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

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May 27, 2008

Abstract

The empirical analysis in this paper adopts logit models to study the hazard rate of ceasing from work by the next year for Italian older employees. The specifications are estimated resorting to the framework proposed by Güell and Hu (2006), which extracts information from repeated cross-sections to recover the hazard rate of interest at the individual level. The sample is drawn from the ISTAT survey Aspetti della Vita Quotidiana and includes employees aged 50-65 in 1993-2002. Our results show that, even conditioning on a wide set of socioeconomic factors, the age profile of the hazard rate is increasing and confirms the low labor market attachment of older workers. Further, the time evolution of the risk of becoming not employed appears to be hump-shaped and achieves its maximum for employees at work in the period 1995-2000, which is characterized by the introduction of important changes in the Social Security system aimed at extending the working life of the elderly.

JEL Codes: C41, J26.

Keywords: Labor supply dynamics, duration analysis, generalized method of moments.

*I wish to thank Raffaele Miniaci, Federico Biagi, Lorenzo Cappellari, Adriano Paggiaro and participants at the WISE Workshop (2008) for useful comments. Financial support from the MIUR and the EU grant COMPARE under research contract no. CIT5-028857 is gratefully acknowledged. The usual disclaimer applies. Corresponding address: Department of Economics and Management, University of Padova, Via del Santo 33, 35123 Padova, Italy. E-mail: danilo.cavapozzi@unipd.it. Telephone: +390498274255.

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.

In Italy 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

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

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 different 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, in the Bank of Italy Survey of Household Income and Wealth (SHIW) less than one half of the households interviewed in each wave belong to the panel section[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 collects in each wave multipurpose information for a sample of approximately 20,000 households representative of

[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.

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 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 men not employed 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 SHIW, the European Community Household

[4] The waves 1993-1999 and 2001-2003 of Aspetti della Vita Quotidiana are released according to the contract no. 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.

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

Panel (ECHP) and the public releases of the INPS archive.

The structure of this paper is the following. Section 2 briefly describes the principal changes in the Social Security system ongoing in the period considered in this analysis. A brief survey of the related literature is provided in Section 3. Section 4 introduces the GMM estimator adopted. Section 5 presents the dataset used to obtain the main results of this work, which are the focus of Section 6. Finally, Section 7 concludes.

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 ageing 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. While 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 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

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

either old-age or seniority pensions and have contributed to the scheme for more than 40 years.

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.

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

[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).

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 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

[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.

[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.

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 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. It is worth remembering that our study complements the contributions listed in this section since it builds upon

[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.

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.

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

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

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

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

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

$$\underbrace{\sum_i x_i \Lambda(x_i'\beta)}_{\substack{\text{all employees}\\ \text{aged 50-65 at time } t}} = \underbrace{\sum_i y_i x_i}_{\substack{\text{only those still in the sample and}\\ \text{employed at time } t+1}}. \tag{2}$$

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

On the one hand, the LHS of (2) 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 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 (2) by the set of moment conditions

$$\underbrace{\sum_i x_{it}\Lambda(x'_{it}\beta)}_{\substack{\text{all employees aged 50-65} \\ \text{in the cross-section of time } t}} = \underbrace{\sum_j x_{jt+1}}_{\substack{\text{all employees aged 51-66} \\ \text{in the cross-section of time } t+1}} . \tag{3}$$

The LHS of (3) 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 unweighted 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), 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 (2), 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) can be motivated going through the law of iterated expectation,

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

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) after a sample size normalization and the same applies to $E[x_{it}\Pr(\mathbf{1}(t+1)=1|x_{it})]$ 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

[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.

quadratic function of the vector

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

More specifically,

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

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

$$\begin{aligned} W = V\left[\frac{1}{\sqrt{n_t}}\sum_j x_{jt+1} - \frac{1}{\sqrt{n_t}}\sum_i x_{it}\Lambda(x'_{it}\beta)\right] = \\ = V\left[\frac{1}{\sqrt{n_t}}\sum_j x_{jt+1}\right] + V\left[\frac{1}{\sqrt{n_t}}\sum_i x_{it}\Lambda(x'_{it}\beta)\right] - \\ -2Cov\left[\frac{1}{\sqrt{n_t}}\sum_j x_{jt+1}, \frac{1}{\sqrt{n_t}}\sum_i x_{it}\Lambda(x'_{it}\beta)\right] = \\ = \frac{n_{t+1}}{n_t}V\left[x_{jt+1}\right] + V\left[x_{it}\Lambda(x'_{it}\beta)\right] - \frac{n^c_{t,t+1}}{n_t}2Cov[x_{jt+1}, x_{it}\Lambda(x'_{it}\beta)]. \end{aligned} \tag{4}$$

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 (4)[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

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

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,

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

where[15]

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

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 where 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 paper the outcome of interest will be $1-\Lambda(x'_{it}\beta)$, 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, since 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

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

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

5 Data

The repeated cross-sections called Aspetti della Vita Quotidiana have 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, employment, leisure, health and utilization of medical care services[16].

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 (1) reports the results for males and suggests that between

[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).

ages 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 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 cohorts[18] 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 (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 cohort-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 cohorts 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 have been 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 (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 made up of the same individuals, as long as they furnish valid estimates for

[18] The impossibility of observing the employment rates experienced by each cohort in the whole age-range considered prevents us from fully disentangling between age and cohort effects.

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 the 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 $t+1$.

6 Results

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 (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[19].

Table (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 (3) and (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 in Table (2).

The related literature agrees on pointing out peaks in the probability of quitting employment around ages 55, 60 and 65[20]. Although the evidence summarized in Figures (3) and (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[21]. The gradual modifications in the requirements for old-age and seniority pensions 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.

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

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

[21] Exits at the age of 57 are in line with the age-requirements for the seniority pensions described in Section 2.

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 (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 has actually begun 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 (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 (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 (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[22] 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 (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

[22] As written in Section 2, in 1992 old-age pensions were granted to women aged at least 55.

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 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[23], 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.

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 from the rest 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.

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. These dynamics can be driven by

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

whatever uncontrolled time-varying factor, such as cohort effects and 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[24]. 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[25]. The results summarized in Table (5) are in line with those discussed

[24] 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.

[25] We are forced to drop the set of dummies for the region of residence in view of their strong

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 (6) overall confirm the considerations provided in the previous section.

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 (7) shows that in 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.

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 below the levels settled by EU, especially when we look at the employment outcomes of the elderly. As most 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

relationship with the employment rates.

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 sample of about 60,000 individuals randomly selected from the whole Italian population. 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 Italian institutional background is shown 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 longer working lives. 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.

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Table 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 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 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 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 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 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 ≤ 0.01, **: $0.01 <$ p-value ≤ 0.05, *: $0.05 <$ p-value ≤ 0.1.

Table 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 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 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.

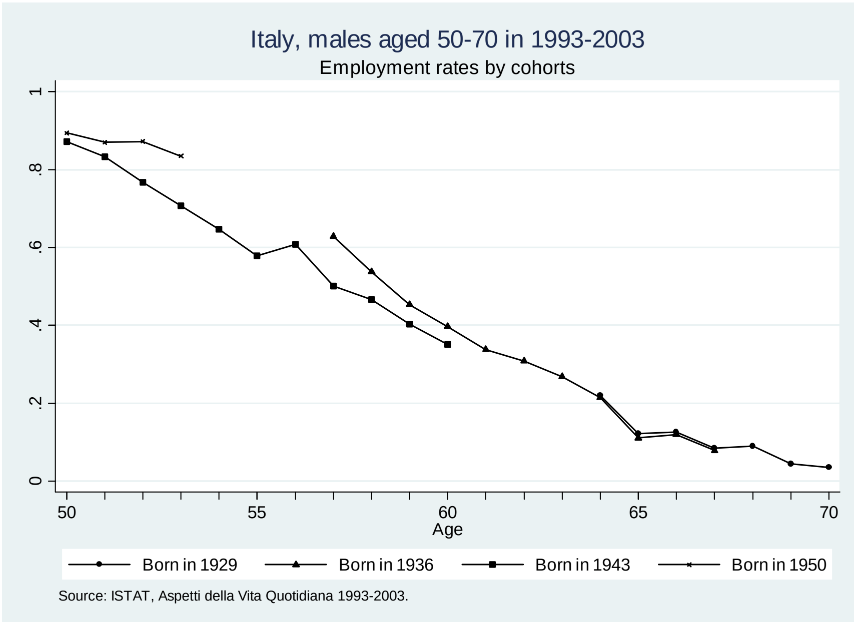

Figure 1: Employment rates.

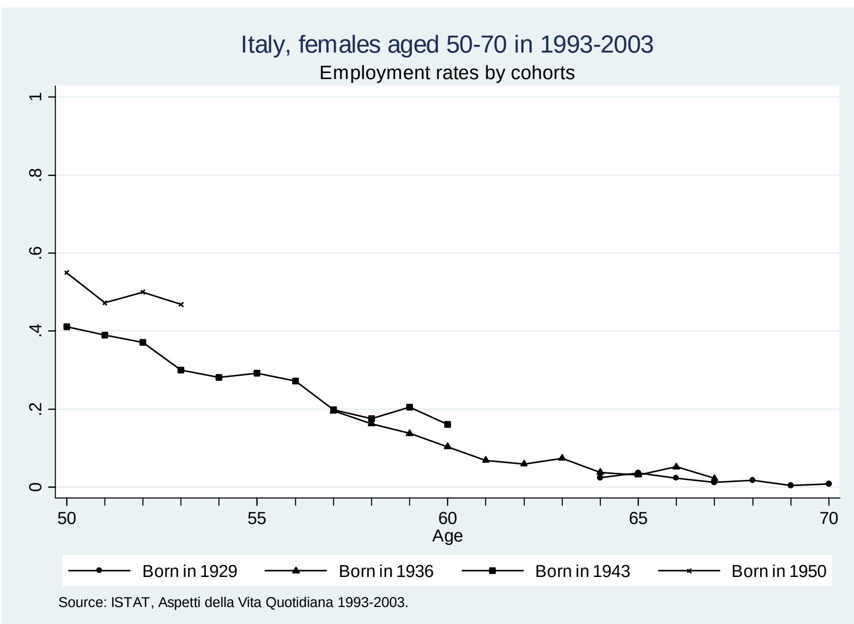

Figure 2: Employment rates.

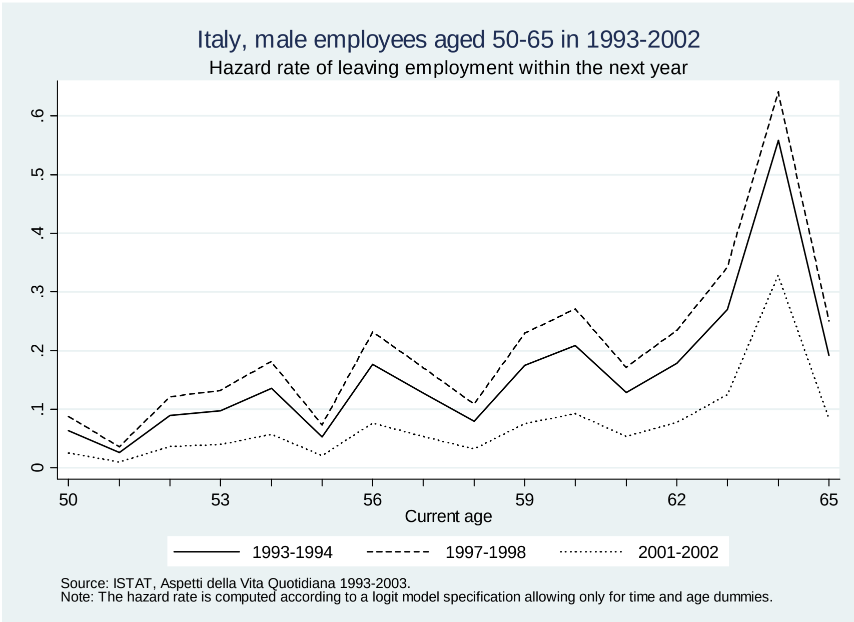

Figure 3: Age-profile of the hazard rate of exiting employment.

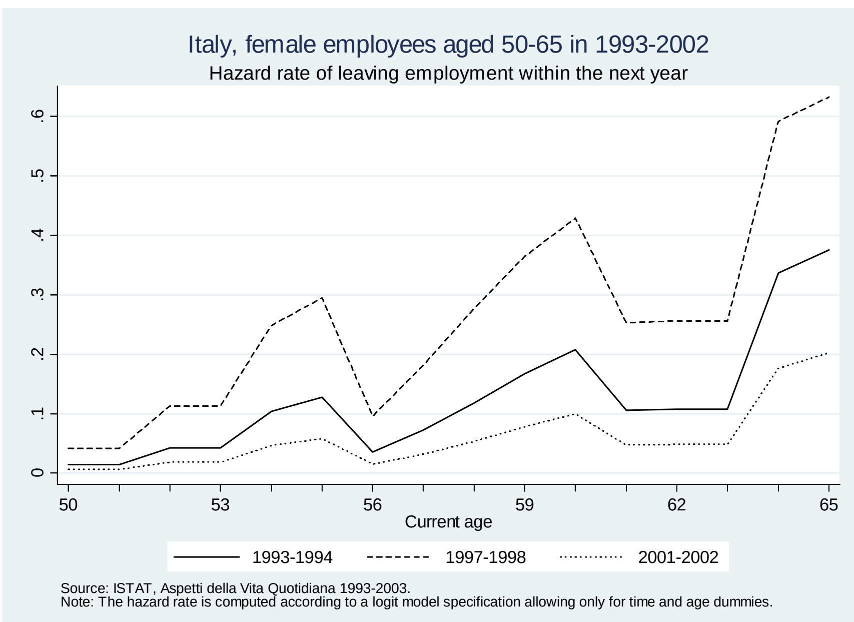

Figure 4: Age-profile of the hazard rate of exiting employment.