e_{it}. \tag{2.1}$$

Whereas the vector $\mathbf{x}_{it}$ contains control factors, $\mathbf{h}_{it}$ and $\mathbf{sh}_{it}$ indicate the set of variables characterizing the health conditions of, respectively, the individual i and her spouse at time t. The labor market position of the partner at time t is described by the dummy se_{it}, taking on value 1 if she is employed and 0 otherwise. Finally, the error term e_{it} is for the moment assumed to be uncorrelated with the explanatory variables[14].

Although this set-up discards the evaluation of the state-dependence[15] in labor supply determination, it explicitly models the potential transition out of employment of individuals currently at work. Hence, this specification strategy is particularly suited to capture the effects of the variables of interest on the individual decision of *keeping on* working in the future, which is the major concern of this work.

Our framework allows for the existence of time-invariant unobserved heterogeneity, denoted by the term c_i, relevant for both the labor supply of a married individual and all the factors in the right-hand-side of the model. Neglecting this issue may end up in obtaining only measures of the partial association between y_{it} and the explanatory variables. Following Chamberlain (1984), we assume that

$$c_i = \tau + \gamma' \overline{\mathbf{x}_i^c} + a_i. \tag{2.2}$$

where τ is an intercept, $\overline{\mathbf{x}_i^c}$ collects the averages over time of time-varying explanatory variables and a_i is a stochastic component. Plugging (2.2) in (2.1) leads to

$$y_{it} = \widetilde{\beta}_0 + \boldsymbol{\beta}_1' \mathbf{x}_{it} + \boldsymbol{\beta}_2 \mathbf{h}_{it} + \boldsymbol{\beta}_3 \mathbf{sh}_{it} + \beta_4 se_{it} + \gamma' \overline{\mathbf{x}_i^c} + u_{it}, \tag{2.3}$$

where $\widetilde{\beta}_0 = \beta_0 + \tau$ and $u_{it} = e_{it} + a_i$. Particularly important for our purposes, this specification strategy permits us to allow for unobserved time-invariant characteristics, such as past socio-economic conditions, affecting both the employment and the health outcomes of couple members.

[14] All the inference presented in the paper is robust to unknown (i) heteroskedasticity and (ii) autocorrelation at the individual level of the error term.

[15] See Heckman (1981) and Wooldridge (2002) for an introduction to the econometric techniques usually adopted to estimate the state-dependence in stochastic processes and Cappellari et al. (2007) for an application of such specifications to the study of labor supply of older workers in UK.

At first, health conditions are described by a set of dummies defined according to self-assessments. As in the previous section, we distinguish among three levels of psychophysical well-being: good, fair and bad, which is the baseline.

Next, we tackle the issue of the measurement error potentially affecting self-assessments by estimating an alternative health index as in Bound et al. (1999). In this case, the health variables in $\mathbf{h}_{it}$ and $\mathbf{sh}_{it}$ are the linear predictions of ordered probit specifications that regress couple members health self-ratings at a given time t on all the contemporaneous exogenous variables of the model[16] and a set of more objective indicators, $\mathbf{zh}_{it}$[17]. Information collected in the latter vector should refer to precise aspects of the physical and mental conditions of couples members, whose measurement cannot be altered by their beliefs and perceptions. In other words, such health indicators should be gauged according to a scale common to all population and not affected by person-specific heterogeneity. Exploiting the health section of ECHP questionnaire, the vector $\mathbf{zh}_{it}$ collects, for each spouse, three dummy variables, the first equals to one if the individual has spent at least one night in a hospital during the last twelve months, whereas the remaining two indicate the consultation with either a general practitioner or a medical specialist during the same reference period[18]. Once health indexes are calculated, their predicted values are plugged in $\mathbf{h}_{it}$ and $\mathbf{sh}_{it}$.

Although health indexes are obtained by means of first-stage regressions, OLS point estimates are still consistent but their variance and covariance matrix is not valid because it does not reflect the sample variability of the generated explanatory variables. Hence, intervallar estimates are computed by bootstrapping nonparametrically the overall estimation procedure[19].

[16]That is, $\overline{\mathbf{x}_i^c}$, $\mathbf{x}_{it}$ and $\mathbf{ze}_{it}$, which includes the additional instruments used to address the possible endogeneity of spouse labor market position.

[17]In order to allow for more flexibility in the estimation of health indexes, the ordered probit regressions are run separately for each wave.

[18]Totally, the vector $\mathbf{zh}_{it}$ consists of six variables.

[19]See Wooldridge (2002) for further details on the bootstrap technique.

Problems arising in rating health conditions lead us to discard the fixed effects (FE) approach in estimating (2.1) since this method crucially relies on the time-variation exhibited by regressors. If the changes over time characterizing the health status of an individual are at least partially spurious and attributable to reporting errors, FE may exacerbate the bias coming from the unobserved heterogeneity involved in the determination of self-assessments. Hence, we argue that in our context Chamberlain technique is safer to purge the estimates of the parameters of interest from the effect of c_i, even when corrections for measurement error are taken into account.

2.5.1 Endogeneity of spouse labor supply

As long as spouse employment status is considered exogenous, the parameters of interest in (2.3) are estimated by standard OLS technique. Removing this assumption translates into switching towards an instrumental variable (IV) approach.

The feasibility of the IV strategy relies on the availability of additional instruments $\mathbf{ze}_{it}$. In particular, we need an exogenous source of variation correlated with the employment position of the spouse se_{it}. A natural candidate is the yearly employment rate calculated for the population of individuals of the same gender and living in the same country as the spouse at time t. This indicator is reckoned by exploiting the waves 1995-2000 of ECHP as released by Eurostat. As long as they provide representative samples for the countries participating to the survey, valid estimates of the actual domestic employment rates can be calculated on the basis of this data source. Further, this choice permits to calculate the employment rates for age-intervals narrower than those usually considered by official statistics. We group the respondents aged at most 20 and assemble their older counterparts according to age-classes three-year long.

Each wave is stratified by country, gender and the age-classes. For each respondent in each stratum we calculate the proportion of the *other* group-members at work. Excluding individuals from the calculation of their associated employment rate is crucial to solve our endogeneity problem. In fact, instrumenting the spouse labor supply with an index depending on this variable itself would not tackle appropriately this issue because it might be endogenous as well. Instead, our approach allows us to obtain a measure that is (i) informative of the labor market conditions faced by individuals similar[20] to the spouse and (ii) constructed without exploiting the information on her own actual labor supply.

This *ad-hoc* employment rate is expected to be a valid instrument since it should be correlated with the labor market participation of the spouse but, conditional on the other observables, have a negligible influence on the employment outcomes of the other couple member.

More specifically, the consistency of the IV estimates based on $\mathbf{ze}_{it}$ requires (i) the information in $\mathbf{ze}_{it}$ to be correlated with the potential endogenous variable (*relevance of additional instruments*) and (ii) the whole set of instruments to be uncorrelated with the error term u_{it} (*validity of instruments*). Finally, (iii) if the right-hand-side variables in equation (2.3) are exogenous, OLS estimates should be preferred because of their higher efficiency (*exogeneity*).

While the hypotheses of validity of instruments and exogeneity are checked by carrying out respectively the usual Hansen[21] and Hausman[22] tests, the relevance of additional instruments is tested by regressing the potential en-

[20]Two individuals are defined as similar if they belong to the same group. As it is clear, there is a trade-off between the precision in measuring the similarity among individuals and the ensuing sample size of the groups.

[21]The overindentification is attained by including in $\mathbf{ze}_{it}$ an additional variable, along with the employment rate described above. For males and female employees $\mathbf{ze}_{it}$ is augmented with the interaction between the *ad-hoc* employment rate and the number of children aged 16 or less living in the household. For self-employed females $\mathbf{ze}_{it}$ takes in the squared *ad-hoc* employment rate.

[22]We consider the usual regression-based Hausman test. See Wooldridge (2002) for ulterior explanations.

dogenous variable on the whole set of instruments[23] and conducting a test of joint insignificance for the variables in $\mathbf{ze}_{it}$.

When generated health indexes are adopted, inference is carried out by bootstrapping the whole IV strategy. Notably, whereas the variance and covariance matrices considered to test the properties (i) and (iii) are estimated by bootstrapping the corresponding auxiliary regressions, the p-value of the Hansen test is calculated as the fraction of times the bootstrapped test statistics are higher than the one obtained from the original sample.

Finally, it is worth noting that when dealing with the endogeneity of binary explanatory variables, linear probability models turn out to be computationally advantageous if compared to other discrete choice specifications. While the latters usually require the implementation of cumbersome non-linear estimation methods, the consistency of the standard IV technique is unaffected by the nature of the endogenous covariate.

2.6 Results

Labor supply dynamics is described allowing for time dummies, country dummies, couple members age and education[24], household size, number of children aged less than 16, labor income, other household income, job characteristics (blue/white collar), sector of employment and number of years of contribution as of the time the worker enters the sample[25]. Following Chamberlain (1984), the time-invariant unobserved heterogeneity is controlled for

[23] In every time period t the set of instruments makes up of $\mathbf{ze}_{it}$, $\overline{\mathbf{x}_i^c}$, $\mathbf{x}_{it}$, and the vector of health variables $\mathbf{h}_{it}$ and $\mathbf{sh}_{it}$, regardless of their definition.

[24] Workers age is described by a set of four dummies, (i) aged 50 or less, (ii) aged 51-54, (iii) aged 55-59, (iv) aged 60 or over, which is the baseline. Instead, partners age is controlled by a second degree polynomial. Education is described by means of three dummies based on the ISCED code, (i) primary education (ISCED 0-2), secondary education (ISCED 3), high education (ISCED 5-7), which is the baseline.

[25] We include the interactions between time dummies and a dummy taking on value 1 if the respondent live in Southern Europe, namely Portugal, Spain, Italy and Greece. Also, we include the interaction between number of years of contribution and having at least secondary education (ISCED$\geq$3). Years of contribution are approximated using the age at the first job.

by enriching the specifications with the averages over time of worker age and age squared, household size, number of children aged less than 16, labor income and other household income[26].

We look at the results of two different specifications. The former takes the self-defined health status of couple members as reliable indicators of their psychophysical well-being. In this case, the parameters on health variables represent the average differential effect on the probability of becoming not employed with respect to the baseline, i.e. being in poor health. The latter addresses the issue concerning the potential measurement error affecting health self-ratings by adopting the alternative indicator suggested in Bound et al. (1999). This specification identifies the average changes in the outcome of interest produced by marginal variations in the health indexes. In both cases, the impact of couple members health is allowed to vary across workers age by interacting the health variables with a dummy taking on value 1 if the worker is aged 54 or less (*younger worker*) and 0 otherwise (*older worker*).

Although the OLS estimates are illustrated in Tables (B.8)-(B.11), for sake of brevity we comment only on IV results since the Hausman test reveals that the exogeneity assumption for spouse labor supply is generally rejected by data. The consistency of the IV strategy is based on (i) the relevance of the additional information collected in $\mathbf{ze}_{it}$ and (ii) the not-rejection by the Hansen test of the null hypothesis of validity for the overall set of instruments. Tables (B.6) and (B.7) summarize the IV point estimates of the causal effects of changes in partner labor supply and couple members health on the probability of transition towards not employment of a married worker. Tables (B.12)-(B.15) contain an extended set of results, including the tests of relevance and validity of instruments.

In general, the estimates of the parameters on the main control factors report the expected sign and the unobserved heterogeneity c_i seems to play

[26] In other words, the vector $\overline{\mathbf{x}_i^c}$ collects these averages.

an active role in the determination of transitions towards not employment. This latter evidence is produced by conducting a test of joint insignificance for all the parameters on the variables in $\overline{\mathbf{x}_i^c}$ assumed to describe the deterministic component of the unobserved dynamics.

Table (B.6) shows the results of the specifications describing health by means of self-assessments. The first column refers to female employees. As expected, healthier women present a significantly lower probability of leaving the initial state of employed. Irrespective of age, being in good health conditions is associated with an average increase of about 8 percentage points in the likelihood of remaining at work. Conversely, whereas the positive impact of experiencing fair health conditions amounts to 6 percentage points for younger workers, it attenuates and becomes statistically negligible when the older counterparts are considered. Comparing these IV point estimates with the corresponding raw differences highlights that netting out observable and unobservable characteristics usually reduces the effect object of study. Nevertheless, it remains sizeable and generally significant.

For younger employees, living with a partner in good (fair) health conditions induces an increase in the probability of ceasing from work of 5 (4) percentage points. Instead, for older workers, while the good health of the partner induces a significant rise of 4 percentage points in the likelihood of leaving the state of employed, the effect of fair health dissolves. It is noteworthy that the raw variation suggests an opposite and smaller effect, probably due to different sample compositions faded away at least partially by the set of control factors used in this regression analysis. Finally, our results point to the coordination between couple members employment status. When the spouse is at work, the future labor force attachment increases on average by 10 percentage points.

Next, we move to examine the results for self-employed females. Regardless of their age, healthier workers face a lower risk of leaving the state of employed. On the contrary, cohabiting with healthier spouses reduces the labor force attachment. For self-employed aged 54 or less, living with a partner in good (fair) health conditions makes stopping work in the future more likely by 10 (6) percentage points. This impact shrinks with age and for older individuals only having a partner in good health conditions is found to yield a significant drop in the likelihood of keeping on working. Finally, as in the employees case, an employed spouse induces an increase in the probability of carrying on work.

All these findings are overall upheld when male employees are looked at. Instead, a different pattern emerges for self-employed males. In this case, changes in their partners psychophysical well-being are not expected to bring about significant variations in their probability of ceasing from work.

The adoption of the alternative health index does not alter qualitatively the results. Table (B.7) documents that improvements in the psychophysical well-being of a worker are still estimated to have a positive impact on her likelihood of remaining employed, whereas the opposite pattern is found for the health of the cohabiting partner. Note that in the case of self-employed males the Hausman test fails to reject the null (Table B.15), and then spouse labor supply can be taken as exogenous. This amounts to say that OLS results in Table (B.11) should be considered in view of their higher efficiency.

2.7 Conclusions

The empirical analysis in this paper exploits the information conveyed by the waves 1995-2001 of the European Community Household Panel (ECHP) to examine the labor supply determinants of married workers aged 46-65. More specifically, we look at individuals employed at time $t = 1995, ..., 2000$ and focus on their likelihood of moving towards not employment within $t + 1$. Using linear probability model specifications, our goal is to assess the impact on the probability of keeping on working determined by variations in the health conditions of workers and in those of their cohabiting partners.

The overall physical and mental conditions of individuals are difficult to measure correctly. In this study they are first described by their self-assessments. This information is documented to be correlated with mortality indexes and collapses in an unique variable all the aspects relevant to gauge the psychophysical well-being of a person. In spite of this, its determination may depend on unobserved heterogeneity producing a measurement error and prejudicing the comparability of self-ratings in the population. We follow the approach proposed by Bound et al. (1999) in order to filter out self-ratings from the unobserved dynamics causing their unreliability.

The specifications adopted to describe labor supply dynamics allow for a wide set of control factors, in particular the employment position of partners. This choice is suggested by (i) the evidence provided by the wide research vein asserting the presence of coordination between the labor market outcomes of couple members and (ii) the close relationship existing between the health and the employment status of an individual. Hence, not allowing for partners labor supply could lead to (i) the exclusion of a relevant determinant of the optimal allocation of leisure for married workers and (ii) the parameter on partner health to capture even the effect that should be imputed to her own employment status, providing a misleading assessment of the causal effect of interest.

Health of married workers is estimated to affect significantly their employment conditions. In fact, in line with the existing literature, those who are healthier are always shown to report higher chances of remaining at work in the future.

In addition, apart from the case of self-employed males, we provide evidence of a positive relationship between spouse health and the probability of giving up working. Our findings are in line with Pozzebon and Mitchell

(1989), who find that in US married females with a sick spouse experiment a higher labor market attachment. Their results can be rationalized adducing the necessity to afford the costs of professional care and the incentives to work yielded by the fact that their job may furnish health insurance coverage to their partners. Although the need of purchasing medical assistance is likely to represent a major reason even in our context, the same does not hold for job-related health insurance, which is an institutional arrangement considerably more widespread in the US than in Europe. Our findings shows that European married workers with sick partners are shown to increase their likelihood of keeping on working even in the absence of this policy. The introduction of such plans may strengthen this link but at the cost of producing a negative effect on the propensity to work of the spouses themselves, as described by Buchmueller and Valletta (1999) and Chou and Staiger (1997).

A further explanation for the pattern proposed by our results could be that better physical and mental conditions of the partner induce a rise in leisure attractiveness, which diminishes the labor market attachment of the worker. However, discerning whether the impact of couple members health on their labor market outcomes is driven by an effect on the system of preferences or on the budget constraint requires the specification of structural models and it is beyond the scopes of this work.

Finally, when the partner is employed, the likelihood of ceasing from work is always estimated to drop. Our results support the hypothesis that, even conditioning on their health, couple members prefer to spend their time in the same employment state and that there exists coordination between their labor supply due to observable and unobservable determinants. This well-known pattern should play an important role in the design of policies aimed at enhancing the employment rates of older workers. For instance, pension systems allowing women to retire earlier than men are expected to produce a twofold effect on the labor supply of couple members. On the one hand, wives are expected to be more prone to leave the state of employed, *ceteris paribus*. On the other hand, their increased propensity to stop working induces a detrimental impact on their husbands labor supply driven by the coordination mechanism. Discarding this dynamics may translate in missing an important part of the process underlying the employment decisions of married workers.

Chapter 3

Estimating the Labor Supply Dynamics of Older Workers Using Repeated Cross-sections

3.1 Introduction

The Lisbon and Stockholm agreements plan to modernize the economy of the European Union along the lines of competition, knowledge, sustainable growth and social cohesion. Fulfilling these general purposes requires the achievement of a number of partial goals, such as the one of enhancing the labor market participation at all ages by 2010. More specifically, the employment rate is expected to approach 70% for the overall active population and 50% for the population group aged 55-64.

Although Italy has subscribed both the above-quoted treaties, these targets are far from being fully attained. As reported in Forze di Lavoro (2007) by the Italian National Statistical Institute (ISTAT), during the period 1995-

2006 employment rates of the active population exhibited a clear positive trend but still in 2006 only the labor market attachment presented by males was in line with the levels recommended by EU. Additionally, just 43.7% of men and 21.9% of women in the age interval 55-64 were at work. Unlike their younger counterparts, elderly individuals who do not carry out a job are likely to be eligible for retirement benefits. Hence, stimulating their employment rates may also translate in complementing the effects brought about by the pension reforms introduced in the last fifteen years, which encourage retirement postponement in order to (i) alleviate the financial burden borne by the National Institute for Social Security (INPS) and (ii) strengthen its long-run sustainability.

The design of effective policies pursuing such a task needs a thorough understanding of the decision process underlying the labor supply dynamics of older workers. This study looks at Italian employees aged 50-65 and performs a discrete-time duration analysis focusing on their hazard rate of stopping working within the next year. The choice of the age interval is dictated by the wide empirical evidence[1] asserting that the probability of retiring attains its peaks in this population group because of the institutional characteristics of the Italian pension system. In line with Blau (1994) we opt to focus on labor supply dynamics rather than assuming some arbitrary definition of retirement in order (i) to deal with a sample representative of all the employees in the age-range of interest and (ii) to consider any possible trajectory of their employment patterns. Therefore, our strategy takes into account exits from the state of employee occurring for whatever reason, such as following classical retirement routes or experiencing unemployment spells due to firm-downsizings that bridge older workers towards *ad-hoc* early retirement schemes.

Microeconometric investigations probing the transitions between differ-

[1] See Miniaci (1998) or Brugiavini and Peracchi (2003).

ent labor market positions are typically based on longitudinal datasets. They track respondents over time and fully characterize the dynamic structure of the process as well as the one of relevant factors supposed to influence its development. However, this kind of data source may present some disadvantages, such as severe attrition affecting the representativeness of the sample, limited cross-sectional dimension or restrictions on information availability[2]. For instance, the panel section of SHIW consists of less than one half of the overall interviewed households[3], whereas the Spanish Labor Force Survey does not provide access to family characteristics in its longitudinal version.

To overcome such limitations, which are common to many other countries, we decide to estimate the hazard rate of interest according to an application of the Generalized Method of Moments (GMM) technique proposed by Güell and Hu (2006). This framework turns out to be valuable in a duration context since it allows to calculate at the *individual level* the likelihood of *giving up* working only resorting to repeated cross-sections. In other words, despite employees are not tracked over time and their actual employment paths are not observed, this set-up makes us still able to reckon their probability of remaining at work in the future.

Data are drawn from the ISTAT survey Aspetti della Vita Quotidiana, which is released yearly for the period 1993-2003 and collect in each wave multipurpose information for a sample of approximately 20,000 households representative of the whole Italian population[4].

When analyzing the hazard rate of exiting employment for Italian older workers, we should not overlook the role played by the several reforms undergone by the Social Security system since the early nineties. As written

[2] See Heckman and Robb (1985) for further details.

[3] In the 1991 wave of SHIW less then 30% of households have been previously interviewed. This proportion rises to about 40% for all the following waves.

[4] Waves 1993-1999 and 2001-2003 of the data used in this chapter are released according to the contract 4842 with the Department of Economics and Management, University of Padova. I thank Fausta Ongaro (Department of Statistics, University of Padova) for providing me with the wave 2000 of this survey.

above, retirement from the labor force is a popular state among the elderly who are not working. According to Aspetti della Vita Quotidiana 1993-2003, while 85% of not employed men aged 50-65 self-define as retired, this proportion falls to 34% in the case of females. However, although 60% of women not at work in this age-range are housewives, almost one-fifth of them rate retirement benefits as their main source of income[5]. Hence, the incentives to work provided by the Social Security system are likely to be an important determinant of the labor supply dynamics of the workers in our sample.

In this respect, the adoption of the survey Aspetti della Vita Quotidiana is particularly advisable in view of the long time-span covered. This feature permits to show the evolution of the labor market attachment of the elderly and to check whether it is consistent with the ameliorative effects pursued by policy-makers. Similar research questions could be additionally addressed using the ISTAT Quarterly Labor Force Survey (RTFL) for the years 1992-2004, which describes in detail both current employment status and past working history. However, Aspetti della Vita Quotidiana provides a richer description of the overall socioeconomic condition of workers, for instance in terms of health and behavioral risks. Since these factors are expected to play a major role in determining the probability of keeping on working, we prefer to use the latter survey. Finally, in light of these considerations and of the large sample size, we think that Aspetti della Vita Quotidiana compares favorably to the longitudinal datasets currently available for studying the employment patterns of Italian older workers, such as the Bank of Italy Survey of Household Income and Wealth (SHIW), the European Community Household Panel (ECHP) and the public releases of the INPS archive.

The structure of the chapter is the following. Section 3.2 briefly describes the principal changes in the Social Security system ongoing in the period

[5] It should be remembered that these statistics refer to the whole population of individuals not at work irrespective to their working history.

considered in this analysis. A brief survey of the related literature is provided in Section 3.3. Section 3.4 introduces the GMM estimator adopted. Section 3.5 presents the dataset used to obtain the main results of this work, which are the focus of Section 3.6. Finally, Section 3.7 concludes.

3.2 Social security reforms in Italy

The main reforms of the Social Security System passed during the nineties take their names from the Prime Ministers at the time. Specifically, they are the Amato (1992), Dini (1995) and Prodi (1997) acts. Further, minor modifications have been promulgated almost yearly since 1992. All these changes tighten the eligibility criteria and introduce less generous rules for the pension benefit computation in order to stem the runaway growth of the outlays managed by INPS. Such modifications were needed in view of the aging process characterizing the Italian society, the improved life-expectations of the elderly and the ensuing necessity of adjusting the amount of pension benefits granted to the worker during her retirement years to the actual amount of contributions paid to the Social Security throughout her working life.

The Dini reform is the most incisive because it switches the original defined benefit system to the defined contribution method. This conversion is phased in over a long transitional period and will interest only the most recent cohorts of individuals. In fact, it distinguishes between the workers with at least 15 years of contribution in 1992 and all other workers. In particular, the former group is totally exempted by the transformation of the scheme and it is exposed only to less radical adjustments in eligibility rules and benefit formulas.

This section considers only the legislated changes modifying the old-age and seniority pensions available for employees. In spite of this, it should be remembered that the Social Security system also covers self-employed workers and provides social, survivor and disability benefits[6].

Until 1992 males (females) could claim old-age benefits not before age 60 (55) and conditional on having contributed to the scheme for at least 15 years. The Amato reform gradually increases these requirements in order to grant by 2002 old-age pensions only to males (females) aged at least 65 (60) years old with at least 20 years of contribution.

Additionally, till 1992 eligibility for seniority pensions differs among sectors of employment. Whereas employees of the private sector had to collect at least 35 years of contribution, this requirement was much looser for public sector workers (20 years of contribution for males and 15 for females). Since 1996 these rules have been tightened and harmonized across sectors of employment. In 2003 workers of the private (public) sector with 35 years of contribution could retire only if they were at least 57 (56) years old. Alternatively, retirement was allowed at any age if workers had collected 37 years of contribution. Further, during the years 1995-1999 the eligibility for seniority pensions was additionally limited by the so-called exit windows, which forced workers to defer retirement by a period between six and twelve months.

Prior to 1995, in line with the defined benefit schemes, pension amounts depended on pensionable earnings and a proportionality factor increasing with the length of contribution history. Although the Dini reform enforces the defined contribution scheme for younger cohorts of workers, the old rules still apply to compute a part of their benefits. In general, pensionable earnings result from a weighted sum of past earnings. The reforms of the system basically extended the period taken into account for this computation. Until 1992 only the last 5 years of work were considered. In 2001 this time-interval included (i) the last 10 years for workers covered by the defined benefit scheme and (ii) all the working history for those under the defined

[6] See Battistin et al. (2007) and Brugiavini (1999) for further details.

contribution system.

In conclusion, it should be noted that the Amato reform of 1992 started to provide incentives, through actuarial rewards in the benefit computation, to those workers who decide to postpone retirement even if they are eligible for either old-age or seniority pensions and have contributed to the scheme for more than 40 years.

3.3 Literature review

The microeconometric analysis of the labor supply dynamics of Italian older workers has received the attention of many studies differing with respect to the theoretical framework, the econometrics and the sample adopted. We recall some contributions to briefly outline the state of the art in this research field[7].

Miniaci (1998) studies individual decisions of retirement by means of a fully reduced-form duration analysis. He exploits the retrospective information conveyed by the wave 1995 of SHIW in order to characterize retirement patterns distinguishing between alternative exit routes from the labor market. The sample includes males and females who are, respectively, aged 50-70 or 45-65[8] and household heads or their spouses. The main results come from the estimation of multinomial logit and Cox models. They point out that workers of younger cohorts retire earlier and, consequently, remain economically inactive for a higher proportion of their lives in view of better life expectancies. As expected, better educational attainments and lower replacement ratios enhance the likelihood of being at work. *Ceteris paribus*, public sector employees are not shown to retire earlier and no significant differences arise between the North and the South in terms of the probability of applying for invalidity or social benefits.

[7] For further details, see the references quoted in these works.

[8] More precisely, the inclusion in the sample requires males (females) to become retirees after the age of 50 (45).

Brugiavini and Peracchi (2003) analyse the determinants of the transitions towards retirement using a sample extracted from the INPS archive. Their theoretical framework suggests that at any age individuals decide to keep on working or retire by (i) comparing the expected present values associated with these two outcomes and (ii) opting for the employment state producing the highest pay-off. Taking advantage of the detailed income information provided by the data source chosen, this work spends great attention in defining future earnings, pensionable earnings and social security wealth, which are expected to play a prominent role in this decision. Their sample includes only private sector non-agricultural employees aged 50-69 who have started an employment spell between 1977 and 1996. Reduced-form probit models are utilized to evaluate the relationship between the transitions towards retirement and a set of explanatory factors including, among the others, a full set of age dummies and the above-mentioned income variables. According to their estimates, the actual age-profile of retirement rates is well-replicated and the parameters on the income variables have the expected sign. In particular, the probability of ceasing from work decreases with future labor earnings and rises with higher levels of pension wealth and pensionable earnings. The estimated models are then used to simulate how the retirement rates change under different institutional scenarios that alter both eligibility requirements and pension wealth computation[9].

Spataro (2000) describes retirement decisions according to an option-value model as in Stock and Wise (1990) and imposes specific assumptions on the utility function of individuals. The parameters characterizing agents preferences are estimated exploiting the waves 1991 and 1993 of SHIW and focusing upon individuals aged 45-65 and at work at the end of 1990. Unlike

[9]Brugiavini and Peracchi (2001) consider similar research questions and use an older release of the same dataset. Finally, Brugiavini (1999) estimates the impact of social security parameters on actual retirement patterns and on expectations about the retirement age.

Miniaci (1998), only old-age and seniority retirement routes are considered[10]. Modelling the risk of giving up working by means of probit models he obtain the estimates of the structural parameters along with the evolution over age of the hazard rate of interest. The fitted hazard rates are close to the actual ones but, as in comparable studies with American data, they do not fully capture the peak in the probability of retiring occurring at age 60. In general, Italian workers are more prone to retire earlier and are characterized by a higher risk aversion as well as a lower intertemporal discount rate than their US homologues. Finally, the comparison between the actual expected retirement ages and those coming out from the fitted model[11] confirms the predictive power of this framework only for the group of workers aged 51 or over. This limitation may be rationalized by the fact that the retirement option is actually taken into account only by this population group. Hence, the inclusion in the sample of younger workers is likely to introduce bias in the estimates.

Colombino (2003) departs from the focus on the transitions out of employment but still specify a structural model of retirement. He follows Gustman and Steinmeier (1986) and assumes that individuals choose their optimal labor market position on the basis of the comparison between the instantaneous utility levels of retiring and being at work. It can be shown that in such framework the estimation of the structural parameters can be carried out by following a standard logit analysis of the current employment state. Data are drawn from the wave 1993 of SHIW and the sample takes in only the household heads and their spouses aged at least 40 years old and either at work or job pensioners. Likewise Brugiavini and Peracchi (2003), the fitted model permits to simulate the response of individual behaviors to legislated changes to the pension system. It is found that a marginal cut in

[10]This choice also reflects the spirit of the option value model which specifies different utility functions for those who are at work and those who retire.

[11]In SHIW all individuals at work are asked to report the expected age of retirement. The comparison implemented exploits the waves 1991-1995 of the survey.

the benefits ends up in a small but not irrelevant reduction of the number of retirees. On the contrary, dropping the eligibility requirements produces only a slight increase in the proportion of pensioners within the population. This latter evidence suggests that at least in 1993 eligibility constraints were not binding.

The results obtained in our analysis does not go through the development of a structural model of retirement. It is no doubt that recovering structural parameters describing individual preferences is relevant for policy purposes but this comes at the cost of imposing unverifiable assumptions on the utility functions of individuals. Instead, we follow Miniaci (1998) and Brugiavini and Peracchi (2001 and 2003) in order to specify a reduced form approach to estimate the labor supply dynamics of older workers. Finally, it is worth remembering that our study complements the contributions listed in this section since it builds upon a dataset that allows to describe the evolution of the labor market attachment of the elderly in a period featured by a long series of changes to the pension system.

3.4 The econometric framework

Let y_i be a binary random variable taking on value 1 if an employee in the age interval 50-65 at time t keeps on working in $t+1$ and 0 if she moves towards not employment for whatever reason. We assume that

$$y_i = 1\left\{x_i'\beta + e_i > 0\right\}, \tag{3.1}$$

where x_i is a vector collecting explanatory factors of interest, β is the set of parameters we intend to estimate and e_i is a stochastic component following the logistic distribution[12]. As a result, the probability of remaining at work is defined as

[12] Strictly speaking, in a duration analysis $y_i = 0$ identifies the exit from the initial state of employee.

$$\Pr(y_i = 1) = \Lambda(x_i'\beta) = \frac{\exp(x_i'\beta)}{1 + \exp(x_i'\beta)}. \tag{3.2}$$

When the coefficients in β are estimated via the maximization of the standard log-likelihood function

$$y_i \sum_i \log \Lambda(x_i'\beta) + (1 - y_i) \sum_i \log(1 - \Lambda(x_i'\beta)),$$

we obtain the first-order conditions

$$\underbrace{\sum_i x_i \Lambda(x_i'\beta)}_{\substack{\text{all employees} \\ \text{aged 50-65 at time } t}} = \underbrace{\sum_i y_i x_i}_{\substack{\text{only those still in the sample and} \\ \text{employed at time } t+1}}. \tag{3.3}$$

On the one hand, the LHS of (3.3) is the sum of x_i over all the employees in the age range of interest at time t weighted by their conditional probability of remaining at work given x_i. On the other hand, the RHS considers only those still in the sample and employed at time $t + 1$. The consistency of the maximum likelihood estimates requires the eventual attrition in the sample to be uncorrelated with the outcome of interest once we condition on x_i.

The implementation of this standard method crucially relies on the availability of panel data that track individuals over time and allow to observe the variable y_i. Güell and Hu (2006) propose an alternative approach that calculates at the individual level the probability of leaving a given initial state using independent cross-sections representative of the same population at different time periods.

In on our specific context the identification of β requires the definition of (i) an *entry* cross-section collecting employees aged 50-65 at a given time t and (ii) an *exit* cross-section collecting their counterparts aged 51-66 at

$t+1$. The fundamental underpinning for all our results is that individuals aged 51-66 in $t+1$ should behave as those aged 50-65 in t tracked down one year later. Notably, in this set-up the assumption of dealing with two cross-sections representative of the Italian employees entails that the inclusion in the sample should not depend on unobserved characteristics affecting the labor market position.

Under these restrictions, we mimic equation (3.3) by the set of moment conditions

$$\underbrace{\sum_i x_{it}\Lambda(x'_{it}\beta)}_{\substack{\text{all employees aged 50-65} \\ \text{in the cross-section of time } t}} = \underbrace{\sum_j x_{jt+1}}_{\substack{\text{all employees aged 51-66} \\ \text{in the cross-section of time } t+1}} . \tag{3.4}$$

The LHS of (3.4) is the sum of the vectors of explanatory variables over the sample of employees aged 50-65 at time t weighted by their conditional probability of keeping on working in $t+1$. Instead, the RHS is its un-weighted homologue calculated for the entire set of employees aged 51-66 at time $t+1$. It is worth stressing that since y_i does not show up in (3.4), the feasibility of this expression does not rely on the necessity of following individuals over time[13]. This amounts to say that, unlike the FOCs in (3.3), this way of proceeding circumvents the presence of attrition damaging the representativeness of longitudinal datasets.

Building crucially upon the assumption that the cross-sections of times t and $t+1$ are randomly drawn from the same underlying population, the adoption of the moment conditions (3.4) can be motivated going through the law of iterated expectation,

[13] The indexes i and j in the notation x_{it} and x_{jt+1} intend to emphasize that the cross-sections do not make up of the same set of individuals.

$$E[x_{jt+1}\mathbf{1}(t+1)] = E[x_{it}E[\mathbf{1}(t+1)|x_{it}]] = E[x_{it}\Pr(t+1|x_{it})],$$

where $\mathbf{1}(t+1)$ takes on value 1 if the individual is at work in $t+1$ and 0 otherwise. The analogy principle shows easily that $E[x_{jt+1}\mathbf{1}(t+1)]$ is the population analogue of the RHS of (3.4) after a sample size normalization and the same applies to $E[x_{it}\Pr(t+1|x_{it})]$ and the corresponding LHS.

As pointed out earlier, we draw data from the waves 1993-2003 of the ISTAT survey Aspetti della Vita Quotidiana. In order to exploit massively the available information, the entry cross-section should pool together all the individuals for whom we observe their counterparts one year later, who will be in turn collected in the exit cross-section. Hence, the entry dataset consists of the sample of Italian employees aged 50-65 in 1993-2002, whereas the exit cross-section of their homologues aged 51-66 in 1994-2003. As it is evident, the entry and exit cross sections are by construction not independent as employees aged 51-65 in the years 1994-2002 figure in both of the datasets.

The GMM estimator $\widehat{\beta}_{GMM}$ results from the minimization of a weighted quadratic function of the vector

$$g(\beta) = \sum_j x_{jt+1} - \sum_i x_{it}\Lambda(x'_{it}\beta). \tag{3.5}$$

More specifically,

$$\widehat{\beta}_{GMM} = \arg\min_{\beta} g(\beta)W^{-1}g(\beta), \tag{3.6}$$

where W^{-1} is a symmetric, positive semidefinite weighting matrix.

Given the moment conditions stacked in $g(\beta)$, the efficient GMM estimator is obtained if the weighting matrix W^{-1} is the inverse of the variance and covariance matrix of $g(\beta)$. Naming n_t and n_{t+1} the sample size of, respectively, the entry and exit cross-sections, we define

$$
\begin{aligned}
W = V\left[\frac{1}{\sqrt{n_t}}\sum_j x_{jt+1} - \frac{1}{\sqrt{n_t}}\sum_i x_{it}\Lambda(x'_{it}\beta)\right] = \\
= V\left[\frac{1}{\sqrt{n_t}}\sum_j x_{jt+1}\right] + V\left[\frac{1}{\sqrt{n_t}}\sum_i x_{it}\Lambda(x'_{it}\beta)\right] - \\
-2Cov\left[\frac{1}{\sqrt{n_t}}\sum_j x_{jt+1}, \frac{1}{\sqrt{n_t}}\sum_i x_{it}\Lambda(x'_{it}\beta)\right] = \\
= \frac{n_{t+1}}{n_t}V\left[x_{jt+1}\right] + V\left[x_{it}\Lambda(x'_{it}\beta)\right] - \frac{n^c_{t,t+1}}{n_t}2Cov[x_{jt+1}, x_{it}\Lambda(x'_{it}\beta)]. \quad (3.7)
\end{aligned}
$$

It is evident how the weighting matrix W^{-1} reflects the sample design of our analysis generating not independent entry and exit cross-sections. In fact, the information conveyed by the set of $n^c_{t,t+1}$ individuals showing up in both the datasets drives the covariance term in (3.7)[14].

As usual, the implementation of the optimal GMM estimator needs a set of starting values $\widetilde{\beta}$ coming from a consistent estimator of β. We decide to obtain them by running the GMM machinery and assuming W to be the identity matrix. Our choice is suggested by the fact that GMM estimators are always consistent regardless of the weighting matrix adopted.

Finally, it can be shown that $\widehat{\beta}_{GMM}$ is asymptotically normally distributed,

$$\sqrt{n_t}(\widehat{\beta}_{GMM} - \beta) \approx N(0, (A'W^{-1}A)^{-1}).$$

[14] In Güell and Hu (2006) the moments conditions specified are uncorrelated and produce a diagonal weighting matrix.

where[15]

$$A = E[\frac{\partial}{\partial\beta}\frac{1}{n_t}g(\beta)] = -\frac{1}{n_t}E[\frac{\partial}{\partial\beta}\sum_i x_{it}\Lambda(x'_{it}\beta)] =$$
$$= -\frac{1}{n_t}E[\sum_i \Lambda(x'_{it}\beta)(1-\Lambda(x'_{it}\beta))x_{it}x'_{it}] =$$
$$= -E[\Lambda(x'_{it}\beta)(1-\Lambda(x'_{it}\beta))x_{it}x'_{it}].$$

The variance and covariance matrix of $\widehat{\beta}_{GMM}$ is estimated replacing A and W with their sample analogues.

Our study develops a duration analysis in which time is indicated by the age at which employees are found at work. As written above, the probability $\Lambda(x'_{it}\beta)$ indicates the likelihood of remaining employed in the future given the characteristics included in the vector x_{it}. To preserve the comparability with the findings provided by the related literature, in the remainder of the chapter all the results will be presented with respect to $1 - \Lambda(x'_{it}\beta)$, which identifies the hazard rate of exiting employment within the next year.

Unlike proportional hazard models, logit specifications used in this analysis allow the effect on the hazard rate of a given explanatory variable to change with the values taken on by all the other covariates. Therefore, as long as age figures among the regressors, the impact of the other explanatory variables is assumed to be duration dependent.

Finally, the GMM estimator proposed in this work can be additionally implemented using the waves of surveys consisting of both a longitudinal and a refresher component, like SHIW and the Survey of Health, Ageing and Retirement in Europe (SHARE). The advantage of our approach over classical panel data models lies in that it exploits the information coming from the whole dataset and not from its longitudinal section only. If each wave is cross-sectionally representative, this framework circumvents the potential

[15] This result comes from Newey and McFadden (1994).

attrition in the sample and avoids complicating the estimation method to explicitly account for this issue.

3.5 Data

The repeated cross-sections called Aspetti della Vita Quotidiana has been collected yearly by ISTAT from 1993 to 2003. This multipurpose survey gathers unique information about a variety of aspects concerning the life-style of Italian households. Each wave furnishes a sample of approximately 60,000 individuals who are asked about education, training, accommodation, market and nonmarket activities, leisure, health and utilization of medical care services[16].

3.5.1 Descriptive cohort analysis

We develop a descriptive cohort analysis because it is of help to present both our dataset and the information utilized by the GMM framework discussed above. More specifically, we intend to describe the age-profile of the employment rates exhibited by individuals born in 1929, 1936, 1943, 1950 during the period 1993-2003. In line with the general focus of this study on older workers, the age-interval of interest is restricted to 50-70[17].

This analysis shows the evolution over time of the probability of being at work disentangling the different roles played by age and year of birth in its determination. Figure (C.1) reports the results for males and suggests that between 50 and 53 there is a positive time trend according to which younger cohorts exhibit a higher propensity towards being at work. For instance, at

[16] However, since Aspetti della Vita Quotidiana is not explicitly designed for labor supply analyses it misses out information on characteristics, like income and built-up pension wealth, expected to affect the employment status. A discrete classification of the household income is available only for waves 1996 and 1997.

[17] Since the computation of an in-depth cohort analysis is not among the principal purposes of this work, we select only four birth-cohorts in order to preserve the clarity in the figures reporting the results of this section. Totally, the selected sample consists of 11,017 (11,407) observations for males (females).

age 53 those born in 1950 are on average about 10 percentage points more likely to be employed than those born in 1943. However, the reverse pattern is found when older ages are taken into account. In fact, at age 60, males born in 1936 experiment a higher probability of carrying out a job than their homologues in cohort 1943.

Figure (C.2) shows that for women there is a clearer evidence pointing to an increase in the labor participation over time. Nevertheless, this pattern is still not age-invariant. While at age 53 females born in 1950 are more than 20 percentage points more likely to be at work than those born in 1943, this difference becomes negligible when we compare the employment rates at age 57 of the cohorts of women born in 1943 and 1936.

The lack of a time trend common to all ages is maybe driven by the several pension reforms of the last fifteen years. Whereas workers born in 1929 and 1936, face a more stable institutional set-up, the labor market outcomes of the younger cohorts were deeply influenced by the long series of changes to the pension system passed by the Parliament almost every year. In view of such uncertainty, the workers of more recent cohorts, e.g. 1943, probably decided to retire from the labor force as soon as possible in order to avoid expected unfavorable modifications in the Social Security system.

This cohort analysis has not been carried out by following over time the same set of individuals, as instead occurs in longitudinal datasets. In fact, we are dealing with pooled cross-sections collecting in different time periods representative samples of individuals born in different calendar years. As an example, in Figure (C.1) the age-profile of the employment rates experimented by males born in 1950 at ages 52 and 53 comes from considering males aged 52 in 2002 and their counterparts aged 53 in 2003. Although these sub-samples are not composed of the same individuals, as long as they furnish valid estimates for the employment outcomes in 2002 and 2003 of males born in 1950, we are able to reckon the evolution of the chances of being at work for this birth-cohort.

Analogously, on the basis of cross-sections representative of the same population of employees at times t and $t+1$, we can recover at the individual level their hazard rate of stopping working within the next year. The GMM approach proposed in this analysis calculates the likelihood of becoming not employed for *each* individual at work at time t without need of longitudinal datasets but only resorting to the information conveyed by her counterparts of the same cohort included in the cross-section of time $t + 1$.

3.6 Results

3.6.1 Age and cohort profiles of the hazard rate

We consider employees aged 50-65 in 1993-2002 and calculate the age-profile of their hazard rate of becoming not employed within the next year. In particular, we allow the hazard rate to vary over all pairs of years from 1993-1994 to 2001-2002 and according to a full set of age dummies. In general, we define one dummy for each year of age. Lack of variability in the sample of women forces us to impose linear restrictions in the specification. We aggregate in single age-classes female employees aged (i) 50-51, (ii) 52-53 and (iii) 62-63.

The interpretation of the parameters on year dummies should be cautious. On the one hand, these variables may capture the impact of the reforms of the Social Security system but also reflect whatever macro-change occurred in the Italian economy during the period of reference. On the other hand, they are also allowed to pick up cohort fixed-effects, namely the variations on the hazard rate due to the variability in the socioeconomic conditions characterizing individuals at the same stage of their life-cycle but born in different calendar years. This second limitation can be neglected only if we assume that the cohort-specific heterogeneity is fully controlled by the explanatory factors included in the model.

Table (C.1) presents the age-composition and the sample size of the entry and exit cross-sections used to implement the estimation strategy. The same samples will be used for all the following analyses. Regardless of the gender, while more than 50% of the individuals in the entry cross-sections are aged 50-54, the proportion of those aged 60 or over amounts to less than 15%. This raw evidence highlights the low labor market participation of the elderly[18].

Table (C.2) suggests that the time trend of the hazard rate is hump-shaped and that, conditional on age, male (female) employees in 1997-2000 (1995-1998) present the lowest likelihood of keeping on working in the future. Although the hazard rate significantly reduces for the youngest cohorts of employees (those at work in 2001-2002), some cross-gender differences emerge. While males of these cohorts exhibit a probability of ceasing from work significantly lower than the baseline group (i.e. those at work in 1993-1994), no significant deviations with respect to the benchmark are found for their females counterparts. Figures (C.3) and (C.4) illustrate the age-profiles of the hazard rate for male and female employees in 1993-1994, 1997-1998 and 2001-2002 calculated on the basis of the estimates of Table (C.2).

The related literature agrees on pointing out peaks in the probability of quitting employment around ages 55, 60 and 65[19]. Although the evidence summarized in Figures (C.3) and (C.4) is overall in line with the expected pattern, some notable differences are found. As an example, higher risks of leaving employment are found at ages 57 and 61 for males and 56 for females. These deviations from the classical timing of retirement may be rationalized by, again, the presence of ongoing reforms of the pension system[20]. The gradual modifications in the requirements for old-age and seniority pensions

[18] For the exit cross-sections similar considerations hold.

[19] See Miniaci (1998) and Brugiavini and Peracchi (2001 and 2003).

[20] Exits at the age of 57 are in line with the age-requirements for the seniority pensions described in Section 3.2.

alter the set of incentives to remain employed provided to workers and, consequently, bring about changes in the shapes of their hazard rate of exiting employment as compared to those found in studies considering datasets referring to antecedent periods.

3.6.2 Allowing for further explanatory variables

Our specifications are enriched by a more exhaustive set of regressors including year dummies, age, household size, hospitalization during the last twelve months, subscription of either health or life insurances, smoking habits, homeownership, education, job characteristics, sector of employment and region of residence. The results are contained in Table (C.3).

Parameters on year-variables still assert that, even conditioning on ulterior factors, the evolution over time of the hazard rate is hump-shaped. Maintaining cohorts of employees in 1993-1994 as baseline, the risk of becoming not employed within the next year significantly rises for their counterparts at work in 1995-2000 and goes down for youngest cohorts.

Our estimates suggest the puzzling evidence that while two important pension reforms like Dini (1995) and Prodi (1997) have been passed in order to induce workers to postpone retirement, employees experiment the highest propensity towards leaving employment. However, it should be remembered that the full implementation of the legislated changes promulgated during the nineties requires a considerable transitional period and the ameliorative effects, if any, will be sizeable only in the middle-run.

The youngest cohorts in the sample may exhibit a fall in the hazard rate because the Social Security actually begins to provide incentives designed to extend their permanence in the labor market. On the contrary, as argued before, the uninterrupted series of reforms lowering the overall generosity of the pension system may have led employees at work in the years 1995-2000 to retire as they became eligible in order to avoid further coercive extensions of their working life and poorer pension benefits.

In Table (C.4) we compute the variations in the hazard rate exerted by changes in time of employment with respect to a reference category consisting of white-collar employees of the primary or industry sector in 1993-1994, homeowners, with an upper secondary school degree, not having been hospitalized during the last twelve months, not having subscribed either a health or a life insurance, non-smokers, living in the North in a household with 3 components.

Males aged 60-65 in this reference category have a hazard rate of becoming not employed equal to 31 percentage points (ppt) and are on average 35% less likely to stop working than their homologues at work in 1997-1998. Conversely, their hazard rate is 46% higher if compared to the one for employees in 2001-2002. This pattern is even more marked for females of the same age. Their baseline risk amounts to 46 ppt, goes up by 79 percent for cohorts of employees in 1997-1998 and finally reduces by 33.70 percent when those at work in 2001-2002 are looked at. Table (C.4) also contains the results of the same exercise carried out for individuals aged 56-57. We notice that, in particular for women, while the benchmark probabilities diminish, the fluctuations around the baseline are wider than in the previous case. In particular, although the hazard rate for females aged 56-57 in 1993-1994 is remarkably low (7 ppt), it is overall in line with the evidence proposed in Figure (C.4), which associates a higher labor market attachment to females in this age range and detects a peak in the risk of ceasing from work within the next year for their slightly younger counterparts. A rationale for the negligible hazard rate of this population group derives from the design of the pension system, which grants seniority benefits only to workers with a contribution history much longer than that required for old-age pensions. In fact, our results suggest that women *still* at work at age 56-57 in 1993-1994 are not going to apply for the old-age exit route[21] but plan to keep on working in order to accumulate the necessary years of contribution set by the eligibility rules for seniority pensions. On the contrary, female employees of the same age in 1997-1998 are less prone to increase their pension wealth and prefer to anticipate their exit from the labor market.

Table (C.3) reports that, irrespective of the gender the risk of giving up working exhibits a clear positive age-profile witnessing the small incentives to remain employed provided to the elderly by the Italian labor market institutions. This expected evidence might represent not only a genuine age effect indicating that the disutility of work rises with age but it can also be driven by the higher levels of pension wealth built-up by older workers, which makes the retirement option more favorable.

Our findings also distinctly point to a negative relationship between household size and the probability of keeping on working. Living in larger families may entail a heavier burden of non-market activities, like looking after grand-children or minding elderly parents, which are associated with a higher propensity towards exiting the labor market, *ceteris paribus*.

Hospitalization is a proxy for the overall health status. Consistently with what suggested by the literature[22], those who have been admitted to a hospital during the last twelve months present higher hazard rates of ceasing from work. We also allow for smoking habits and subscriptions of life or health insurances because they are indicators of the likelihood of future shocks on health conditions. In particular, it can be argued that purchasing an insurance is an index of both the risk aversion of the agent and of her self-rated risk of incurring in bad health episodes. While none of these variables is significantly related with the hazard rate of males, female smokers present a significantly lower probability of remaining at work.

[21] As written in Section 3.2, in 1992 old-age pensions were granted to women aged at least 55.

[22] See Lumsdaine and Mitchell (1999) for a survey.

Finally, it should be highlighted how, *ceteris paribus*, both working in the public sector and living in the South are associated with a drop in the hazard rate of becoming not employed. This latter evidence is at least partly due to the less favorable economic conditions distinguishing this region with respect to the remainder of Italy. Indeed, at a given age, workers in the South are less likely to have accumulated the necessary years of contribution needed to claim old-age or seniority pensions owing to the higher difficulties in finding and preserving a job as compared to their counterparts living in the North and in the Centre.

3.6.3 Robustness checks

So far, the parameters on the year-variables express the evolution over time of the hazard rate of becoming not employed. Besides the mentioned cohort-effects, this dynamics can be driven by whatever uncontrolled time-varying factor, such as variations in the institutional setting or in the business-cycle. In this section we intend to check whether the previous results are still valid once we control for an ulterior time-varying measure of the state of the Italian labor market. Disentangling between different sources of time-varying dynamics permits to filter out the parameters on year dummies from the heterogeneity that should not be imputed to either changes in the Social Security or heterogeneity between cohorts.

Upholding the previous results once an alternative source of time dynamics is controlled for, may indicate stronger evidence in favor of the hypotheses that (i) younger cohorts experiment lower hazard rates because the reforms of the pension system are gradually inducing retirement postponement and (ii) during the period characterized by several pension reforms, employees have the highest propensity towards stopping working.

The choice of an appropriate economic indicator reflects two opposite requisites. On the one hand, we need an index closely related to the hazard rate of leaving employment, on the other hand we should avoid the adoption of a statistic too much depending on the labor market outcomes of the elderly because it would detect variations produced by the pension systems reforms, which are a prerogative of the time-dummies. This dilemma is addressed by considering the ISTAT time-series of quarterly employment rates for the population aged 15-64 in the period 1993-2003[23]. Although some of its time-variation can be produced by the pension system reforms, the information conveyed by this index should embrace all the dynamics influencing the labor market conditions of workers at all ages. We rearrange the original information released by ISTAT in order to obtain the average yearly employment rates of four macro-regions, namely North-West, North-East, Centre, South and Islands.

The previous specifications are augmented by the so-defined employment rates[24]. The results summarized in Table (C.5) are in line with those discussed before. The estimates of the parameters on year dummies still suggest that, *ceteris paribus*, the time evolution of the hazard rate is hump-shaped and that the youngest cohorts in the sample are more likely to remain at work than their homologues in 1995-2000. Differential effects reported in Table (C.6) for the same reference category of the previous section overall confirm the considerations made above.

Employment rates are always statistically significant predictors of the risk of stopping working. Strikingly, whereas for males higher employment rates in the population are associated with a lower probability of remaining at work, the reverse effect is found for females. This evidence is probably driven by the different evolution of the probability of being at work exhibited by the two genders over the period of reference. Table (C.7) shows that in

[23] The time series of quarterly employment rates disaggregated by macro-region for the period 1993:1-2003:4 can be downloaded from the ISTAT website http://www.istat.it/lavoro/lavret/forzedilavoro/Ric05-05/Indicatori_regionali.xls.

[24] We are forced to drop the set of dummies for the region of residence in view of their strong relationship with the employment rates.

our sample it increases at all ages for females, while for males it rises only for those aged 50-54 and decreases for older individuals. This amounts to say that although the overall employment rates increase over time, the labor market attachment of a wide part of male workers included in our sample actually diminishes bringing about the unexpected sign of this relationship.

3.7 Conclusions

Italy has engaged in aligning domestic employment rates with the targets set by the European Council in the Lisbon and Stockholm agreements. However, labor market participation is globally still below the levels settled by EU, especially when we look at the employment outcomes of the elderly. As a wide part of individuals not at work in this age-range are eligible to retirement benefits, their labor market attachment is strictly related to the characteristics of the Social Security system, which has undergone several modifications since the early nineties. In this sense, after more than a decade from the first organic reform (Amato, 1992), it is meaningful to investigate the effects of these institutional changes on the likelihood of remaining employed of older workers.

Typically, empirical investigations studying labor market transitions extract information from longitudinal datasets. In spite of their widespread utilization, panel data may present some disadvantages, such as attrition affecting the representativeness of the sample, small cross-sectional dimension and absence of variables relevant to describe the phenomenon of interest. These drawbacks are overcome by the GMM technique proposed by Güell and Hu (2006), which recovers at the individual level the probability of leaving a given initial state by combining independent cross-sections representative of the same population in different time periods.

Our empirical analysis focuses on employees aged 50-65 in 1993-2002 and calculates their hazard rate of ceasing from work within the next year by drawing data from the repeated ISTAT cross-sections Aspetti della Vita Quotidiana. The reasons motivating the adoption of this dataset are manifold. First, each wave provides us with a large sample of about 60,000 observations randomly selected from the whole population of Italian households. Second, the wide time-window covered is suited to analyze the employment consequences of a long period of institutional changes. Finally, the questionnaire is explicitly designed to describe the socioeconomic status of respondents and allows us to include in the econometric specifications an exhaustive set of individual and household characteristics.

As expected, age, household size, health, sector of employment and the region of residence are significantly associated with the labor supply dynamics of both the genders. In particular, the increasing age-profile of the hazard rate denotes the absence of incentives that can make the elderly willing to prolong their working life. Hence, the current institutional background turns out to be not in line with EU recommendations, which instead urge the adoption of policies in order to raise the labor market participation of this population group.

The time trend of the hazard rate is estimated to be hump-shaped. Conditional on age, employees at work in 1995-2000 present the highest risk of becoming not employed by the next year, whereas their counterparts in more recent cohorts exhibit a statistically significant drop in their hazard rate. These results suggest the unexpected pattern that the introduction of important pension reforms fostering retirement postponement is associated with a contemporaneous rise in the propensity towards leaving the labor force.

On the one hand, the findings of this analysis support the view that only the most recent cohorts of employees are associated with a longer working life. This effect may be due to both (i) individual willingness to work longer because of economic incentives actually discouraging early retirement and

(ii) the mechanical implementation of tighter eligibility criteria. On the other hand, we propose evidence corroborating the hypothesis that in the short-run repeated reforms lowering the overall generosity of the Social Security system induce older workers to anticipate their exit from the labor market in order to exploit the more favorable features of the current legislative scenario.

Appendix A

Appendix to Chapter 1

Table A.1: Italy, employees aged 45-70 in 2000. For each population of interest we report the percentage of PC skilled workers.

	Males	Females	Total
Age			
45-49	46.39	47.00	46.66
50-54	43.13	42.35	42.88
55-59	39.44	26.85	35.20
60+	23.76	29.73	25.36
Education			
Elementary	5.11	2.65	4.24
Middle	25.51	23.08	24.78
High	62.18	58.36	60.67
University	81.25	67.72	75.14
Region			
North	53.26	45.26	49.89
Centre	44.33	42.93	43.80
South	29.72	35.16	31.36
Total	42.25	41.89	42.12

Source: SHIW 2000. Only household heads and their spouses are considered. Sample for males (females) consists of 1330 (783) observations.

Table A.2: Italy, employees aged 45-70 in 2000. For each population of interest we report the percentage of workers using a PC at work.

	Males	Females	All
Age			
45-49	35.67	35.25	35.48
50-54	33.33	28.24	31.68
55-59	29.11	13.89	23.99
60+	14.85	13.51	14.49
Education			
At most primary	2.55	0.66	1.88
Lower secondary	14.39	17.16	15.22
Upper secondary	49.79	42.62	46.96
Tertiary	67.71	42.41	56.29
Region			
North	42.34	34.21	38.91
Centre	30.67	30.43	30.58
South	22.44	18.72	21.32
Total	32.11	28.99	30.95

Source: SHIW 2000. Only household heads and their spouses are considered. Sample for males (females) consists of 1330 (783) observations.

Table A.3: Italy, employees aged 45-70 in 2000. For each population of interest we report the sample size and the proportion of transitioners towards retirement by PC skills and PC utilization at work.

	Strict definition		Broader definition	
	Within 2002	Within 2004	Within 2002	Within 2004
Males				
Sample size (#)	604	436	635	451
Retired (overall)	14.57	25.69	18.74	28.16
With PC skills	8.00	20.20	9.96	21.36
Without PC skills	20.06	30.47	25.71	33.88
Using a PC at work	6.13	16.98	8.29	18.01
Not using a PC at work	19.13	30.69	24.16	33.79
Females				
Sample size (#)	344	247	370	274
Retired (overall)	8.72	19.84	15.14	27.74
With PC skills	7.89	18.92	11.95	21.05
Without PC skills	9.38	20.59	17.54	32.50
Using a PC at work	9.62	15.79	12.15	18.99
Not using a PC at work	8.33	21.64	16.35	31.28

Source: SHIW 2000-2004. Only household heads and their spouses are considered. Proportions are expressed in percentage terms.

Table A.4: Italy, employees aged 45-70 in 2000. Effects of PC use at work on the probability of retirement estimated by linear probability model. Specifications allow for age, education, number of days spent at home for illness during 2000, region of residence, number of household components, labor income, other household income, the age at the time of the first job, the number of years of contribution to Social Security up to 2000, job characteristics, sector of employment and firm dimensions. For each parameter we report the point estimate and the standard error between brackets.

	2000-2002			
	Strict definition		Broader definition	
	Males	Females	Males	Females
OLS estimates	-0.1043 ***	0.024	-0.0858 ***	0.0158
	(0.0299)	(0.0339)	(0.0316)	(0.0392)
2SLS estimates	-0.1048 *	-0.0402	-0.1194 *	-0.1773
	(0.0631)	(0.1003)	(0.0704)	(0.1292)
F-test	33.06 ***	10.12 ***	34.90 ***	11.19 ***
Hansen test	0.56	0.98	0.12	3.87
Hausman test	0.01	0.68	0.55	1.68 *
Num. of Obs.	604	344	635	370
	2000-2004			
	Strict definition		Broader definition	
	Males	Females	Males	Females
OLS estimates	-0.1114 ***	-0.0541	-0.0939 **	-0.0438
	(0.0416)	(0.0451)	(0.0422)	(0.0533)
2SLS estimates	-0.2402 ***	-0.2214	-0.2419 ***	-0.3462 **
	(0.0906)	(0.1431)	(0.0925)	(0.1739)
F-test	27.95 ***	8.26 ***	27.52 ***	9.53 ***
Hansen test	0.74	5.11	0.29	2.17
Hausman test	1.64	1.26	1.82 *	1.94 *
Num. of Obs.	436	247	451	274

Source: SHIW 2000-2004. Only household heads and their spouses are considered. Inference is robust to arbitrary heteroskedasticity. F-test is a test of joint insignificance of the additional instruments in the IV first stage regression. ***: p-value≤0.01, **: 0.01<p-value≤0.05, *: 0.05<p-value≤0.1.

Table A.5: Italy, employees aged 45-70 in 2000. Effects of PC skills and PC use at work on the probability of retirement within 2002 estimated by linear probability model. Specifications allow for age, education, number of days spent at home for illness during 2000, region of residence, number of household components, labor income, other household income, the age at the time of the first job, the number of years of contribution to Social Security up to 2000, job characteristics, sector of employment and firm dimensions. For each parameter we report the point estimate and the standard error between brackets.

	2000-2002			
	Strict definition		Broader definition	
	Males	Females	Males	Females
OLS				
PC use at work	-0.0446	0.0631 *	-0.0245	0.0477
	(0.0461)	(0.0375)	(0.0474)	(0.0453)
Higher education				
PC skills	-0.0800	-0.0415	-0.0795	-0.0267
	(0.0561)	(0.0400)	(0.0584)	(0.0460)
Over. eff. of PC use at work	-0.1246***	0.0216	-0.1040 **	0.0210
	(0.0390)	(0.0405)	(0.0418)	(0.0452)
Lower education				
PC skills	-0.0848 *	-0.1251 **	-0.0907	-0.1287
	(0.0508)	(0.0592)	(0.0566)	(0.0882)
Over. eff. of PC use at work	-0.1294 ***	-0.0619	-0.1152 *	-0.0811
	(0.0489)	(0.0642)	(0.0539)	(0.0887)
2SLS				
PC use at work	0.2966	-0.9387	0.1858	-1.2870
	(0.7162)	(1.6806)	(0.9819)	(1.0803)
Higher education				
PC skills	-0.3433	0.5651	-0.2510	0.7242
	(0.5912)	(1.0376)	(0.8088)	(0.6640)
Over. eff. of PC use at work	-0.0467	-0.3737	-0.0652	-0.5629
	(0.1410)	(0.6567)	(0.1867)	(0.4503)
Lower education				
PC skills	-0.1563	0.3438	-0.1514	0.3078
	(0.3618)	(0.7507)	(0.4760)	(0.5244)
Over. eff. of PC use at work	0.1403	-0.5949	0.0344	-0.9792
	(0.3875)	(0.9977)	(0.5319)	(0.7002)
Hausman test	0.46	0.42	0.07	1.58
Num. of Obs.	604	344	635	370

Source: SHIW 2000-2002. Only household heads and their spouses are considered. Inference is robust to arbitrary heteroskedasticity. ***: p-value≤0.01, **: 0.01<p-value≤0.05, *: 0.05<p-value≤0.1.

Table A.6: Italy, employees aged 45-70 in 2000. Effects of PC skills and PC use at work on the probability of retirement within 2004 estimated by linear probability model. Specifications allow for age, education, number of days spent at home for illness during 2000, region of residence, number of household components, labor income, other household income, the age at the time of the first job, the number of years of contribution to Social Security up to 2000, job characteristics, sector of employment and firm dimensions. For each parameter we report the point estimate and the standard error between brackets.

	2000-2004			
	Strict definition		Broader definition	
	Males	Females	Males	Females
OLS				
PC use at work	-0.0374	-0.0679	-0.0277	-0.0467
	(0.0662)	(0.0665)	(0.0673)	(0.0709)
Higher education				
PC skills	-0.1465 *	-0.0024	-0.1328 *	-0.0178
	(0.0769)	(0.0729)	(0.0772)	(0.0766)
Over. eff. of PC use at work	-0.184 ***	-0.0702	-0.1604 ***	-0.0645
	(0.0536)	(0.0509)	(0.0532)	(0.0617)
Lower education				
PC skills	0.0129	0.1146	0.0115	0.0905
	(0.0862)	-(0.1035)	(0.0867)	(0.1197)
Over. eff. of PC use at work	-0.0246	0.0467	-0.0162	0.0438
	(0.0740)	(0.1029)	(0.0758)	(0.1186)
2SLS				
PC use at work	0.1761	-1.0617	0.0894	-1.2608
	(0.5831)	(0.7652)	(0.6375)	(1.0110)
Higher education				
PC skills	-0.3549	0.5526	-0.2704	0.6106
	(0.4893)	(0.4900)	(0.5255)	(0.6535)
Over. eff. of PC use at work	-0.1788	-0.5091	-0.1810	-0.6502
	(0.1291)	(0.3211)	(0.1431)	(0.4053)
Lower education				
PC skills	-0.1722	0.4602	-0.1692	0.3988
	(0.3026)	(0.3832)	(0.3298)	(0.5442)
Over. eff. of PC use at work	0.0039	-0.6015	-0.0798	-0.8619
	(0.3500)	(0.6166)	(0.3722)	(0.7438)
Hausman test	0.38	1.73	0.61	2.04
Num. of Obs.	436	247	451	274

Source: SHIW 2000-2004. Only household heads and their spouses are considered. Inference is robust to arbitrary heteroskedasticity. ***: p-value$\leq$0.01, **: 0.01<p-value$\leq$0.05, *: 0.05<p-value$\leq$0.1.

Table A.7: Italy, employees aged 45-70 in 2000. Effects of PC skills and PC use at work on the probability of retirement within 2002 estimated by logit model. Specifications allow for age, education, number of days spent at home for illness during 2000, region of residence, number of household components, labor income, other household income, the age at the time of the first job, the number of years of contribution to Social Security up to 2000, job characteristics, sector of employment and firm dimensions. For each parameter we report the point estimate, the standard error between brackets and the marginal effects at the average level in italics.

	2000-2002			
	Strict definition		Broader definition	
	Males	Females	Males	Females
PC use at work	-0.8609	1.1022	-0.4554	0.3231
	(0.5726)	(1.1672)	(0.4879)	(0.7290)
	-0.0317	*0.0113*	*-0.0417*	*0.0222*
Higher education				
PC skills	-1.0875 *	-0.6374	-0.7566	-0.0127
	(0.6496)	(1.1803)	(0.5548)	(0.7371)
	-0.0437	*-0.0049*	*-0.0713*	*-0.0008*
Over. eff. of PC use at work	-1.948 ***	0.4648	-1.2112 ***	0.3105
	(0.5660)	(0.8043)	(0.4597)	(0.5974)
	-0.0755	*0.0064*	*-0.1130*	*0.0214*
Lower education				
PC skills	-0.7378	-2.3927	-0.5808	-2.0398
	(0.6198)	(1.7383)	(0.5167)	(1.1820)
	-0.0276	*-0.0122*	*-0.0534*	*-0.0676*
Over. eff. of PC use at work	-1.5987 ***	-1.2905	-1.0362 *	-1.7166
	(0.6384)	(1.4809)	(0.5284)	(1.0508)
	-0.0593	*-0.0009*	*-0.0950*	*-0.0454*
Rivers-Vuong test	2.25	2.44	0.6	4.41
Num. of Obs.	604	344	635	370

Source: SHIW 2000-2002. Only household heads and their spouses are considered. ***: p-value≤0.01, **: 0.01<p-value≤0.05, *: 0.05<p-value≤0.1.

Table A.8: Italy, employees aged 45-70 in 2000. Effects of PC skills and PC use at work on the probability of retirement within 2004 estimated by logit model. Specifications allow for age, education, number of days spent at home for illness during 2000, region of residence, number of household components, labor income, other household income, the age at the time of the first job, the number of years of contribution to Social Security up to 2000, job characteristics, sector of employment and firm dimensions. For each parameter we report the point estimate, the standard error between brackets and the marginal effects at the average level in italics.

	2000-2004			
	Strict definition		Broader definition	
	Males	Females	Males	Females
PC use at work	-0.6882	-1.3760	-0.5001	-0.6387
	(0.5680)	(0.9091)	(0.5154)	(0.6532)
	-0.0540	*-0.0143*	*-0.0579*	*-0.0864*
Higher education				
PC skills	-1.4203 ***	0.3687	-1.1532 *	-0.0871
	(0.6724)	(0.8579)	(0.6209)	(0.6315)
	-0.1192	*0.0047*	*-0.1374*	*-0.0128*
Over. eff. of PC use at work	-2.1084 ***	-1.0070	-1.6534 ***	-0.7258
	(0.5634)	(0.9152)	(0.5195)	(0.5878)
	-0.1732	*-0.0096*	*-0.1954*	*-0.0991*
Lower education				
PC skills	1.6098 *	3.2943 **	0.1931	0.8910
	(0.7468)	(1.3392)	(0.5147)	(0.9071)
	0.2219	*0.1580*	*0.0926*	*0.1701*
Over. eff. of PC use at work	-0.4987	1.9183	-0.3070	0.2523
	(0.6163)	(1.3213)	(0.5471)	(0.8995)
	-0.0487	*0.1437*	*-0.0346*	*0.0837*
Rivers-Vuong test	1.54	1.89	0.94	3.22
Num. of Obs.	436	247	451	274

Source: SHIW 2000-2004. Only household heads and their spouses are considered.
***: p-value≤0.01, **: 0.01<p-value≤0.05, *: 0.05<p-value≤0.1.

Table A.9: Italy, employees aged 45-70 in 2000. Effects of PC skills and PC use at work on the probability of retirement estimated by Cox model. Specifications allow for education, number of days spent at home for illness during 2000, region of residence, number of household components, labor income, other household income, the age at the time of the first job, the number of years of contribution to Social Security up to 2000, job characteristics, sector of employment and firm dimensions. For each parameter we report the point estimate and the standard error between brackets.

	Strict definition		Broader definition	
	Males	Females	Males	Females
PC use at work	-0.7895 ***	-0.1313	-0.4181	-0.1499
	(0.3803)	(0.5697)	(0.3438)	(0.4393)
Higher education				
PC skills	-0.7779*	-0.0188	-0.6511	-0.0353
	(0.4453)	(0.5815)	(0.4031)	(0.4392)
Over. eff. of PC use at work	-1.5674 ***	-0.1501	-1.0692 ***	-0.1852
	(0.3905)	(0.5225)	(0.3365)	(0.3964)
Lower education				
PC skills	0.1503	0.2834	0.0324	-0.3855
	(0.3903)	(0.9301)	(0.3345)	(0.7001)
Over. eff. of PC use at work	-0.6392	0.1521	-0.3857	-0.5354
	(0.4319)	(0.9274)	(0.3663)	(0.6641)
Num. of Obs.	961	556	1001	596
Num. of Ind.	608	344	639	370

Source: SHIW 2000-2004. Only household heads and their spouses are considered. Tied events are controlled by the exact partial likelihood method. ***: p-value≤0.01, **: 0.01<p-value≤0.05, *: 0.05<p-value≤0.1.

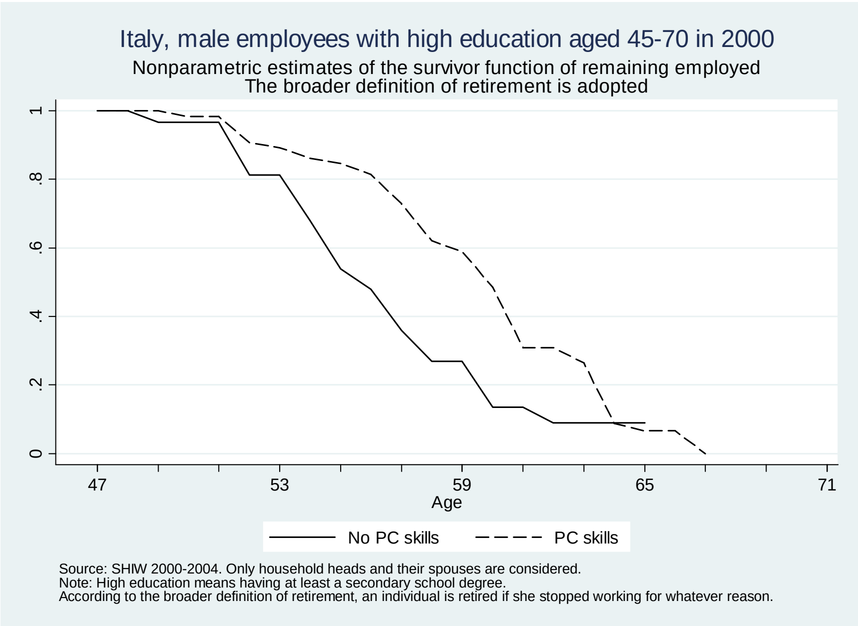

Figure A.1: Nonparametric estimates.

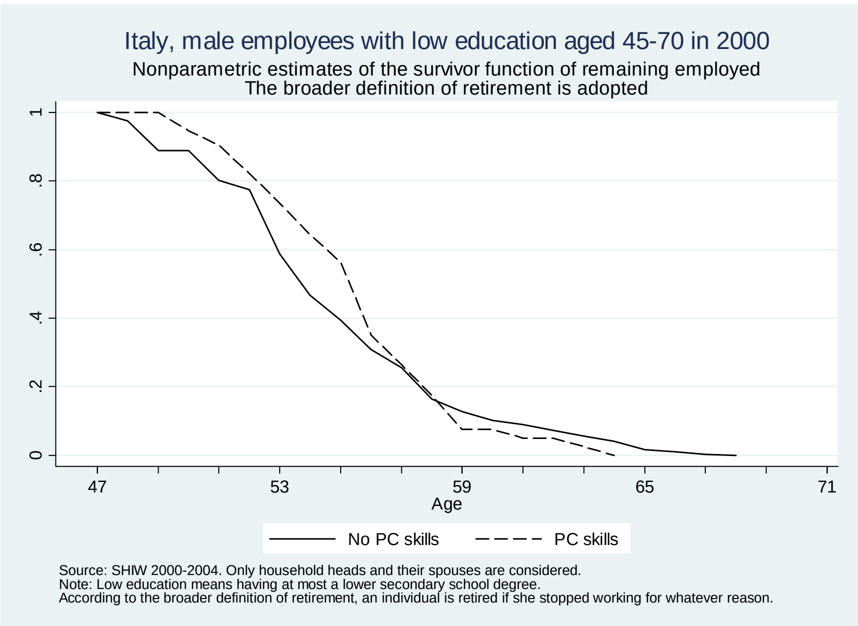

Figure A.2: Nonparametric estimates.

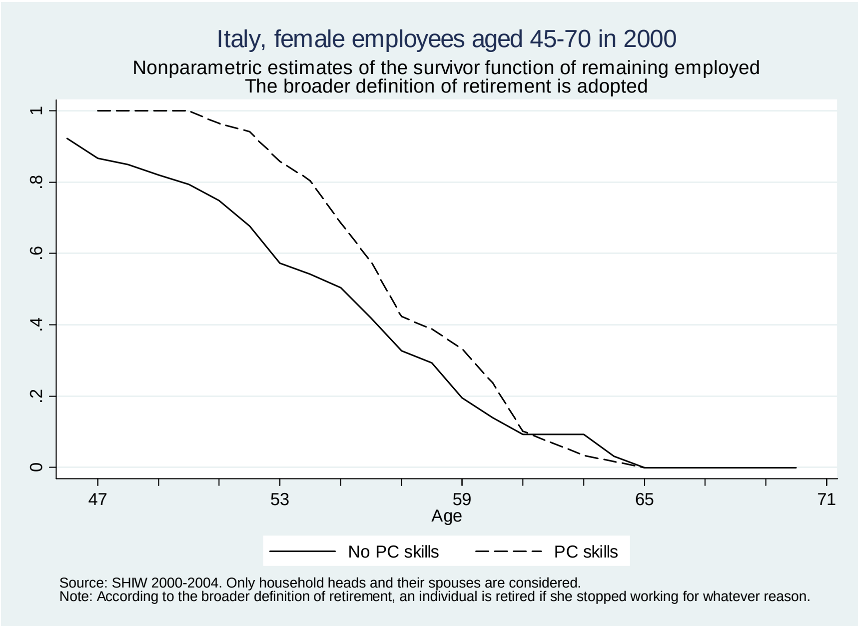

Figure A.3: Nonparametric estimates.

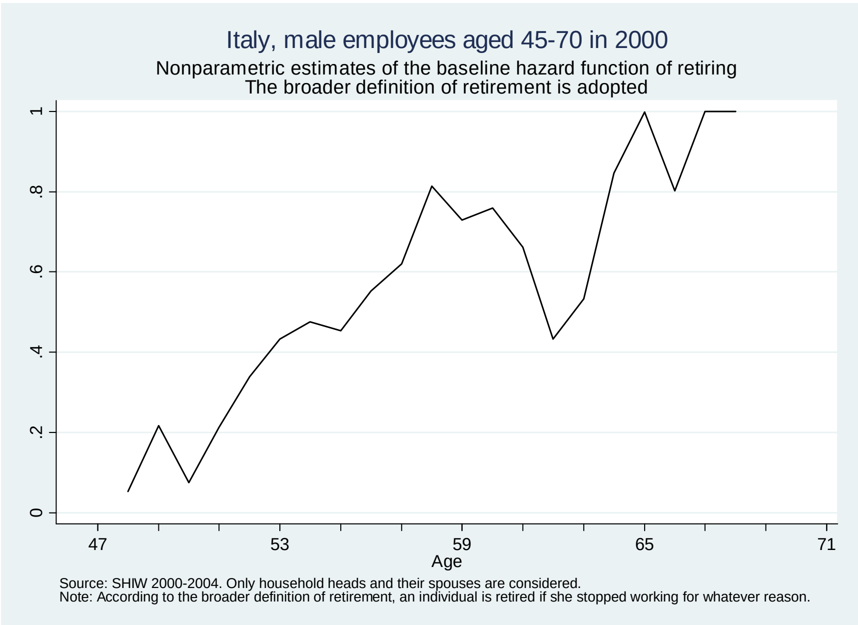

Figure A.4: Cox baseline.

Appendix B

Appendix to Chapter 2

Table B.1: Europe, married workers aged 46-65. For each population of interest we report the sample size and the number of individuals between brackets.

Country	Females		Males	
	Employees	Self-employed	Employees	Self-employed
Denmark	1,560	89	1,863	239
	(466)	(35)	(511)	(81)
Belgium	1,984	192	3,751	333
	(641)	(79)	(1102)	(123)
Netherlands	807	179	1,632	327
	(272)	(60)	(488)	(99)
Ireland	955	229	1,708	1,113
	(333)	(92)	(527)	(346)
Italy	2,040	835	4,603	2,501
	(628)	(290)	(1364)	(756)
Greece	718	1,424	2,281	3,215
	(249)	(464)	(697)	(904)
Spain	1,097	581	3,560	1,615
	(409)	(222)	(1111)	(502)
Portugal	1,919	1,505	3,282	2,253
	(577)	(481)	(923)	(668)
Austria	1,007	494	2,015	484
	(330)	(165)	(618)	(171)
UK	2,576	266	2,704	825
	(687)	(93)	(719)	(239)
Total	14,663	5,794	27,399	12,905
	(4,592)	(1,981)	(8,060)	(3,889)

Source: ECHP 1995-2001.

Table B.2: Europe, married workers aged 46-65. For each population of interest we report the sample composition in percentage terms by age, health, spouse health and spouse labor participation.

Variable	Females		Males	
	Employees	Self-employed	Employees	Self-employed
By age				
54 or less	75.60	58.03	68.59	51.39
55 or over	24.40	41.97	31.41	48.61
By health				
Good	68.44	56.51	72.47	67.05
Fair	25.79	32.33	23.25	26.35
Bad	5.77	11.17	4.28	6.60
By spouse health				
Good	67.02	54.19	66.18	64.72
Fair	24.99	31.77	25.66	25.73
Bad	7.99	14.03	8.16	9.55
By spouse labor participation				
Not employed	23.67	22.68	48.34	48.52
Employed	76.33	77.32	51.66	51.48

Source: ECHP 1995-2001. Note: Health status is defined according to self-assessments.

Table B.3: Europe, married workers aged 46-65. For each population of interest we report the proportion of transitions towards not employment stratifying the sample according to age, health, spouse health and spouse labor participation.

Variable	Females		Males	
	Employees	Self-employed	Employees	Self-employed
By age				
54 or less	7.24	12.17	3.95	2.70
55 or over	18.14	20.76	15.74	11.51
By health				
Good	8.15	14.66	6.43	4.98
Fair	12.08	14.84	9.31	9.18
Bad	20.80	24.11	19.42	18.54
By spouse health				
Good	8.84	15.35	6.56	5.16
Fair	11.62	16.02	9.20	9.25
Bad	13.32	16.85	11.67	13.22
By spouse labor participation				
Not employed	16.02	18.72	10.27	8.24
Employed	8.00	14.91	5.20	5.79
Total	9.90	15.77	7.65	6.98

Source: ECHP 1995-2001. Note: Health status is defined according to self-assessments. Proportions are expressed in percentage terms.

Table B.4: Europe, married workers aged 46-65. For each population of interest we report by country the proportions of not-employed individuals moving towards employment within the next year.

Country	Females	Males
Denmark	10.08	10.16
Belgium	5.06	4.25
Netherlands	2.60	3.28
Ireland	6.40	11.34
Italy	2.80	5.85
Greece	4.65	8.22
Spain	4.13	8.50
Portugal	7.42	10.26
Austria	3.46	3.42
UK	6.44	7.72
Total	4.67	7.02

Source: ECHP 1995-2001.

Table B.5: Europe, married workers aged 46-65. Results of the log-rank test maintaining the null hypothesis of equality among the survivor functions of the groups obtained stratifying the sample by health, spouse health and spouse labor participation.

Variable	Females		Males	
	Employees	Self-employed	Employees	Self-employed
Health				
	105.32	22.19	168.21	122.58
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Spouse health				
	8.82	1.75	19.22	40.84
	(0.0122)	(0.4163)	(0.0001)	(0.0000)
Spouse labor participation				
	35.96	0.45	71.37	3.24
	(0.0000)	(0.5027)	(0.0000)	(0.0718)

Source: ECHP 1995-2001. Note: We report the test statistics and the p-value between brackets. Health status is defined according to self-assessments.

Table B.6: Europe, married workers aged 46-65. IV linear probability model estimates of the effects on the likelihood of transition towards not employment produced by discrete changes in spouse employment position and couple members health. Health is defined according to individual self-assessments. Raw differences in the sample are reported in italics.

Variable	Females		Males	
	Employees	Self-employed	Employees	Self-employed
All				
Employed spouse	-9.85 **	-33.90 ***	-18.89 **	-23.13 **
	-8.03	*-3.81*	*-5.07*	*-2.45*
Aged 54 or less				
Good health	-7.65 ***	-6.11 **	-7.88 ***	-4.51 **
	-11.07	*-5.48*	*-11.71*	*-8.62*
Fair health	-5.93 ***	-6.64 ***	-7.51 ***	-2.49
	-8.63	*-7.07*	*-10.25*	*-6.13*
Spouse good health	4.57 ***	9.74 ***	4.13 *	1.38
	-1.96	*1.63*	*-3.15*	*-4.25*
Spouse fair health	3.52 **	6.49 ***	3.49 **	1.61
	-0.54	*0.64*	*-1.57*	*-2.24*
Aged 55 or over				
Good health	-8.15 ***	-11.20 ***	-6.67 ***	-7.29 ***
	-11.18	*-9.39*	*-10.82*	*-14.12*
Fair health	-2.56	-9.51 ***	-4.58 **	-4.71 **
	-5.05	*-9.19*	*-7.83*	*-10.17*
Spouse good health	4.48 *	7.03 *	4.22 **	3.16
	-3.98	*-1.49*	*-2.87*	*-7.04*
Spouse fair health	3.25	3.47	3.44 **	3.79
	-0.82	*-1.63*	*-0.60*	*-2.82*

Source: ECHP 1995-2001. Note: Effects are expressed in terms of percentage points. The baselines are (i) spouse not at work, (ii) bad health and (iii) spouse bad health. Details on the estimation results are provided by Tables (B.12) and (B.13). ***: p-value ≤ 0.01, **: $0.01 <$ p-value ≤ 0.05, *: $0.05 <$ p-value ≤ 0.1.

Table B.7: Europe, married workers aged 46-65. IV linear probability model estimates of the effects on the likelihood of transition towards not employment produced by changes in spouse employment position and couple members health. Health is measured according to a linear index obtained following Bound et al. (1999). Raw differences in the sample are reported in italics.

Variable	Females		Males	
	Employees	Self-employed	Employees	Self-employed
All				
Employed spouse	-9.69 **	-31.75 ***	-19.33 **	-23.00
	-8.03	*-3.81*	*-5.07*	*-2.45*
Aged 54 or less				
Health	-1.98 ***	-3.20 ***	-3.23 ***	-3.28 ***
Spouse health	1.14	2.78 **	1.33 **	0.69
Aged 55 or over				
Health	-2.43 ***	-3.66 ***	-2.78 ***	-3.39 ***
Spouse health	1.44 **	3.11 ***	1.28 **	0.89

Source: ECHP 1995-2001. Note: Effects are expressed in terms of percentage points. The baseline for spouse employment position is being not at work. Details on the estimation results are provided by Tables (B.14) and (B.15). ***: p-value ≤ 0.01, **: $0.01 <$ p-value ≤ 0.05, *: $0.05 <$ p-value ≤ 0.1.

Table B.8: Europe, married female workers aged 46-65. Effects on the probability of transition towards not employment within the next year estimated by OLS linear probability model. Specifications allow for time and country dummies, spouses age and education, household size, number of children aged less than 16, labor income, job characteristics, sector of employment, years of contribution up to the time the worker enters the sample, other household income. For each parameter we report the point estimate and the standard error between brackets.

Variable	Employees	Self-employed
Employed spouse	-0.0211 ***	-0.0442 ***
	(0.0073)	(0.0140)
Aged 54 or less		
Good health	-0.0741 ***	-0.0558 **
	(0.0158)	(0.0248)
Fair health	-0.0575 ***	-0.0581 **
	(0.0162)	(0.0249)
Spouse good health	0.0228 **	0.0141
	(0.0106)	(0.0193)
Spouse fair health	0.0176	0.0048
	(0.0110)	(0.0195)
Aged 55 or over		
Good health	-0.0762 ***	-0.1022 ***
	(0.0251)	(0.0294)
Fair health	-0.0219	-0.0910 ***
	(0.0259)	(0.0269)
Spouse good health	0.0219	0.0009
	(0.0194)	(0.0267)
Spouse fair health	0.0180	-0.0115
	(0.0201)	(0.0241)
$H_0 : c_i = 0$	21.62 ***	6.56 ***
Num. Of Obs.	14,663	5,794
Num. Of Ind.	4,592	1,981

Source: ECHP 1995-2001. Note: Inference is robust to arbitrary heteroskedasticity and autocorrelation of the error terms at the individual level. Health is defined according to individual self-assessments. $H_0 : c_i = 0$ is a test of joint insignificance for all the parameters of the specification assumed to describe c_i. ***: p-value $\leq$ 0.01, **: 0.01 $<$ p-value $\leq$ 0.05, *: 0.05 $<$ p-value $\leq$ 0.1.

Table B.9: Europe, married male workers aged 46-65. Effects on the probability of transition towards not employment within the next year estimated by OLS linear probability model. Specifications allow for time and country dummies, spouses age and education, household size, number of children aged less than 16, labor income, job characteristics, sector of employment, years of contribution up to the time the worker enters the sample, other household income. For each parameter we report the point estimate and the standard error between brackets.

Variable	Employees	Self-employed
Employed spouse	-0.0188 ***	-0.0198 ***
	(0.0038)	(0.0048)
Aged 54 or less		
Good health	-0.0702 ***	-0.0512 ***
	(0.0134)	(0.0162)
Fair health	-0.0711 ***	-0.0365 **
	(0.0135)	(0.0165)
Spouse good health	-0.0023	-0.0225 *
	(0.0072)	(0.0118)
Spouse fair health	0.0044	-0.0117
	(0.0075)	(0.0122)
Aged 55 or over		
Good health	-0.0555 ***	-0.0669 ***
	(0.0193)	(0.0186)
Fair health	-0.0391 **	-0.0522 ***
	(0.0197)	(0.0185)
Spouse good health	0.0173	-0.0104
	(0.0130)	(0.0134)
Spouse fair health	0.0188	0.0048
	(0.0135)	(0.0136)
$H_0 : c_i = 0$	23.58 ***	9.96 ***
Num. Of Obs.	27,399	12,905
Num. Of Ind.	8,060	3,889

Source: ECHP 1995-2001. Note: Inference is robust to arbitrary heteroskedasticity and autocorrelation of the error terms at the individual level. Health is defined according to individual self-assessments. $H_0 : c_i = 0$ is a test of joint insignificance for all the parameters of the specification assumed to describe c_i. ***: p-value $\leq$ 0.01, **: 0.01 $<$ p-value $\leq$ 0.05, *: 0.05 $<$ p-value $\leq$ 0.1.

Table B.10: Europe, married female workers aged 46-65. Effects on the probability of transition towards not employment within the next year estimated by OLS linear probability model. Specifications allow for time and country dummies, spouses age and education, household size, number of children aged less than 16, labor income, job characteristics, sector of employment, years of contribution up to the time the worker enters the sample, other household income. For each parameter we report the point estimate and the standard error between brackets.

Variable	Employees	Self-employed
Employed spouse	-0.0191 ***	-0.0425 ***
	(0.0072)	(0.0131)
Aged 54 or less		
Health	-0.0188 ***	-0.0348 ***
	(0.0069)	(0.011)
Spouse health	0.0039	0.0062
	(0.0059)	(0.0091)
Aged 55 or over		
Health	-0.0232 ***	-0.0386 ***
	(0.007)	(0.0111)
Spouse health	0.0070	0.0087
	(0.0061)	(0.0094)
$H_0 : c_i = 0$	139.78 ***	45.50 ***
Num. Of Obs.	14,663	5,794
Num. Of Ind.	4,592	1,981

Source: ECHP 1995-2001. Note: Inference is robust to arbitrary heteroskedasticity and autocorrelation of the error terms at the individual level. Health is measured according to a linear index obtained following Bound et al. (1999). Standard errors are computed bootstrapping 500 times the two-step estimation strategy. $H_0 : c_i = 0$ is a test of joint insignificance for all the parameters of the specification assumed to describe c_i. ***: p-value $\leq$ 0.01, **: 0.01 $<$ p-value $\leq$ 0.05, *: 0.05 $<$ p-value $\leq$ 0.1.

Table B.11: Europe, married male workers aged 46-65. Effects on the probability of transition towards not employment within the next year estimated by OLS linear probability model. Specifications allow for time and country dummies, spouses age and education, household size, number of children aged less than 16, labor income, job characteristics, sector of employment, years of contribution up to the time the worker enters the sample, other household income. For each parameter we report the point estimate and the standard error between brackets.

Variable	Employees	Self-employed
Employed spouse	-0.0185 ***	-0.0209 ***
	(0.0039)	(0.0047)
Aged 54 or less		
Health	-0.0311 ***	-0.0253 ***
	(0.0044)	(0.0049)
Spouse health	0.0055	-0.0077
	(0.0040)	(0.0050)
Aged 55 or over		
Health	-0.0264 ***	-0.0267 ***
	(0.0047)	(0.0051)
Spouse health	0.0051	-0.0061
	(0.0041)	(0.0050)
$H_0 : c_i = 0$	132.84 ***	62.27 ***
Num. Of Obs.	27,399	12,905
Num. Of Ind.	8,060	3,889

Source: ECHP 1995-2001. Note: Inference is robust to arbitrary heteroskedasticity and autocorrelation of the error terms at the individual level. Health is measured according to a linear index obtained following Bound et al. (1999). Standard errors are computed bootstrapping 500 times the two-step estimation strategy. $H_0 : c_i = 0$ is a test of joint insignificance for all the parameters of the specification assumed to describe c_i. ***: p-value $\leq$ 0.01, **: 0.01 $<$ p-value $\leq$ 0.05, *: 0.05 $<$ p-value $\leq$ 0.1.

Table B.12: Europe, married female workers aged 46-65. Effects on the probability of transition towards not employment within the next year estimated by IV linear probability model. Specifications allow for time and country dummies, spouses age and education, household size, number of children aged less than 16, labor income, job characteristics, sector of employment, years of contribution up to the time the worker enters the sample, other household income. For each parameter we report the point estimate and the standard error between brackets.

Variable	Employees	Self-employed
Employed spouse	-0.0985 **	-0.3390 ***
	(0.0463)	(0.0998)
Aged 54 or less		
Good health	-0.0765 ***	-0.0611 **
	(0.0158)	(0.0254)
Fair health	-0.0593 ***	-0.0664 ***
	(0.0162)	(0.0254)
Spouse good health	0.0457 ***	0.0974 ***
	(0.0171)	(0.0351)
Spouse fair health	0.0352 **	0.0649 ***
	(0.0151)	(0.0293)
Aged 55 or over		
Good health	-0.0815 ***	-0.1120 ***
	(0.0254)	(0.0307)
Fair health	-0.0256	-0.0951 ***
	(0.0262)	(0.0281)
Spouse good health	0.0448 *	0.0703 *
	(0.0240)	(0.0373)
Spouse fair health	0.0325	0.0347
	(0.0225)	(0.0299)
$H_0 : c_i = 0$	133.00 ***	40.35 ***
F-test	112.07 ***	26.23 ***
Hansen test	0.29	0.31
Hausman test	3.05 *	9.81 ***
Num. Of Obs.	14,663	5,794
Num. Of Ind.	4,592	1,981

Source: ECHP 1995-2001. Note: Inference is robust to arbitrary heteroskedasticity and autocorrelation of the error terms at the individual level. Health is defined according to individual self-assessments. $H_0 : c_i = 0$ is a test of joint insignificance for all the parameters of the specification assumed to describe c_i. The IV approach controls for the endogeneity of spouse employment status. F-test is a test of joint insignificance of the additional instruments in the IV first-stage regression. ***: p-value $\leq$ 0.01, **: 0.01 $<$ p-value $\leq$ 0.05, *: 0.05 $<$ p-value $\leq$ 0.1.

Table B.13: Europe, married male workers aged 46-65. Effects on the probability of transition towards not employment within the next year estimated by IV linear probability model. Specifications allow for time and country dummies, spouses age and education, household size, number of children aged less than 16, labor income, job characteristics, sector of employment, years of contribution up to the time the worker enters the sample, other household income. For each parameter we report the point estimate and the standard error between brackets.

Variable	Employees	Self-employed
Employed spouse	-0.1889 **	-0.2313 **
	(0.0878)	(0.1243)
Aged 54 or less		
Good health	-0.0788 ***	-0.0451 **
	(0.0148)	(0.0186)
Fair health	-0.0751 ***	-0.0249
	(0.0143)	(0.0196)
Spouse good health	0.0413 *	0.0138
	(0.0237)	(0.0261)
Spouse fair health	0.0349 **	0.0161
	(0.0176)	(0.0221)
Aged 55 or over		
Good health	-0.0667 ***	-0.0729 ***
	(0.0208)	(0.0202)
Fair health	-0.0458 **	-0.0471 **
	(0.0207)	(0.0196)
Spouse good health	0.0422 **	0.0316
	(0.0188)	(0.0290)
Spouse fair health	0.0344 **	0.0379
	(0.0164)	(0.0249)
$H_0 : c_i = 0$	127.65 ***	54.46 ***
F-test	17.95	8.51 ***
Hansen test	0.32	0.07
Hausman test	4.18 **	3.64 *
Num. Of Obs.	27,399	12,905
Num. Of Ind.	8,060	3,889

Source: ECHP 1995-2001. Inference is robust to arbitrary heteroskedasticity and autocorrelation of the error terms at the individual level. Health is defined according to individual self-assessments. $H_0 : c_i = 0$ is a test of joint insignificance for all the parameters of the specification assumed to describe c_i. The IV approach controls for the endogeneity of spouse employment status. F-test is a test of joint insignificance of the additional instruments in the IV first-stage regression. ***: p-value ≤ 0.01, **: $0.01 <$ p-value ≤ 0.05, *: $0.05 <$ p-value ≤ 0.1.

Table B.14: Europe, married female workers aged 46-65. Effects on the probability of transition towards not employment within the next year estimated by IV linear probability model. Specifications allow for time and country dummies, spouses age and education, household size, number of children aged less than 16, labor income, job characteristics, sector of employment, years of contribution up to the time the worker enters the sample, other household income. For each parameter we report the point estimate and the standard error between brackets.

Variable	Employees	Self-employed
Employed spouse	-0.0969 **	-0.3175 ***
	(0.0481)	(0.1004)
Aged 54 or less		
Health	-0.0198 ***	-0.0320 ***
	(0.0070)	(0.0114)
Spouse health	0.0114	0.0278 **
	(0.0073)	(0.0123)
Aged 55 or over		
Health	-0.0243 ***	-0.0366 ***
	(0.0070)	(0.0116)
Spouse health	0.0144 **	0.0311 ***
	(0.0075)	(0.0126)
$H_0 : c_i = 0$	141.70 ***	45.63 ***
F-test	197.64 ***	53.14 ***
Hansen test	0.16	0.30
Hausman test	2.73 *	7.64 ***
Num. Of Obs.	14,663	5,794
Num. Of Ind.	4,592	1,981

Source: ECHP 1995-2001. Note: Inference is robust to arbitrary heteroskedasticity and autocorrelation of the error terms at the individual level. Health is measured according to a linear index obtained following Bound et al. (1999). Standard errors are computed bootstrapping 500 times the two-step estimation strategy. $H_0 : c_i = 0$ is a test of joint insignificance for all the parameters of the specification assumed to describe c_i. The IV approach controls for the endogeneity of spouse employment status. F-test is a test of joint insignificance of the additional instruments in the IV first-stage regression. ***: p-value ≤ 0.01, **: $0.01 <$ p-value ≤ 0.05, *: $0.05 <$ p-value ≤ 0.1.

Table B.15: Europe, married male workers aged 46-65. Effects on the probability of transition towards not employment within the next year estimated by IV linear probability model. Specifications allow for time and country dummies, spouses age and education, household size, number of children aged less than 16, labor income, job characteristics, sector of employment, years of contribution up to the time the worker enters the sample, other household income. For each parameter we report the point estimate and the standard error between brackets.

Variable	Employees	Self-employed
Employed spouse	-0.1933 **	-0.2300
	(0.0862)	(0.1453)
Aged 54 or less		
Health	-0.0323 ***	-0.0328 ***
	(0.0048)	(0.0077)
Spouse health	0.0133 **	0.0069
	(0.0057)	(0.0117)
Aged 55 or over		
Health	-0.0278 ***	-0.0339 ***
	(0.0050)	(0.0078)
Spouse health	0.0128 **	0.0089
	(0.0057)	(0.0119)
$H_0 : c_i = 0$	122.46 ***	52.65 ***
F-test	34.41 ***	15.68 ***
Hansen test	0.20	0.15
Hausman test	4.09 **	2.09
Num. Of Obs.	27,399	12,905
Num. Of Ind.	8,060	3,889

Source: ECHP 1995-2001. Note: Inference is robust to arbitrary heteroskedasticity and autocorrelation of the error terms at the individual level. Health is measured according to a linear index obtained following Bound et al. (1999). Standard errors are computed bootstrapping 500 times the two-step estimation strategy. $H_0 : c_i = 0$ is a test of joint insignificance for all the parameters of the specification assumed to describe c_i. The IV approach controls for the endogeneity of spouse employment status. F-test is a test of joint insignificance of the additional instruments in the IV first-stage regression. ***: p-value ≤ 0.01, **: $0.01 <$ p-value ≤ 0.05, *: $0.05 <$ p-value ≤ 0.1.

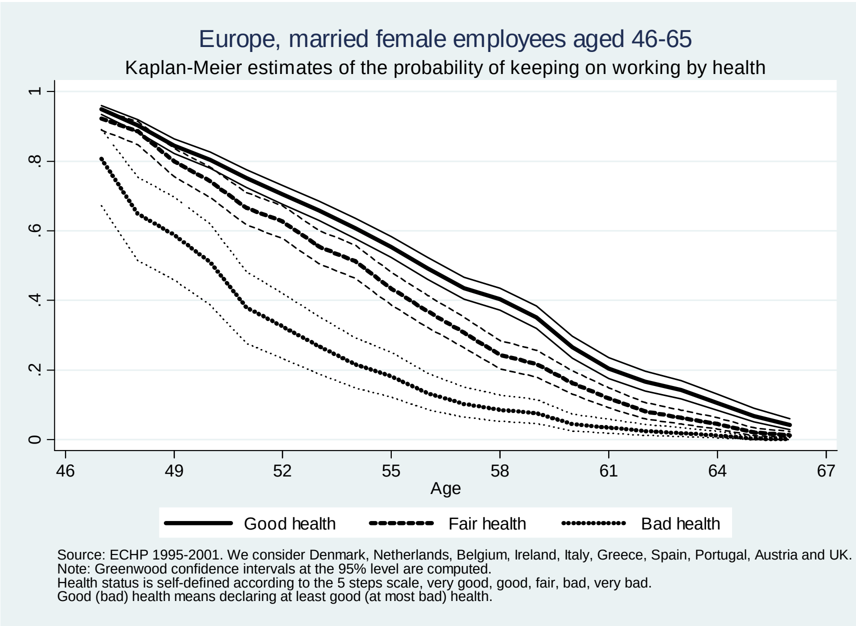

Figure B.1: Nonparametric duration analysis.

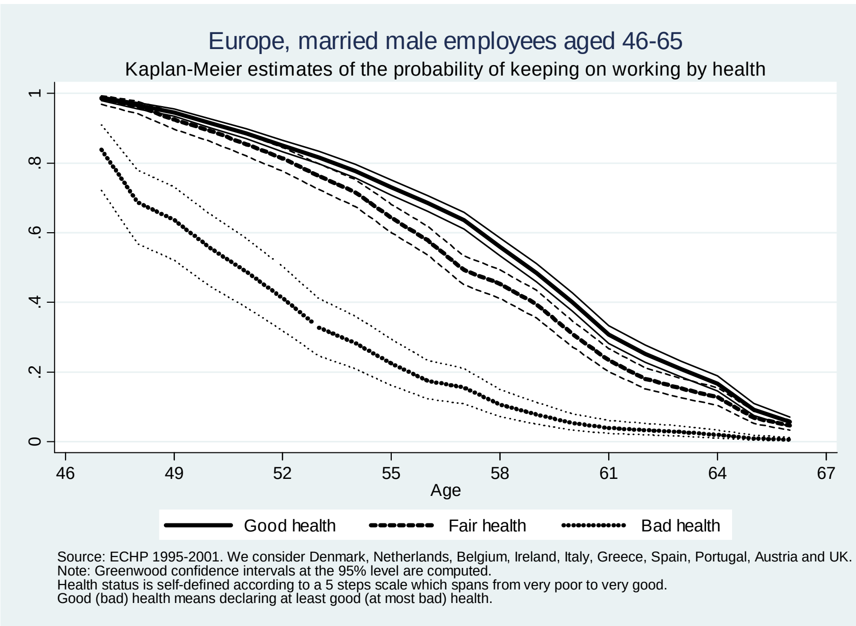

Figure B.2: Nonparametric duration analysis.

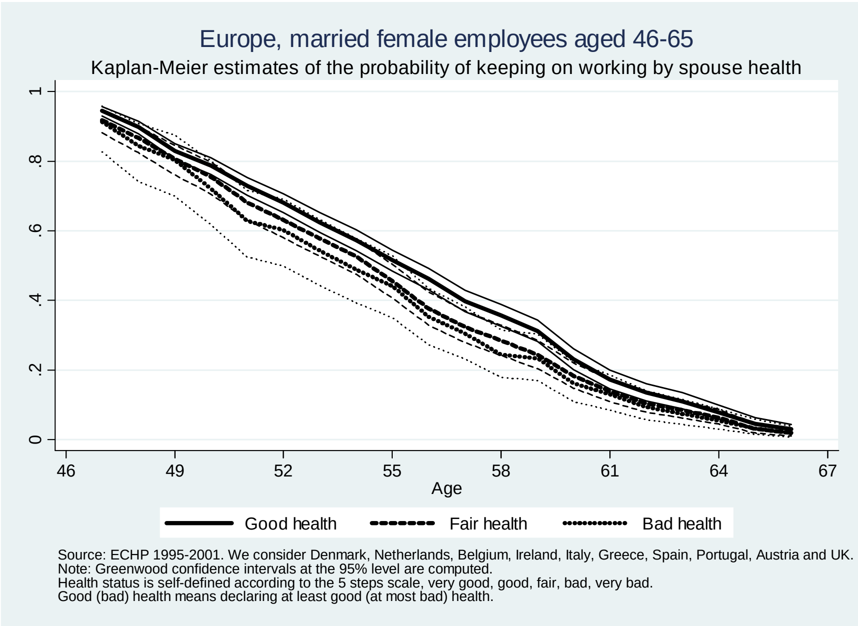

Figure B.3: Nonparametric duration analysis.

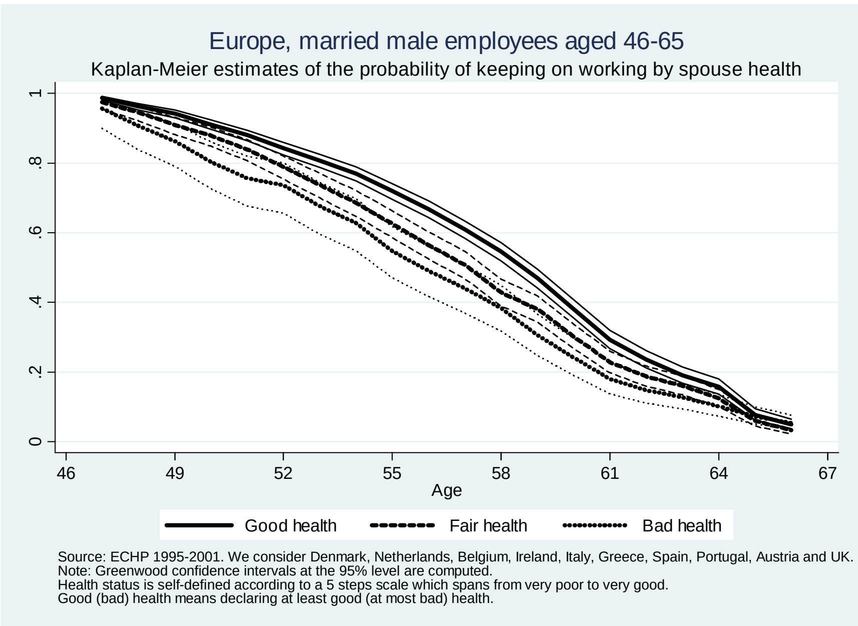

Figure B.4: Nonparametric duration analysis.

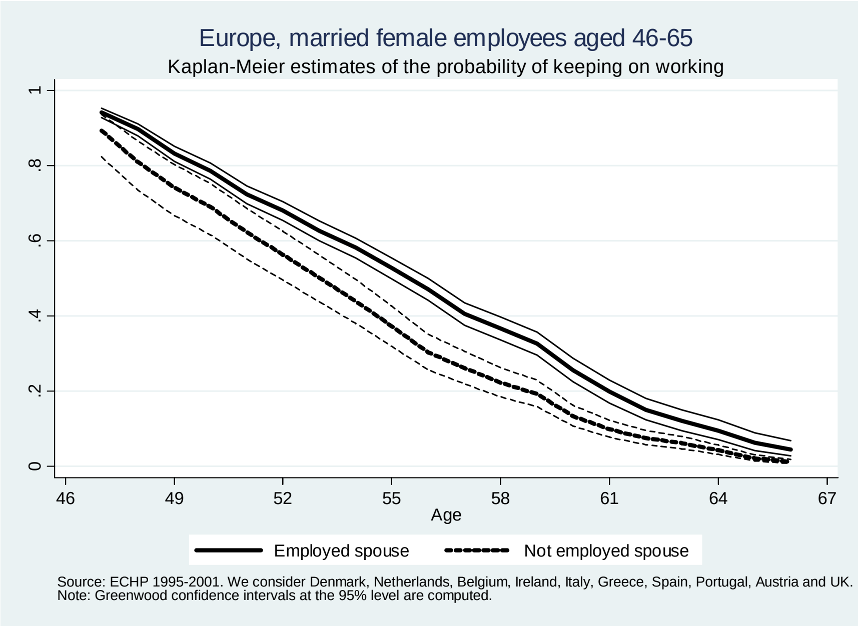

Figure B.5: Nonparametric duration analysis.

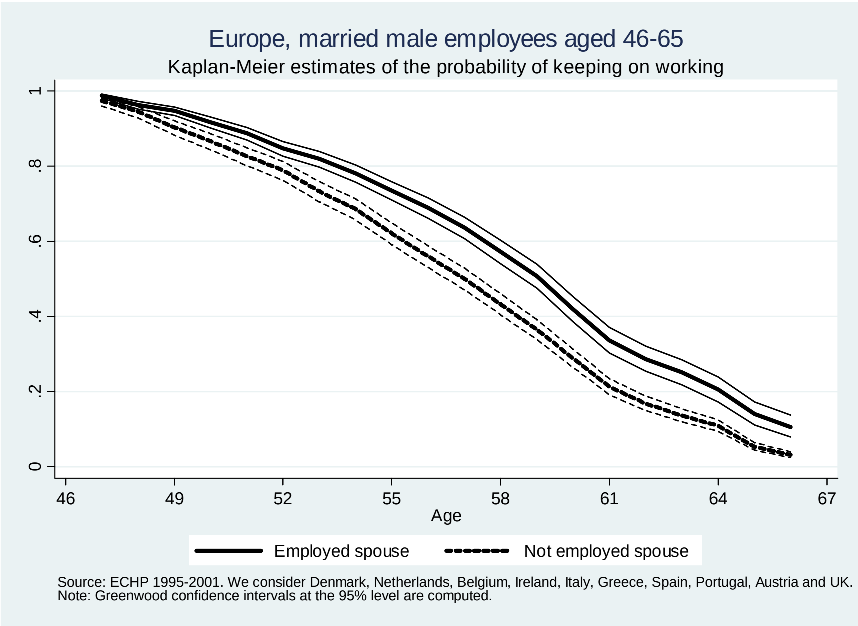

Figure B.6: Nonparametric duration analysis.

Appendix C

Appendix to Chapter 3

Table C.1: Italy, employees aged 50-65 in 1993-2002 and 51-66 in 1994-2003. Age distribution (%) and sample size of the entry and exit cross-sections used to implement the GMM estimator.

Age	Males	Females
	Entry Cross-section	
50	13.61	15.57
51	12.46	14.54
52	11.76	12.75
53	10.51	10.87
54	9.13	10.40
55	7.70	8.29
56	7.10	6.27
57	5.68	5.36
58	4.92	4.44
59	4.52	3.40
60	3.66	2.36
61	2.76	1.70
62	2.35	1.38
63	1.92	1.38
64	1.34	0.81
65	0.57	0.48
Sample size	17,725	8,393
	Exit Cross-section	
51	14.35	17.07
52	13.75	15.55
53	12.04	12.46
54	10.69	12.04
55	8.90	9.84
56	8.24	7.56
57	6.52	6.55
58	5.51	5.26
59	5.08	4.00
60	4.13	2.85
61	3.17	1.86
62	2.70	1.59
63	2.16	1.62
64	1.57	0.95
65	0.68	0.50
66	0.51	0.30
Sample size	15,580	7,544

Source: ISTAT, Aspetti della Vita Quotidiana 1993-2003.

Table C.2: Italy, employees aged 50-65 in 1993-2002. GMM logit model estimates of the effects on the hazard rate of leaving employment within the next year. For each parameter we report the point estimate and the standard error between brackets.

Variable	Males	Females
Years 1995-1996	0.3974	0.9139 *
	(0.2474)	(0.5398)
Years 1997-1998	0.3429 *	1.0508 **
	(0.2058)	(0.4465)
Years 1999-2000	0.3713 **	0.6072
	(0.2132)	(0.4925)
Years 2001-2002	-0.9509 **	-0.8626
	(0.3991)	(0.8389)
Age 50	-2.6840 ***	/
	(0.4039)	-4.1920 ***
Age 51	-3.6278 ***	(0.7880)
	(1.0458)	/
Age 52	-2.3225 ***	/
	(0.3586)	-3.1092 ***
Age 53	-2.2235 ***	(0.5499)
	(0.3555)	/
Age 54	-1.8489 ***	-2.1577 ***
	(0.3014)	(0.5150)
Age 55	-2.8826 ***	-1.9239 ***
	(0.6757)	(0.5011)
Age 56	-1.5400 ***	-3.2975 ***
	(0.2689)	(1.1374)
Age 57	-1.9228 ***	-2.5546 ***
	(0.3561)	(0.7274)
Age 58	-2.4451 ***	-2.0121 ***
	(0.5625)	(0.5797)
Age 59	-1.5534 ***	-1.6060 ***
	(0.3122)	(0.5499)

Table C.2 *continued*

Variable	Males	Females
Age 60	-1.3316 ***	-1.3373 ***
	(0.2978)	(0.5803)
Age 61	-1.9169 ***	-2.1332 **
	(0.5051)	(0.8942)
Age 62	-1.5249 ***	/
	(0.4192)	-2.1174 ***
Age 63	-0.9936 ***	(0.5847)
	(0.3425)	/
Age 64	0.2358	-0.6798
	(0.2755)	(0.6283)
Age 65	-1.4372 *	-0.5087
	(0.7568)	(0.7402)
Entry cross-section	17,725	8,393
Exit cross-section	15,580	7,544

Source: ISTAT, Aspetti della Vita Quotidiana 1993-2003. Note: The baseline is being employed in 1993-1994. ***: p-value $\leq$0.01, **: 0.01 $<$ p-value $\leq$ 0.05, *: 0.05 $<$ p-value $\leq$ 0.1.

Table C.3: Italy, employees aged 50-65 in 1993-2002. GMM logit model estimates of the effects on the hazard rate of leaving employment within the next year. For each parameter we report the point estimate and the standard error between brackets.

Variable	Males	Females
Year 1995-1996	0.4751 *	1.3620 **
	(0.2608)	(0.6935)
Year 1997-1998	0.4657 **	1.6760 ***
	(0.2184)	(0.5988)
Year 1999-2000	0.5304 **	1.0461 *
	(0.2249)	(0.6355)
Year 2001-2002	-0.8135 **	-0.6613
	(0.4063)	(0.9776)
Age 50-51	-2.4457 ***	-4.5338 ***
	(0.3014)	(0.9421)
Age 52-53	-1.5928 ***	-2.9879 ***
	(0.2048)	(0.5502)
Age 54-55	-1.4717 ***	-1.4752 ***
	(0.2297)	(0.3407)
Age 56-57	-0.7782 ***	-2.3106 ***
	(0.1766)	(0.6084)
Age 58-59	-0.9760 ***	-0.6327
	(0.2523)	(0.4182)
HH size=1	-1.5572 ***	-4.2351 ***
	(0.2840)	(0.7122)
HH size=2	-1.7508 ***	-3.2535 ***
	(0.2121)	(0.4687)
HH size=3	-0.9524 ***	-2.1983 ***
	(0.1263)	(0.3708)
HH size=4	-0.4073 ***	-1.2236 ***
	(0.1123)	(0.3074)
Hospitalization	0.7618 ***	1.7027 ***
	(0.1280)	(0.3619)

Table C.3 *continued*

Variable	Males	Females
Smoker	0.1451	0.4937 **
	(0.1120)	(0.2436)
Smoker in the past	-0.1349	-0.2812
	(0.1167)	(0.3037)
Health insurance	0.0119	-0.2395
	(0.1123)	(0.2968)
Life insurance	0.1158	-0.0101
	(0.1154)	(0.2764)
Home ownership	0.0330	-0.4789 **
	(0.1082)	(0.2307)
University	-0.2279	1.7383 ***
	(0.2072)	(0.4666)
Upper sec. school	-0.0329	0.6306
	(0.1579)	(0.3996)
Lowe sec. school	0.3670 ***	0.4229
	(0.1256)	(0.3175)
White collar	0.0641	-0.3030
	(0.1305)	(0.3471)
Primary, industry	0.3820 **	1.5305 ***
	(0.1326)	(0.3545)
Services	0.4271 ***	1.2706 ***
	(0.1282)	(0.3038)
North	1.0787 ***	1.2050 ***
	(0.1195)	(0.2994)
Centre	0.8771 ***	0.6964 **
	(0.1276)	(0.3024)
Intercept	-1.3820 ***	-0.5557
	(0.2497)	(0.6477)
Entry cross-section	17,725	8,393
Exit cross-section	15,580	7,544

Source: ISTAT, Aspetti della Vita Quotidiana 1993-2003. Note: The baselines are (i) years 1993-1994, (ii) age 60-65, (iii) household size higher than four components, (iv) not having been hospitalized during the last twelve months, (v) non-smoker, (vi) not having subscribed either a health insurance or a life insurance, (vii) non-homeowner, (viii) at most primary school, (ix) blue collar, (x) public administration, (xi) South and Islands. ***: p-value $\leq$0.01, **: 0.01 $<$ p-value $\leq$ 0.05, *: 0.05 $<$ p-value $\leq$ 0.1.

Table C.4: Italy, employees aged 50-65 in 1993-2002. With respect to the reference category described below and for the age-intervals 56-57 and 60-65, we report the hazard rate levels over the pairs of years considered and the corresponding relative deviations from the baseline period (1993-1994). Details on the specifications are reported in Table C.3.

Variable	56-57	60-65
Males		
Years 1993-1994	0.1697	0.3080
Years 1995-1996	0.2474	0.4172
	45.77%	35.44%
Years 1997-1998	0.2457	0.4149
	44.74%	34.70%
Years 1999-2000	0.2578	0.4307
	51.92%	39.83%
Years 2001-2002	0.0831	0.1648
	-51.04%	-46.49%
Females		
Years 1993-1994	0.0772	0.4576
Years 1995-1996	0.2463	0.7671
	218.88%	67.63%
Years 1997-1998	0.3091	0.8185
	300.15%	78.85%
Years 1999-2000	0.1924	0.7060
	149.12%	54.28%
Years 2001-2002	0.0414	0.3034
	-46.38%	-33.70%

Baseline category: (i) years 1993-1994, (ii) HH size=3, (iii) no hospitalization during the last twelve months, (v) non-smoker, (vi) no subscription of either health or life insurances, (vii) homeowner, (viii) upper secondary school, (ix) white collar, (x) employed in the primary or industry sector, (xi) living in the North.

Table C.5: Italy, employees aged 50-65 in 1993-2002. GMM logit model estimates of the effects on the hazard rate of leaving employment within the next year. For each parameter we report the point estimate and the standard error between brackets.

Variable	Males	Females
Year 1995-1996	0.5150 **	1.5127 **
	(0.2560)	(0.7418)
Year 1997-1998	0.4728 **	1.9308 ***
	(0.2157)	(0.6574)
Year 1999-2000	0.4996 **	1.4835 **
	(0.2196)	(0.6962)
Year 2001-2002	-0.8820 **	-0.2039
	(0.4001)	(1.0213)
Age 50-51	-2.3250 ***	-4.3444 ***
	(0.2945)	(0.9673)
Age 52-53	-1.5027 ***	-2.7970 ***
	(0.1996)	(0.5788)
Age 54-55	-1.3933 ***	-1.2615 ***
	(0.2236)	(0.3498)
Age 56-57	-0.7382 ***	-2.2354 ***
	(0.1703)	(0.6288)
Age 58-59	-0.9352 ***	-0.5156
	(0.2444)	(0.4216)
HH size=1	-1.3439 ***	-3.8234 ***
	(0.2780)	(0.7322)
HH size=2	-1.5844 ***	-2.8044 ***
	(0.2052)	(0.4704)
HH size=3	-0.8036 ***	-1.6718 ***
	(0.1209)	(0.3569)
HH size=4	-0.3245 ***	-0.8351 ***
	(0.1078)	(0.3055)
Hospitalization	0.7344 ***	1.6762 ***
	(0.1233)	(0.3811)

Table C.5 *continued*

Variable	Males	Females
Smoker	0.1240 ***	0.7168 ***
	(0.1090)	(0.2721)
Smoker in the past	-0.1294	0.0206
	(0.1136)	(0.3161)
Health insurance	0.0888 ***	0.1136
	(0.1101)	(0.3103)
Life insurance	0.1532 ***	0.2367
	(0.1130)	(0.2945)
Home ownership	-0.0075 ***	-0.6535 ***
	(0.1055)	(0.2517)
University	-0.1517 ***	1.4798 ***
	(0.2019)	(0.4876)
Upper sec. school	0.0073	0.4791
	(0.1545)	(0.4281)
Lowe sec. school	0.3851 ***	0.5259
	(0.1226)	(0.3364)
White collar	0.0652	-0.5123
	(0.1282)	(0.3865)
Primary, industry	0.5125 ***	2.0534 ***
	(0.1300)	(0.4271)
Services	0.5326 ***	1.9367 ***
	(0.1247)	(0.3771)
Employment rate	4.0455 ***	-8.5732 ***
	(0.9346)	(1.6431)
Intercept	-3.6972 ***	2.3169 ***
	(0.6507)	(0.8680)
Entry cross-section	17,725	8,393
Exit cross-section	15,580	7,544

Source: ISTAT, Aspetti della Vita Quotidiana 1993-2003. Note: The baselines are (i) years 1993-1994, (ii) age 60-65, (iii) household size higher than four components, (iv) not having been hospitalized during the last twelve months, (v) non-smoker, (vi) not having subscribed either a health insurance or a life insurance, (vii) non-homeowner, (viii) at most primary school, (ix) blue collar, (x) public administration. ***: p-value $\leq$0.01, **: 0.01 $<$ p-value $\leq$ 0.05, *: 0.05 $<$ p-value $\leq$ 0.1.

Table C.6: Italy, employees aged 50-65 in 1993-2002. With respect to the reference category described below and for the age-intervals 56-57 and 60-65, we report the hazard rate levels over the pairs of years considered and the corresponding relative deviations from the baseline period (1993-1994). Details on the specifications are reported in Table C.5.

Variable	56-57	60-65
Males		
Years 1993-1994	0.1275	0.2341
Years 1995-1996	0.1965	0.3384
	54.13%	44.57%
Years 1997-1998	0.1899	0.3290
	48.97%	40.55%
Years 1999-2000	0.1940	0.3350
	52.23%	43.10%
Years 2001-2002	0.0570	0.1123
	-55.26%	-52.02%
Females		
Years 1993-1994	0.0309	0.2296
Years 1995-1996	0.1264	0.5750
	309.17%	150.42%
Years 1997-1998	0.1802	0.6727
	483.31%	192.97%
Years 1999-2000	0.1232	0.5678
	298.84%	147.30%
Years 2001-2002	0.0253	0.1955
	-17.98%	-14.84%

Baseline category: (i) years 1993-1994, (ii) HH size=3, (iii) no hospitalization during the last twelve months, (v) non-smoker, (vi) no subscription of either health or life insurances, (vii) homeowner, (viii) upper secondary school, (ix) white collar, (x) employed in the primary or industry sector, (xi) average employment rate in 1993-1994.

Table C.7: Italy, individuals aged 50-66 in 1993-2003. Employment rates (%) by gender, year and age-groups.

Year	50-54	55-59	60-66
Males			
1993	79.44	61.28	26.29
1994	78.55	58.11	24.41
1995	77.08	56.99	25.54
1996	76.32	53.28	24.35
1997	74.57	53.46	24.95
1998	78.07	50.55	24.87
1999	78.74	50.06	21.82
2000	82.21	51.62	25.14
2001	81.99	50.23	21.03
2002	85.22	53.77	24.92
2003	85.97	56.60	24.80
Females			
1993	34.27	21.03	5.76
1994	35.52	19.57	6.00
1995	36.10	18.86	6.39
1996	36.07	19.64	5.34
1997	35.87	20.86	5.55
1998	38.72	21.31	5.78
1999	37.60	21.63	5.77
2000	43.52	21.96	5.77
2001	43.96	21.79	6.12
2002	46.55	26.66	6.36
2003	47.25	29.70	7.86

Source: ISTAT, Aspetti della Vita Quotidiana 1993-2003.

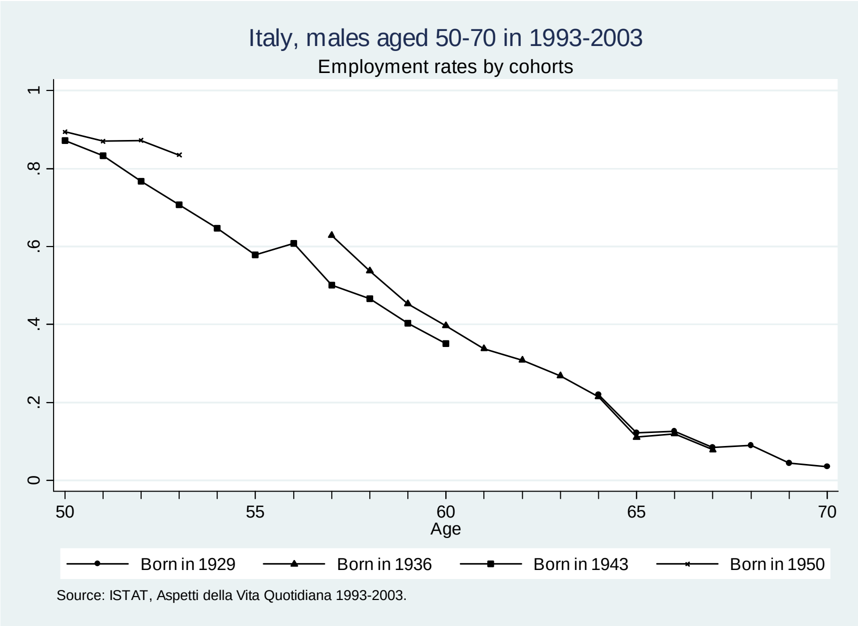

Figure C.1: Employment rates.

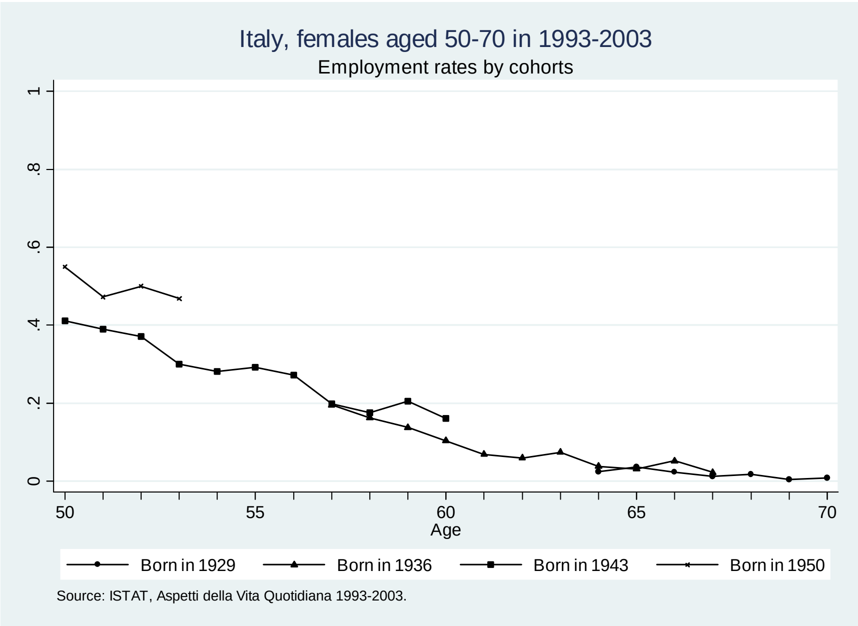

Figure C.2: Employment rates.

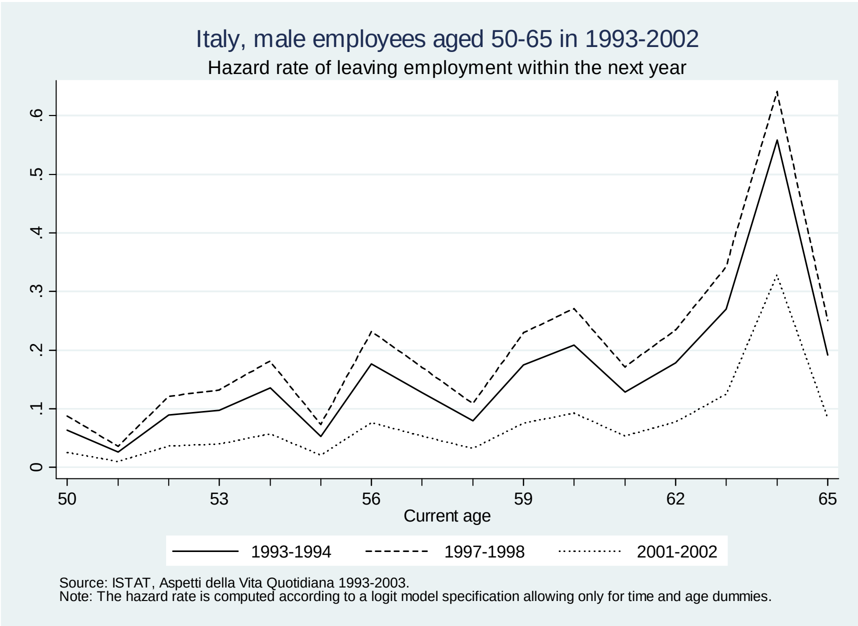

Figure C.3: Age-profile of the hazard rate of exiting employment.

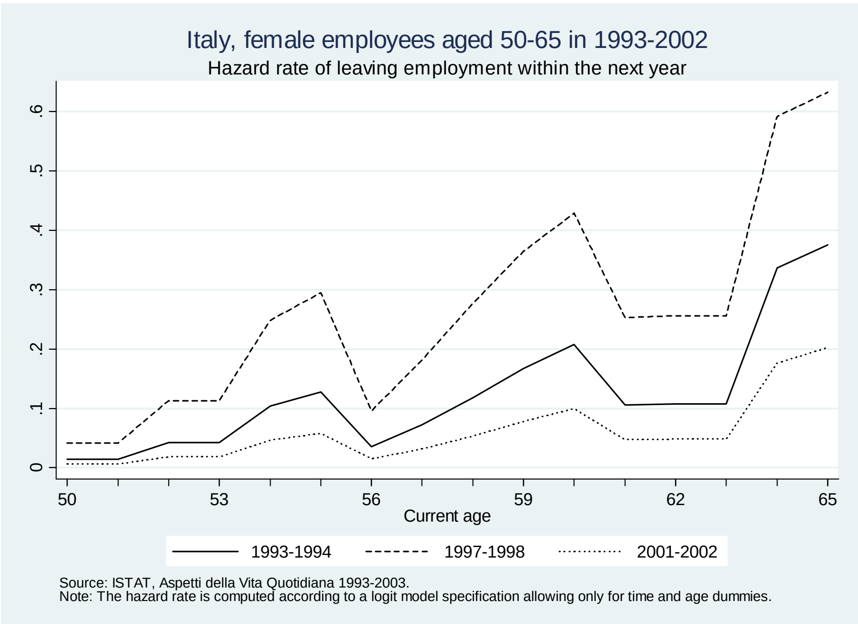

Figure C.4: Age-profile of the hazard rate of exiting employment.

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