DISCUSSION PAPER SERIES

No. 4394

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PUBLIC POLICY

Centre for Economic Policy Research

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Available online at: **www.cepr.org/pubs/dps/DP4394.asp**

ISSN 0265-8003

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Discussion Paper No. 4394
May 2004

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CEPR Discussion Paper No. 4394

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ABSTRACT

What Triggers Early Retirement? Results from Swiss Pension Funds

Early retirement is predominantly considered as the result of incentives set by social security and the tax system. But people seem to retire early even in the absence of such distortions, as the Swiss example demonstrates. We look for determinants of early retirement, in particular the role of lifetime income and family status, using individual data from a selection of Swiss pension funds. Our findings suggest that affordability is a key determinant in retirement decisions: more affluent men and – to a much smaller extent – women tend to leave the work force earlier. The fact that early retirement has become much more prevalent in the last 15 years is another indicator for the importance of affordability as Switzerland's funded pension system has matured over that period leading to higher effective replacement rates. We also find sizeable differences in retirement behaviour across marital status. These may be explained by a constrained rational choice based on differential mortality and the desire of couples to coordinate their entry into retirement.

JEL Classification: D91 and H55
Keywords: duration models, occupational pension and retirement decision

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Submitted 13 April 2004

1 Introduction

Early retirement is a widespread phenomenon throughout Europe causing financial distress to almost all public pension systems. In most countries the main reason for this effect seems clear: High replacement ratios and high implicit tax rates on working beyond a certain age induce workers to opt for an early exit out of the labor market. But early retirement is also prevalent — albeit to a lesser degree — in Switzerland, where implicit tax rates on working on in old age are virtually zero. As of today the public (first pillar) pension system does not even offer early retirement plans. Like in many other countries, the retirement age has fallen in the last decade despite the fact that institutional incentives (statutory retirement age, pension accrual rate, replacement rate) have stayed basically unchanged.

The surge of early retirement in Europe may be explained by several factors. First, gross pension replacement rates have increased in most OECD countries since the early 1960s. Workers are more likely to withdraw from the labor market as soon as they have reached pensionable age if benefits are close to wages. Second, pension accrual rates at older ages have fallen (and differ significantly across OECD countries).[1] If the pension accrual rate is zero there are no penalties from withdrawing from the labor market, whereas if it is high there are incentives for workers to continue working. The implicit tax on continued work, as proposed by Gruber and Wise (1997), has the same impact.[2] A third reason for the increase in early retirement are changes in related benefit systems, notably unemployment and disability benefits. Though not originally intended to support people in retirement, changes in eligibility conditions have *de facto* turned these schemes into early retirement programs in a number of OECD countries.

Although not offered by the first retirement pillar, many (mandatory) Swiss occupational pension plans avail early retirement schemes. These range from simply offering an option for early withdrawal from employment at actuarially fair reductions in the pension benefits to generous early retirement plans including additional payments to make up for first pillar benefits up to the legal retirement age. The observed retirement age in such plans is substantially below the statutory age on average, but varies widely within and between different pension funds.

[1]The pension accrual rate is defined as the gains in old-age pensions from working for an additional period.

[2]The implicit tax (or subsidy) on continued work is the average annual variation in the social security wealth relative to gross earnings obtained by postponing retirement. The social security wealth is the sum of the discounted value of expected benefits (either pensions or other non-employment benefits) minus the discounted cost of obtaining these benefits.

Why do people retire early even when the benefits are adjusted to the retirement age in an actuarially fair way, and when the marginal implicit tax rate on working is unimportant? Economic theory predicts that workers choose their intertemporal consumption and labor supply optimally according to a utility function and with respect to a lifetime budget constraint. If the adjustment for early retirement were the same for everybody, in theory richer individuals should retire later due to their higher life expectancy. Financial constraints of low income workers, on the other hand, may lead to the opposite outcome in which poorer individuals are forced to work longer (affordability). Individuals may thus retire because they can afford to do so, or because they are unable to work any longer because of bad health or the lack of opportunities. Which of the two applies is an empirical question in the end.

In the present paper, we especially focus on the role of lifetime income on the retirement decision. Due to the fact that the second pillar is mandatory (accumulated pension capital has to be transferred to the new plan in case of job changes), accumulated capital at retirement is an excellent measure for lifetime income. We find that affordability is a key determinant in retirement decisions. Wealthier men tend to leave the work force earlier, at least up to a relatively high average lifetime income. This may lead to a socially undesired drain of human capital among elderly workers. Low income workers, on the other hand, often work up to the legal retirement age even in pension funds in which early retirement packages are generous. In these cases the need to generate income seems to be the only explanation for working up to the statutory retirement year. Due to differences in mortality rates across income groups, richer individuals thus tend to enjoy a much longer retirement spell than poorer people. For the pension funds, this means that adverse selection effects are unimportant. Despite data limitations, we find that marital status is another key determinant for retirement decisions. Financial needs and joint retirement problems seem to be the dominating forces.

We also find that the tendency to retire early has increased considerably in the last 15 years. This finding may also be explained by affordability. Due to a maturing of Switzerland's second pillar, more people are now able to accumulate sufficient funds to pay for an early labor market exit than one or two decades ago.

Our findings suggest that the reason for early retirement does not solely lie in the incentive structure implied by public pension plans. The preference for leisure in old age seems to be a dominating driving force for leaving employment. Many poorer individuals only keep working because they cannot afford to retire.

2 The Swiss social security system

Switzerland's pension system is composed of three pillars, of which the first and second are of approximately equal importance.[3] The first pillar AHV/AVS[4] is a predominantly pay-as-you-go (PAYG) system and aims at providing a basic subsistence level of income to all retired residents in Switzerland. The second pillar, the so-called BVG/LPP[5] is a mandatory, employer-based, fully funded occupational pension scheme. Gross replacement ratios in Switzerland increased from 28.4% in 1961 to 49.3 % in 1995. Since 1990, however, these numbers have basically stayed unchanged. The statutory retirement age is 65 for men and currently 62 for women, the latter will be increased gradually to 65 in the next few years. Note that retirement at 65/62 is not mandatory by law, but reaching age 65 for men or age 62 for women is rather an eligibility condition for claiming public pension benefits. Many labor contracts, however, specify a retirement age coinciding with the eligibility age.

In 2000, on average, approximately 50% and 40% of publicly provided transfer retirement income were paid out by the first and second pillar, respectively. This understates the importance of the occupational pension system, however, as contributing agents today can expect more than half of their combined first and second pillar income to come from the second. The second pillar's main goal is to maintain the pre-retirement living standard together with the benefits stemming from the first pillar. Upon attainment of retirement age, the accumulated capital can be withdrawn either as a monthly life-long annuity or as a lump sum (or a mix of the two) provided the pension fund allows for the lump sum option (which is usually the case in defined contribution plans).

Occupational pension benefits are strictly proportional to the accumulated retirement assets (retirement credits plus accrued interest). The accumulated capital K is translated into a yearly pension B using the conversion factor γ, $B = \gamma K$. This conversion also applies to defined benefit plans; the fund has to make sure that enough capital is accumulated to cover the claims made based on previous income.

The BVG/LPP offers joint annuities for men, but not for women. The conversion factor is the same for everybody irrespective of gender, family status or

[3]A detailed description of all aspects of the Swiss social security system is, however, beyond the scope of this paper. The interested reader is referred to Queissar & Vittas (2000, especially concerning institutional details) and Bütler (2004, for the second pillar).

[4]AHV = Alters- und Hinterbliebenen-Versicherung; AVS = Assurance Vieillesse et Survivants.

[5]BVG = Bundesgesetz über die berufliche Alters-, Hinterlassenen- und Invalidenvorsorge, LPP = Loi fédérale sur la prévoyance professionnelle vieillesse, survivants et invalidité.

income. Children under age 18 (or under age 25 if still dependent) of retired persons get an additional pension of 20% of the main claimant's benefit. The widow of a retired receives a benefit amounting to 60% of the previous pension, his dependent children a benefit of 20% each. Survivor benefits lead to sizeable differences in the money worth's ratios with respect to marital status and gender.

Apart from retirement income, the second pillar also provides insurance for disability as well as survivors of insured active men (but not women) during the accumulation period. The availability of disability insurance, though not originally intended to provide retirement income, might have contributed to the observed tendency to retire earlier. This is most certainly also the case for Switzerland, albeit to a lesser degree than in other countries. Privately organized plans are far more reluctant to grant an early exit from the labor force based on disability grounds than public programs (very often potential claimants have to see a company related medical doctor to confirm the disability).

3 Related literature

Kotlikoff (1979) shows that the provision of social security will not affect the retirement decision under the assumption of perfect capital markets, actuarial fairness and known lifespan, as pensions are equivalent to private savings. Crawford and Lilien (1981) relax each assumption in turn, and show that the effect on the date of retirement is in general ambiguous, but that a progressive system tends to advance retirement for low-income workers. Social security also has an impact upon the labor supply decision and on the allocation of labor and consumption over the life cycle. Craig and Batina (1991) simulate the strengths of such effects. Their results show how the introduction of a social security program acts as a disincentive to supply labor in the later stages of life, thus affecting also the level of output produced and the capital-labor ratio.

In theory the social security system thus is a key variable in explaining the retirement decision of older workers. The quantitative effect of old age insurance on retirement has been measured using one of the following three approaches:

1. The "lifetime budget constraint" approach (Burtless and Hausman, 1978; Hausman and Wise, 1980; Venti and Wise, 1984; Burtless, 1986), in which individuals face discontinuous or kinked budget constraints: The lifetime budget constraint is similar to the standard labor-leisure budget constraint, with annual working hours replaced by years of labor force participation, and annual earnings replaced by cumulative lifetime compensation. The optimal age of retirement is determined by a utility function defined over

years of work and cumulative compensation. This approach assumes that individuals know with certainty the opportunities that will be available to them in the distant future.

2. The "option value" approach (Lazear and Moore, 1988; Stock and Wise, 1990), whose central feature is the option value of continued work: The idea is that if one retires before the early retirement age, the option of a later bonus provided by pension plan provisions is lost. Continuing to work preserves the option of retiring later.

3. The "hazard model" approach (Diamond and Hausman, 1984; Hausman and Wise, 1985), which is a reduced-form technique designed to capture the effects on retirement of movements in variables such as Social Security wealth. Implementations of this model are not perfectly "forward looking" as the nonlinear budget constraint specifications, but allow for the updating of information as individuals age. Moreover, these models allow to consider not only pure economic variables (wages, pension benefits, ...) but also other non pecuniary factors, such as the health status, family circumstances and eligibility for different retirement schemes.

We will use the hazard model approach where the retirement decision is treated as a dynamic discrete choice. Similar techniques have been used in other empirical studies, such as Miniacci (1998) for Italy, Antolin & Scarpetta (1998) for Germany and Mastrogiacomo, Alessie & Lindeboom (2002) for the Netherlands.

4 Duration models

Survival-time data documents spans of time ending in an event, called "failure". In our case, the failure event is entering retirement. Let T be a random variable which describes the duration of the working period (i.e. the period before retirement)[6]. We assume T has a continuous probability distribution $f(t)$, where t is a realization of T. The cumulative distribution function is

$$F(t) = \int_0^t f(s)ds = \Pr(T \leq t). \quad (1)$$

The survival function, i.e., the probability that the spell of the working period is of length at least t, takes the form

$$S(t) = 1 - F(t) = \Pr(T \geq t) \quad (2)$$

[6]In our case, the duration is the time period between the age of 55 and the age at retirement. As all individuals have the same time origin, our time axis does not correspond to the "real" time axis.

The hazard rate (or age-specific failure rate) is the rate at which spells are completed after duration t, given that they last at least until t. In our specific case, the hazard function is the probability of entering retirement at a certain age t, conditional on the fact that the agent has not retired before that age. Let $h(t)$ denote the hazard function, so that:

$$h(t) = \lim_{dt \longrightarrow 0^+} \frac{\Pr(t \leq T < t + dt | T \geq t)}{dt} = \lim_{dt \longrightarrow 0^+} \frac{F(t + dt) - F(t)}{dt S(t)} = \frac{f(t)}{S(t)} \quad (3)$$

Duration models can be distinguished between non-parametric, semi-parametric and parametric models, on the basis of whether they predict the probability distribution of a certain event by means of a set of individual characteristics as explanatory variables, in this context called covariates. In this section we first review the Kaplan-Meier (or product-limit) estimator as an example of a non-parametric model that is estimated without covariates. If applied to subgroups of the data set, it is a useful tool to visualize differences across groups and to assess the validity of assumptions used for parametric and semi-parametric survival time models.

We then consider the case of hazard rates that depend on a set of regressors X and parameters β to be estimated. In the present paper we choose the semi-parametric case and describe the Cox proportional hazard model[7].

4.1 Non-parametric duration models: The Kaplan-Meier (or product-limit) estimator

We assume that the hazard function's distribution is discrete, with atoms f_j at finitely many specified points $a_1 < a_2 < ... < a_n$. The survivor function $S(t)$ may be written in terms of the discrete hazard function h_j as

$$S(t) = \prod_{a_j < t} (1 - h_j) \quad (4)$$

The f_j can be expressed in terms of the h_j as

[7]In the parametric case, the density function takes a specific form, such as Weibull, log-normal, log-logistic or Gompertz distribution. The form of the survival function and hazard rate follows. The parametric regression allows to add covariates as in the semi-parametric case. Note that we also carried out parametric regressions with the same covariates. As the results were very similar to the semi-parametric case, we only display the latter.

$$\begin{aligned}
f_1 &= h_1 \\
f_2 &= (1-h_1)h_2 \\
&\dots \\
f_j &= (1-h_1)(1-h_2)\dots(1-h_{j-1})h_j \\
&\dots \\
f_n &= (1-h_1)(1-h_2)\dots(1-h_{n-1})h_n
\end{aligned} \tag{5}$$

The constraints $f_j \geq 0, \sum f_j \leq 1$ become $0 \leq h_j \leq 1$.

Let r_j be the number of individuals in view at a_j, and d_j the number of individuals who fail at a_j. The likelihood function for n independent binomials, with respectively r_j trials, d_j failures and probability of failure h_j, is then

$$L = \prod_j (h_j)^{d_j}(1-h_j)^{r_j-d_j} \tag{6}$$

and the corresponding log-likelihood function is

$$\log L = \sum_j [d_j \log h_j + (r_j - d_j)\log(1-h_j)] \tag{7}$$

The non-parametric maximum-likelihood estimator $\widehat{h}_j$ is the solution of

$$\frac{\partial \log L}{\partial h_j} = \frac{d_j}{h_j} - \frac{r_j - d_j}{1-h_j} = 0 \qquad \text{that is} \qquad \widehat{h}_j = \frac{d_j}{r_j}. \tag{8}$$

The corresponding non-parametric estimator of the survivor function (called Kaplan-Meier, or product-limit, estimator) is then

$$\widehat{S(t)} = \prod_{a_j<t}(1-\widehat{h}_j) = \prod_{a_j<t}(1-\frac{d_j}{r_j}).$$

4.2 Semi-parametric survival-time models: The Cox proportional hazards model

Semi-parametric models enable to derive the relationship between the hazard rate and the explanatory variables without imposing any predefined functional form for the density of the survival time T and, consequently, on the shape of the hazard rate itself. A very common semi-parametric survival-time model is the Cox proportional hazard model.[8] The model does not assume a specific probability

[8] Hausman and Han (1990) and Meyer (1988) propose other semi-parametric specifications for hazard models.

distribution for the time until an event occurs. The effect of the covariates is captured by multiplying the hazard with a term that does not depend on time. The hazard rate for an individual i is modeled as follows:

$$h(x_i, t) = h_0(t)\Psi(x_i, \beta), \tag{9}$$

where x_i is a vector of explanatory variables and $h_0(t)$ is the "baseline" hazard for an individual under standard conditions ($x = 0$), requiring $\Psi(0) = 1$. The baseline hazard varies with time, but not across individuals. It is not specified in terms of a parametric distribution but can be empirically derived from the data. The function $\Psi(x_i, \beta)$ specifies how the covariates are affecting the hazard rate. Its most convenient form is $\Psi(x_i, \beta) = \exp(x_i'\beta)$.

To estimate the parameters of the Cox proportional hazard model, one uses a *partial* likelihood function, i.e., the likelihood without the baseline hazard. Cox's partial likelihood estimator provides a method of estimating β without requiring estimation of h_0. The variance of $\widehat{\beta}$ is estimated by the method of Lin and Wei (1989).

We assume that there are D observed failures from the sample of size N. Let j index the ordered failure times t_j $(j = 1, ..., D)$. D_j is the set of the d_j observations that fail at t_j and R_j is the set of observations k that are at risk at time t_j, i.e., all observations that have not failed up to t_j. The partial log-likelihood function can then be written as

$$\ln L = \sum_{j=1}^{D}\{\sum_{k\in D_j} x_k\beta - d_j \ln[\sum_{i\in R_j} \exp(x_i\beta)]\}. \tag{10}$$

This form of the likelihood is also valid in cases, in which not all survival times are known.[9] Most often one knows the minimum span a person has survived, but not the realized survival time which is longer than the minimum registered. These observations (in the notation used above, there are $N - D$) are called "censored" observation times.

5 Data and empirical strategy

5.1 The data

In the empirical analysis we use data collected at the individual level from 15 Swiss companies, both public and private, active in several industrial branches.

[9]This feature of survival data is especially important in the medical sciences, as people might still be alive when the study closes, or cannot be tracked anymore.

They include the national public railway company, civil servants in two cantons, several industry firms, as well as clothing and food firms. Due to lack of sufficient information, we drop all observations with retirement year earlier than 1990.

The dataset consists of 8452 observations[10] (9441 before dropping pre-1990 retirement, and observations without marital status). We have information about date (or year) of birth, marital status, date (or year) of retirement, yearly pension payments (base level) and yearly additional pensions for children and for first pillar replacement packages. On the firm level, we are also provided with details of early retirement plans, in particular the adjustment in the conversion factor and the availability of first pillar replacement packages.

Some firms also provide us with information about the number of children under 18/25[11], the amount withdrawn as a lump sum (if this option is available), the total capital accumulated at retirement, and an indicator whether the individual has chosen a non-standard retirement option.

As reported in Table 1, males and females represent 63.5 and 36.5 percent of the sample, respectively. The distribution of marital status is very different for men and women, the great majority of men is married (85.4%) at retirement, whereas almost half of the retiring women live alone (52.6% only are married). There are also large differences in annuity across gender and marital status, with women getting approximately half the amount of men on average. The only exception are singles, for which females fare better. This can be explained by the fact that single women are more likely to be well educated than average women, whereas the contrary is the case for men.

The sample consists of individuals whose age at retirement ranges from 55 to 70. We explicitly exclude all observations for which the path to retirement passes via a period of disability benefits. Despite a difference of 3 years between men and women in the statutory retirement age for individuals in the sample, the difference in the factual retirement age is less than half of this number. The median or average retirement age does not seem to vary very much across marital status either. There have been, however, important changes in retirement behavior over the last 15 years. Figure 1 depicts the distributions of the age at retirement for men and women for three different subperiods (1990-1994, 1995-1999, 2000-2003).

[10]The cleaning and editing of the data has been a considerable task. Firstly, the data format provided varied widely across companies. Secondly, much of the relevant information for the project had to be imputed from other sources (regulation of pension fund) or from a combination of available data. In many cases the information could only be gathered from a personal interview with the responsible pension fund manager.

[11]Children under age 18 are always eligible for additional benefits. For those over 18, but under 25, a pension is available for disabled children and those still in school.

The distribution of age at retirement has a peak at the respective (current) statutory/eligibility retirement age of 65 (men) and 62 (women). For the second time period the profile for men has another peak around age 62, which is the age at which some pension funds offer early retirement benefits — sometimes even full — even for men. This peak becomes the most prominent one in the third period. We also notice another peak at age 60. This is often the lowest age for which early retirement packages are offered at relatively good conditions. It is interesting to note that a sizeable fraction of women work beyond the statutory retirement age, though this number has clearly decreased over time. The most striking feature of these distributions is a clear shift of the retirement decision to lower ages for both men and women. This decrease is particularly strong from 1995-1999 to 2000-2003.[12]

It is important to mention that the fraction of people retiring early within the included pension plans far exceeds the corresponding fraction for the whole population. For the companies in our dataset 79% of men and 62% of women retire before the statutory retirement age in the 2002-2003, whereas the corresponding numbers for the whole of Switzerland are 55% (men) and 44% (women) in 2002.[13]

insert Figure 1 & Table 1 here

5.2 The empirical strategy

A first task is to construct a measure of second pillar income that is equivalent across companies. This is basically equivalent to constructing a measure for accumulated capital at retirement plus adding the present value of additional benefits to be received by the pensioners. For this purpose, we use firm specific information on conversion factors, early retirement plans and other benefits. Thus the variable "annuity" corresponds to the yearly pension if all capital were fully annuitized, including the regular yearly pension plus any temporary payments, as well as the annuitized value of any lump sum payment upon retirement. To account for economic growth and inflation, these numbers are deflated by the nominal Swiss GDP (base year 2002). For our empirical analysis we use the log (variable "ln(annuity)") as well as its square (variable "ln(annuity)^2") to capture

[12] The median (mean) retirement age for men is 65 (63.2), 63 (62.7), and 62 (61.7) for the periods 1990-1994, 1995-1999, and 2000-2003, respectively. The corresponding numbers for women are 62 (61.5), 62 (61.3), and 60.1 (60.2), respectively.

[13] Recall that, in general, low income people (and to some extent self-employed) are not covered by second pillar pension plans. This might be a first indicator that individuals who retire early do so because they can afford it.

potential nonlinear effects. Recall that, due to the legal requirement transfer pension capital from a previous to the current employer, second pillar capital or income is a very good proxy for lifetime income. Nonetheless, individual data on retirement wealth cannot convey an exact picture of a person's wealth position as the latter depends on additional income and wealth by the spouse, especially for women.

As is obvious from Figure 1, retirement behavior is very different for men and women. For individuals retiring before 2004, the eligibility age for old age benefits as well as many conditions within company pension plans (notably early retirement conditions) are also very different across gender. We thus analyze men and women separately.

Time is bound to play an important role despite the fact that the proxy for average life-time income has been deflated. Firstly, the effective replacement rate has increased due to a maturation of the system in most companies. This effect is captured by a linear time trend for the annuity variables in all estimations. Secondly, changes to company pension plans may have influenced people's decision to retire or not. This latter effect will be captured by dummy variables by retirement year (we also report a linear retirement effect in two base estimations).

As pension plans differ considerably across pension funds, we always include company fixed effects. For the largest companies in the sample, estimations are reported on the firm level.

We have too few observations on the number of children to make meaningful inference about this parameter. The same is true — to a lesser extent — for the fraction of accumulated capital withdrawn as a lump sum, and whether the person has deviated from the standard option.[14] Perhaps surprisingly, using firm level estimations we find no impact of the latter two variables on the retirement age decision. We do, as a consequence, not report the coefficients of the two variables.

At first sight, all retirement ages are observable, i.e., there is no obvious censoring in the data. However, although not required by law, many companies force people by contractual agreements to retire at the age eligible for first pillar benefits at the latest. A late or early retirement presumably is the result of the interaction of several reasons and options, whereas a retirement at the statutory age is rather an automatic act without further careful considerations. This means that we observe the eligibility age in such cases, although the person might have

[14]We consider a person deviating from the standard option if (s)he chooses another mix of annuity and lump sum payments upon retirement than the standard offer of the company (in most cases a pure annuity). In many cases, individuals have to declare a non-standard pay-out option several months (and up to three years) in advance.

chosen to work longer had she been free to do so. A visual inspection of the histograms in Figure 1, with obvious peaks at 65 (men) and 62 (women) seems to support the incidence of an important bias at ages 65 and 62 for men and women, respectively. We therefore choose to mark all observations with retirement ages around the eligibility age as censored, i.e., we treat them as if we did not know the reason why these individuals have retired at that age. We have experimented with various intervals around the eligibility age, finding very small differences in estimation outcomes. Results are reported for a censoring interval of "eligibility age ± 3 months". As a robustness check, we also present estimations with all data points marked uncensored.

To classify the different estimations with respect to censoring and the impact of the retirement year, the following notation has been chosen:

I	=	no censoring, linear time trend
II	=	no censoring, retirement year dummies
III	=	with censoring, linear time trend
IV	=	with censoring, retirement year dummies

In parenthesis, we add the gender (m = men, f = women), as well as the number of the company or the retirement year if applicable.

6 Empirical results

The following sections report the results of the empirical analysis carried out with the described Swiss data set. Firstly, we present a more descriptive analysis demonstrating the impact of several factors on the retirement decision. We then carry out more formal tests.

6.1 Non-parametric estimation results

We have computed Kaplan-Meier survival function estimates for different subsets of the data (always by gender). The empirical survival functions — only reported without censoring[15] — are shown in Figures 2-4. As mentioned before, Kaplan-Meier survival estimates show the probability of not retiring up to a certain age.

Figure 2 depicts the corresponding estimates for the three time periods 1990-1994, 1995-1999, 2000-2003. In line with Figure 1, we observe a clear downward

[15] The results do not differ very much if censoring is taken into account. The only difference is around the eligibility age of 62 and 65 years for women and men, respectively. As these observations are considered as censored, we do not observe a downward jump in the survival probability at this point, but rather at the end of the censoring interval.

shift in the survival function for both women and men. The huge downward jumps at 62 for women and 65 for men, respectively, are replaced by many smaller jumps over all concerned ages. This reconfirms the observation of a more flexible entry into retirement. Another striking observation is that retirement ages are not equally spaced, but are rather concentrated at full years. This is not surprising given the fact that adjustment rates for early retirement are usually not adjusted in a continuous fashion, but rather in discrete intervals of one year.

To explore the impact of marital status we have split the data along that dimension. Figure 3 shows the results for individuals retiring between 2000 and 2003.[16] For both women and men, the probability of still working after age 55 is lowest for single individuals. Note, however, that single females are also the "richest" women in the sample, while single men have the lowest average annuity of all male retirees. So interpreting the figures without disentangling the effects of marital status and income is delicate. Married men tend to stay in the labor force longer, while married women show a similar exit pattern like single women. Divorced or separated women as well as widows tend to work longer.

Figure 4 shows the estimated survival function by retirement income quartile, again for the period 2000 to 2003. For both men and women, the lowest retirement income quartile tends to stay longest in the work force, at least until the statutory retirement age.[17] The retirement behavior as a function of income is monotonic for women, but clearly not for men. Men in the middle income range tend to retire earlier than both richer and poorer men. It seems as if income played a larger role for the retirement decision of women than for men. However, retirement income is also very much correlated with the family status for women, but far less for men. It is thus important to control for marital status to assess the impact of income.

insert Figures 2, 3 & 4 here

6.2 Cox proportional hazard estimation results

The Cox proportional hazard model does not assume a specific probability distribution for the time until an event occurs. It assumes that the hazard functions of any two individuals are proportional over time, even if the values of one or

[16]Estimates of other periods look similar (not reported here). It is important to do the analysis by period as different aspects, notably changes in the distribution of the marital status over time, may interact and influence the results.

[17]It is worth mentioning again, that second pillar retirement income is roughly proportional to lifetime income above a certain income level in Switzerland. The term "income" thus stands for both retirement income and average lifetime labor income.

more covariates are different. As this is the main assumption of the model, we need to check the proportional hazard assumption for all covariates we want to use.[18] We performed plots of $(-\ln(-\ln(\text{survival})))$curves for each category of a covariate versus (ln(analysis time)).[19] The proportional hazards assumption is not violated if the curves are parallel. In our case we observe that the curves are approximately parallel except at the statutory retirement age of 62 for women and 65 for men. As mentioned before, this is not surprising given the fact that contractual agreements often force people to retire at this age. Even if one considers retirement ages around the eligibility age as censored, the proportional hazard estimation is still often violated for data beyond the statutory retirement age. We believe, however, that these observations should still be included as they convey important information about retirement behavior. Estimations carried out with a truncated data set do not differ very much from the complete data set. From a visual inspection of Kaplan-Meier estimates by income (Figure 4), hazard rates seem to depend on log(income) in a non-monotonic way. We, therefore, also include its square in all preliminary regressions, but only report it if its inclusion leads to a better fit of the model.

Tables 2 to 6 summarize the estimation results for various specifications and for women and men, respectively. The results are displayed as hazard ratios. A hazard ratio greater than 1 means that a unit increase in the covariate increases the hazard rate by *(hazard ratio - 1)* $\times$ 100% . If it is smaller than 1, a unit increase cuts the hazard rate to *(1-hazard ratio)*$\times$ 100%. Estimated coefficients for retirement dummies are not reported in the tables, but are summarized in Figure 5. The number of stars (*) for retirement year dummies in the tables indicate the level of significance for a majority of the estimated hazard coefficients: 10%, 5%, and 1% levels of significance for a majority of coefficients are marked with (*), (**), and (***), respectively. As the year 2003 is taken as a base year, estimates of relative hazard rates below 1 indicate that the retirement probability has increased over the years.

Table 2 presents the results for women. Retirement year is significant at 1% in all regressions (if included in a linear fashion, i.e., as the number of years prior to the base year 2003) and has a hazard ratio smaller than one. This

[18]We use the scaled adjustment for the Schoenfeld residuals and test the null hypothesis that the slope is equal to zero for each covariate in the model. We also perform a global test proposed by Grambsch and Therneau (1994). Note that the test of zero slope is equivalent to testing that the log hazard ratio function is constant over time.

[19]They are often referred to as "log-log" plots. We work with quartiles for the variable "ln(annuity)" and, for the variable "retirement year", with the 3 time periods defined earlier (1990-1994, 1995-1999, 2000-2003).

means that early retirement has become more prevalent over the last decade, which is also confirmed by the estimated coefficients on retirement year dummies (Figure 5). Earlier retirement may have been caused by an improved flexibility in occupational pension plans or the maturing of the system (enabling more women to withdraw earlier from the labor force).

Married women tend to have a higher exit rate than both singles and widows. This result may be explained by two factors. The first is a joint retirement decision of married couples. As wives are younger on average than husbands,[20] they may also be willing to leave the workforce at an earlier age to coordinate the passage into retirement with their spouse. The second reason is that married women are "hedged" by their husbands' income and may thus have lower financial needs than other women. Divorcees, on the other hand, have a significantly lower retirement hazard even if one controls for income.

Retirement income is significant in regressions with censored observations (with a corresponding hazard ratio greater than one). A higher annuity thus induces later retirement, the dependency being linear in logs (we could not identify any non-linearity). Well paid women retire earlier than women with low labor incomes, even if one controls for marital status. This means that the attractiveness of the job does not seem to play a role, but rather the fact that a high pre- and after retirement income makes an earlier retirement age affordable.

The corresponding results for men are summarized in Table 3. All covariates are highly significant in all regressions. The impact of the retirement year is exactly the same as for women: The later we are in the time period, the earlier is retirement on average. However, the role of the marital status is completely different. Married, widowed and divorced men tend to retire *later* than single men. There are no obvious (statistically significant) differences in retirement behavior between the former three groups when controlling for income. For men (and unlike women), the decisive factor in the retirement behavior seems to be the presence or absence of family ties or not.[21] There are several potential explanations for this finding paralleling the reasoning for women. The first is that a later labor market exit of married men is the result of a joint retirement decision. The second may be financial considerations. The overwhelming majority

[20]The age difference in Switzerland is approximately 3 years on average. This number is likely to understate the true age difference of a couple at retirement, as most second marriages display a larger age difference (the divorce rate in Switzerland is approximately 40%).

[21]The importance of familiy ties (particularly for men) seems to be important for another retirement decisions, the choice between an annuity and a lump sum upon retirement (see Bütler & Teppa (2004)). The absence of family ties induces men to opt the annuity option, probably because the annuity is the only form of insurance available to these men.

of today's elderly couples have followed a traditional role model in which the husband is the main (or even the only) bread-winner who has to care not only for himself, but also for his wife and children (who may still be at school). A third potential explanation is that there are large mortality differences between married and non-married men in favor of the former. If reductions to benefits for early retirement are actuarially fair, it is simply not optimal for married men to retire before the statutory age. This effect is reinforced by the joint annuity model in Switzerland (early retirement would entail that future benefits for the surviving wife are reduced at the same rate).

Retirement income has a clear non-monotonic impact on the retirement age. Up to an income that corresponds to the median second pillar income, a higher annuity (and thus a higher average lifetime income) leads to an earlier retirement, although lower life-expectancy for lower income workers should lead to the opposite outcome. For men, this affordability effect is much stronger at lower incomes than for women. It is important to stress that a median retirement income from the second pillar is clearly above the median income of *all* retirees, as low-income earners are not covered by the second pillar. The estimated peaks in the hazard ratio (at a second pillar income of 24'000 SFR $\approx$ 16'000 EU for non-censored regressions, and 35'000 SFR $\approx$ 23'000 EU for uncensored) correspond to a yearly pre-retirement income of approximately 80'000 to 110'000 SFR (55'000 to 72'000 EU). The dependency of the hazard ratio on income is also depicted in Figure 6.

Affordability thus seems to be a key determinant of male retirement behavior. There is a tendency to retire as soon as the financial situation permits (and as soon as early retirement plans are available). Another explanation may be that men have usually worked all their lives, in contrast to many women who had worked only part of their lives. Men may also suffer from worse health and thus retire as soon as the financial situation permits (though this explanation is not fully compatible with the rich — and thus on average healthier — individuals retiring earlier!). For individuals whose retirement income is above the median second pillar pension, affordability is unlikely to play an important role. Why the very rich again retire somewhat later than the medium rich is not obvious. The attractiveness of the job may play a role at high income levels.

insert Tables 2 & 3, and Figures 5 & 6 here

We have also included interaction terms between marital status und retirement income in our regressions to get more informative results. All these variables turn out to be insignificant. Whether this is due to a true absence of interaction effects or due to the quality of the data, we cannot tell at this stage.

To assess the sensitivity of our results, we have conducted regressions with various subsets of the data set. Tables 4 and 5 report some of these results on the firm level for both men and women, Table 6 estimates the coefficients for the three years with the highest number of observations in the dataset. Estimations on the firm level do not differ greatly from the overall regressions for the impact of marital status and income. Due to the much smaller number of observations the significance levels are far lower.

One particularly interesting feature of company level estimations (Tables 4 and 5) is that they convey large fluctuations in the exit rate over the years (see Figure 5). The incidence of early retirement is higher when retirement schemes become more flexible and lower in years following such changes. In some cases no cause for a big fluctuation could be identified. It could well be that due to financial difficulty of a firm or higher returns on invested pension capital, more people were induced to take up early retirement, altough this was not publicly admitted.

To exlude the year effect, we also estimated the model separately for the three years with the highest number of data points (Table 6). Again, results seem fairly robust, although significance levels have substantially deteriorated.

insert Tables 4-6 here

There are, of course, many other determinants for which an impact on the retirement decison can be anticipated, like health status, mortality differences or the number of dependent children at retirement. A bad health status is likely to induce early retirement regardless of the amount of annuity the person could get.[22] Mortality differences may have an impact on both the timing of retirement and the choice of the payout option. As differences in mortality are usually private knowledge,[23] the best we can do is to include proxies like life-time income (the rich live longer than the poor), and marital status (married men live longer than singles). The impact of having dependent children on the retirement decision is unclear, *a priori*. People may want to keep on working to be able to finance their children's expenses. But they also might want to benefit from the generous additional benefits for children (even if reduced due to early retirement) as long

[22]Through the fact that health is usually negatively correlated with (lifetime) income, it is not completely absent from our analysis. It may be the case that less healthy individuals might prefer to retire early, but cannot afford to do so. It is hoped that more complete data sets may help to clarify this issue in the future.

[23]Even if differences in mortality were observable, they would most likely not be eligible as criteria for lower or higher pension benefits.

as they are still eligible. The overall effect will depend on the financial situation of a family as well as the age of the children.

7 Conclusions

Reversing early retirement trends has become a major policy issue in most European countries. It is clear that incentives set by the social security system will be key in this exercise. But there might be other determinants of early retirement that are equally important. If the preference for leisure in old age is sufficiently strong, for example, even negative implicit tax rates on staying in the labor force might not induce people to work much longer. This paper has aimed to shed some light on determinants of the retirement decision other than the impact of social security incentives by analysing individual data from a selection of Swiss pension funds.

The main findings from our exercise can be summarized as follows. Firstly, there is an increasing tendency to retire early in Switzerland even in the absence of legislative changes. The effect is more pronounced for men than for women, and was found to be especially strong in the last few years. Secondly, affordability seems to be a key determinant for the retirement decision, in particular for men. Richer men (as measured by lifetime labor income) retire earlier than poorer men at least up to a relatively high income. For women, the effect of income on the likelihood to exit the labor force is also positive, but far weaker than for men. This affordability interpretation may also partially explain the increase in early retirement over the last 15 years, as Switzerland's second pillar has matured over this period, leading to higher effective replacement rates. Thirdly, marital status plays an important role in an individual's retirement decision. For men, the main difference is between singles, who retire earlier on average, and non-singles. This hints at the importance of family ties (and of potential financial liabilities for children and (ex-)wives) for men. Married women tend to retire earlier than other women, while divorced and separated women clearly work longer, probably due to financial constraints.

Combining our results, it seems that individuals choose their labor market exit in a constrained rational way — at least to a certain degree. If financial constraints are not binding, differential mortality (as mirrored in lifetime income and marital status) and joint retirement decisions have the impact predicted by economic theory. This complements another result of ours with the same dateset (Bütler & Teppa (2004)). The decision between a lump sum and an annuity upon retirement can be explained by a combination of rational choice (based on differences in mortality, the need for insurance and the desire to leave bequests) and an "acquiescence bias" — a majority of individuals choose the standard option offered by their pension plan.

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Variable	Obs.	in %	Median	Mean	(Std.)
Age at retirement	8452		62.0	61.82	(2.70)
female	3084	36.5	62.0	60.90	(2.58)
single	500	16.2	61.1	60.66	
married	1621	52.6	61.1	60.66	
widowed	279	9.1	62.0	61.39	
divorced / separated	684	22.2	62.0	61.42	
Age at retirement male	5368	63.5	62.1	62.35	(2.62)
single	293	5.5	62.2	61.55	
married	4587	85.4	62.0	62.40	
widowed	161	3.0	63.0	62.84	
divorced / separated	327	6.1	62.0	62.08	
Statutory retir. age	2665	31.5			
(female)	1013	32.9			
(male)	1652	30.8			
Annuity deflated	8452		31'488	36'666	(30245)
female	3084		19'017	24'092	(19579)
single	500		34'196	35'364	
married	1621		15'545	21'973	
widowed	279		12'919	18'920	
divorced / separated	684		19'290	22'984	
Annuity deflated male	5368		37'198	43'889	(32820)
single	293		29'397	33'525	
married	4587		38'535	45'156	
widowed	161		30'262	39'705	
divorced / separated	327		31'456	37'471	
Non–standard option	576	6.8%			
(female)	149	4.8%			
(male)	427	8.0%			
Lump–sum capital (in %)	649	7.7%	50.3%	60.0%	(36.7%)
(female)	179	5.8%	100.0%	78.0%	(29.7%)
(male)	470	8.8%	44.6%	53.2%	(36.8%)

Table 1: *Summary statistics for some relevant variables*

Covariate	**I(f)** Haz. Ratio *(p–value)*	**II(f)** Haz. Ratio *(p–value)*	**III(f)** Haz. Ratio *(p–value)*	**IV(f)** Haz. Ratio *(p–value)*
ret. year (linear)	0.9550	—	0.9165	—
	(0.000)	—	*(0.000)*	—
ret. year (dummy)	—	YES	—	YES
	—	*(*)*	—	*(***)*
married	1.1044	1.0979	1.1437	1.1362
	(0.095)	*(0.127)*	*(0.070)*	*(0.094)*
widowed	0.9322	0.9273	0.9005	0.8952
	(0.317)	*(0.287)*	*(0.259)*	*(0.243)*
divorced/separated	0.8752	0.8658	0.8278	0.8184
	(0.029)	*(0.017)*	*(0.013)*	*(0.008)*
ln(annuity)	1.0004	1.0004	1.0016	1.0015
	(0.208)	*(0.281)*	*(0.000)*	*(0.001)*
Censoring	NO	NO	YES	YES
Comp. fixed effects	YES	YES	YES	YES
log p-lik.	-21913.29	-21887.18	-14624.99	-14598.19
observations	3084	3084	3084	3084
failures	3084	3084	2071	2071

Table 2: *Cox proportional hazard regression for women. The variable "ln(annuity)" has been interacted with a linear time trend. Data censored for age at retirement 61.75-62.25 (if censoring = YES).*

Covariate	**I(m)** Haz. Ratio *(p–value)*	**II(m)** Haz. Ratio *(p–value)*	**III(m)** Haz. Ratio *(p–value)*	**IV(m)** Haz. Ratio *(p–value)*
ret. year (linear)	0.9507	—	0.9158	—
	(0.000)	—	*(0.000)*	—
ret. year (dummy)	—	YES	—	YES
	—	*(***)*	—	*(***)*
married	0.7960	0.8034	0.7118	0.7181
	(0.000)	*(0.000)*	*(0.000)*	*(0.000)*
widowed	0.7142	0.7356	0.6475	0.6822
	(0.002)	*(0.002)*	*(0.001)*	*(0.003)*
divorced/separated	0.8414	0.8481	0.8169	0.8180
	(0.032)	*(0.046)*	*(0.043)*	*(0.049)*
ln(annuity)	1.0391	1.0396	1.0490	1.0499
	(0.000)	*(0.000)*	*(0.000)*	*(0.000)*
ln(annuity)^2	0.9981	0.9981	0.9977	0.9977
	(0.000)	*(0.000)*	*(0.000)*	*(0.000)*
hazard max at	24’918 SFR	24’222 SFR	35’792 SFR	33’905 SFR
Censoring	NO	NO	YES	YES
Comp. fixed effects	YES	YES	YES	YES
log p-lik.	-41222.82	-41170.7	-29317.60	-29270.98
observations	5368	5368	5368	5368
failures	5368	5368	3716	3716

Table 3: *Cox proportional hazard regression for men. The variables “ln(annuity)” and “ln(annuity)^2” have been interacted with a linear time trend. Data censored for age at retirement 64.75-65.25 (if censoring = YES).*

Covariate	**IV(f;1)** Haz. Ratio *(p–value)*	**IV(f;10)** Haz. Ratio *(p–value)*	**IV(f;11)** Haz. Ratio *(p–value)*	**IV(f;14)** Haz. Ratio *(p–value)*
ret. year (dummy)	YES (99–00)	YES (90–02)	YES (90–03)	YES (01–02)
	(–)	*(***)*	*(**)*	*(–)*
married	4.9387	1.0530	1.3981	0.9346
	(0.016)	*(0.567)*	*(0.179)*	*(0.684)*
widowed	2.0659	0.8139	1.2882	0.7530
	(0.361)	*(0.087)*	*(0.351)*	*(0.278)*
divorced/separated	3.0080	0.7710	0.9115	0.6715
	(0.109)	*(0.005)*	*(0.671)*	*(0.026)*
ln(annuity)	1.0018	1.0020	0.9969	1.0698
	(0.208)	*(0.001)*	*(0.067)*	*(0.002)*
ln(annuity)^2	—	—	—	0.9963
				(0.002)
hazard max at	—	—	—	9'967 SFR
Censoring	YES	YES	YES	YES
Comp. fixed effects	YES	YES	YES	YES
log p-lik.	-529.97	-8673.22	-704.58	-1101.17
observations	228	1891	192	250
failures	106	1323	163	234

Table 4: *Cox proportional hazard regression for women by company. The variables "ln(annuity)" and "ln(annuity)^2" have been interacted with a linear time trend. Data censored for age at retirement 61.75-62.25 (if censoring = YES).*

Covariate	**IV(m;2)** Haz. Ratio *(p–value)*	**IV(m;9)** Haz. Ratio *(p–value)*	**IV(m;10)** Haz. Ratio *(p–value)*	**IV(m;15)** Haz. Ratio *(p–value)*
ret. year (dummy)	YES (00–03)	YES (00–02)	YES (90–02)	YES (90–03)
	(*)	*(*)*	*(**)*	*(**)*
married	0.5407	0.7548	0.7896	0.6359
	(0.001)	*(0.113)*	*(0.073)*	*(0.006)*
widowed	0.6365	0.9486	0.6744	0.6906
	(0.149)	*(0.848)*	*(0.088)*	*(0.116)*
divorced/separated	0.7059	1.0458	0.9023	0.6784
	(0.206)	*(0.861)*	*(0.543)*	*(0.075)*
ln(annuity)	1.1597	1.0001	1.1271	1.0037
	(0.000)	*(0.946)*	*(0.000)*	*(0.000)*
ln(annuity)^2	0.9931	—	0.9943	—
	(0.000)		*(0.000)*	
hazard max at	42'254 SFR	—	33'984 SFR	—
Censoring	YES	YES	YES	YES
Comp. fixed effects	YES	YES	YES	YES
log p-lik.	-4070.30	-2927.90	-9298.49	-2901.78
observations	762	600	2135	937
failures	695	489	1313	460

Table 5: *Cox proportional hazard regression for men by company (4 largest companies). The variables "ln(annuity)" and "ln(annuity)^2" have been interacted with a linear time trend. Data censored for age at retirement 64.75-65.25 (if censoring = YES).*

Covariate	**IV(m;2000)** Haz. Ratio *(p–value)*	**IV(m;2001)** Haz. Ratio *(p–value)*	**IV(m;2002)** Haz. Ratio *(p–value)*
married	0.7721	0.6334	0.8387
	(0.178)	*(0.004)*	*(0.287)*
widowed	0.9726	1.0533	0.7704
	(0.926)	*(0.868)*	*(0.354)*
divorced/separated	0.8900	0.6298	1.0108
	(0.674)	*(0.044)*	*(0.962)*
ln(annuity)	1.1098	1.0553	1.0451
	(0.001)	*(0.065)*	*(0.133)*
ln(annuity)^2	0.9950	0.9974	0.9978
	(0.001)	*(0.063)*	*(0.130)*
hazard max at	32’096 SFR	27’481 SFR	—
Censoring	YES	YES	YES
Comp. fixed effects	YES	YES	YES
log p-lik.	-4035.04	-4718.67	-3564.48
observations	884	919	749
failures	663	876	598

Table 6: *Cox proportional hazard regression for men by retirement year. The variables “ln(annuity)” and “ln(annuity)^2” have been interacted with a linear time trend. Data censored for age at retirement 64.75-65.25 (if censoring = YES).*

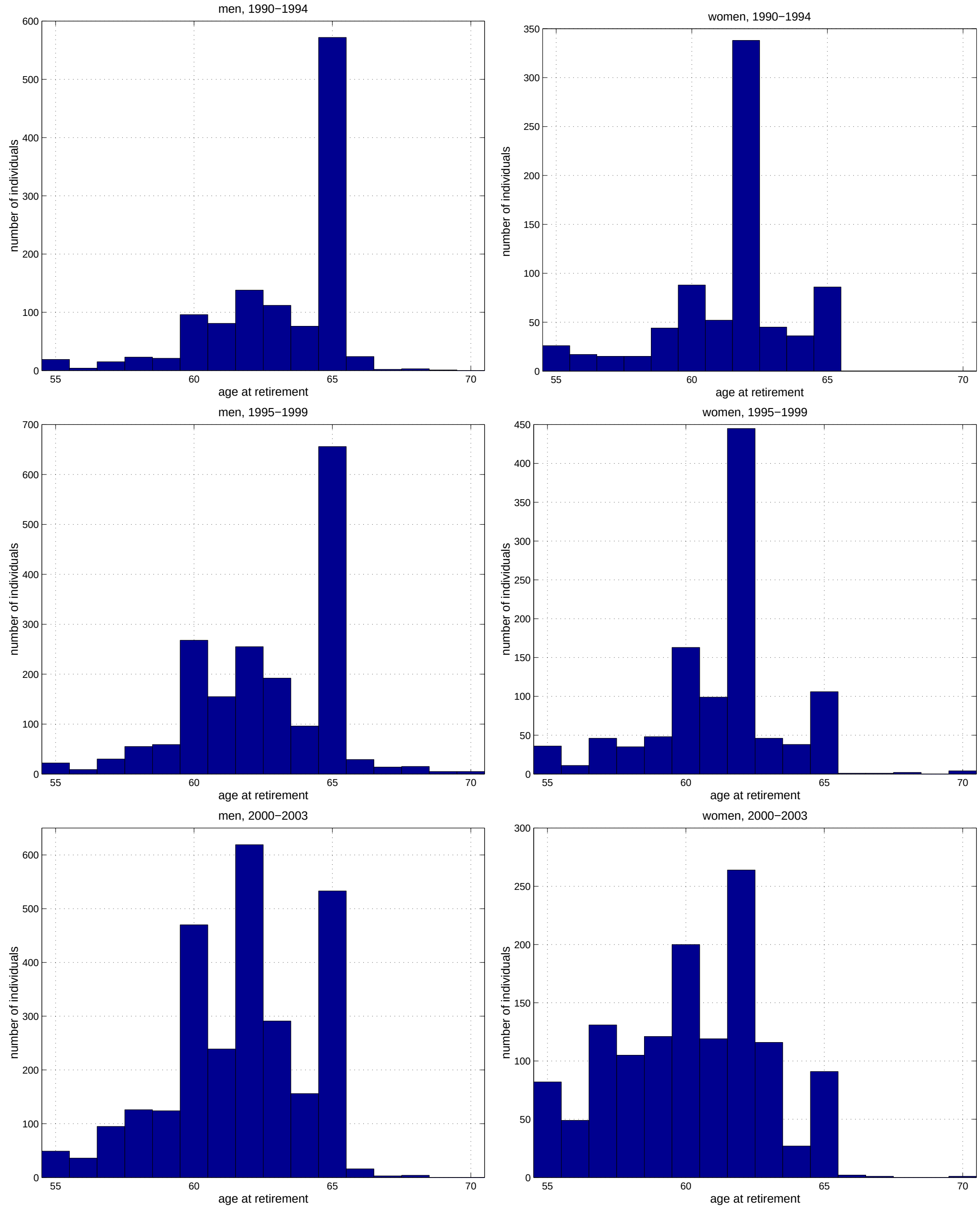

Figure 1: *Distributions of age at retirement for men (left-hand side) and for women (right-hand side)*

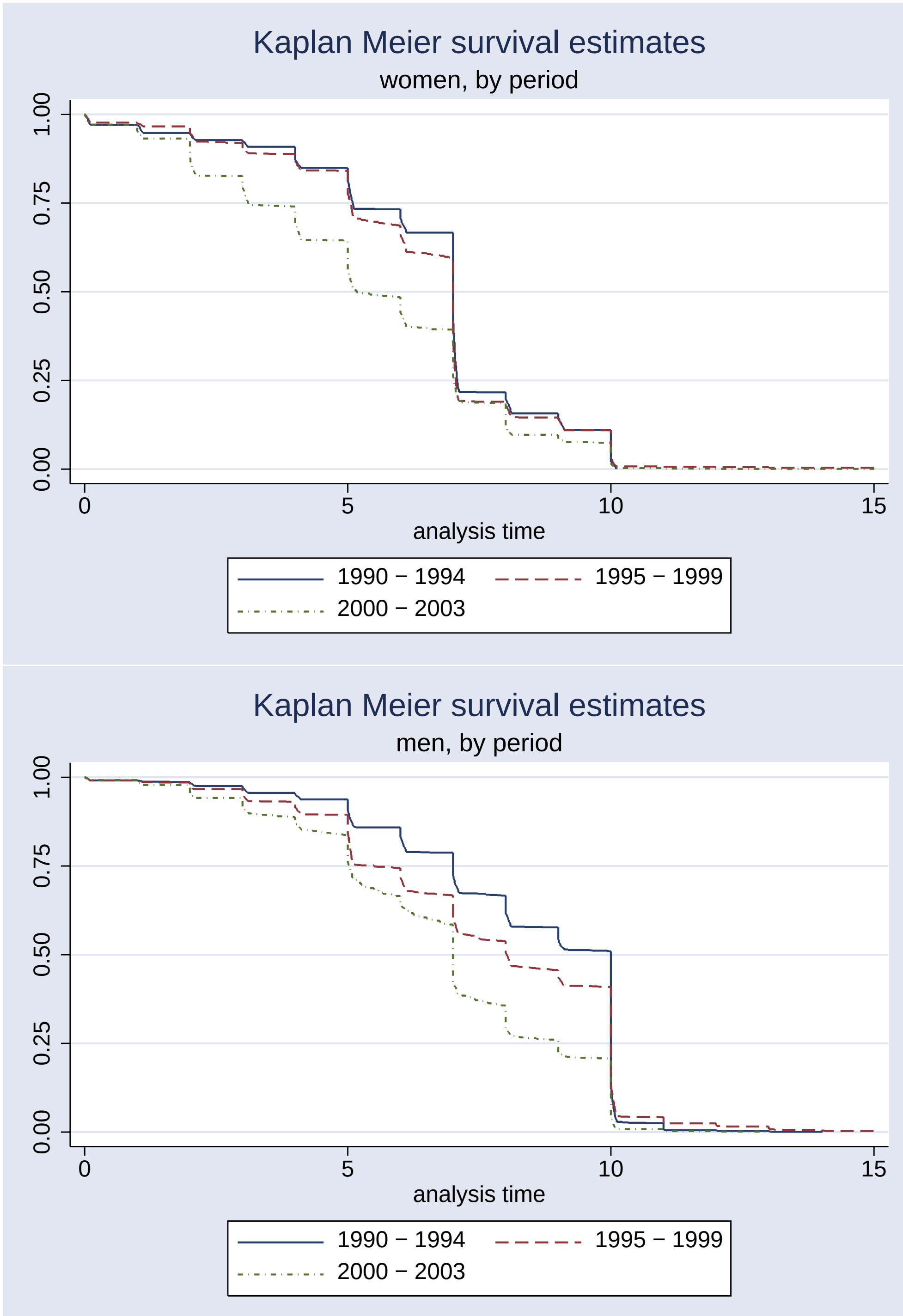

Figure 2: *Kaplan-Meier estimator without censoring by period for women (upper panel) and men (lower panel). The numbers on the horizontal axis denote the years after age 55.*

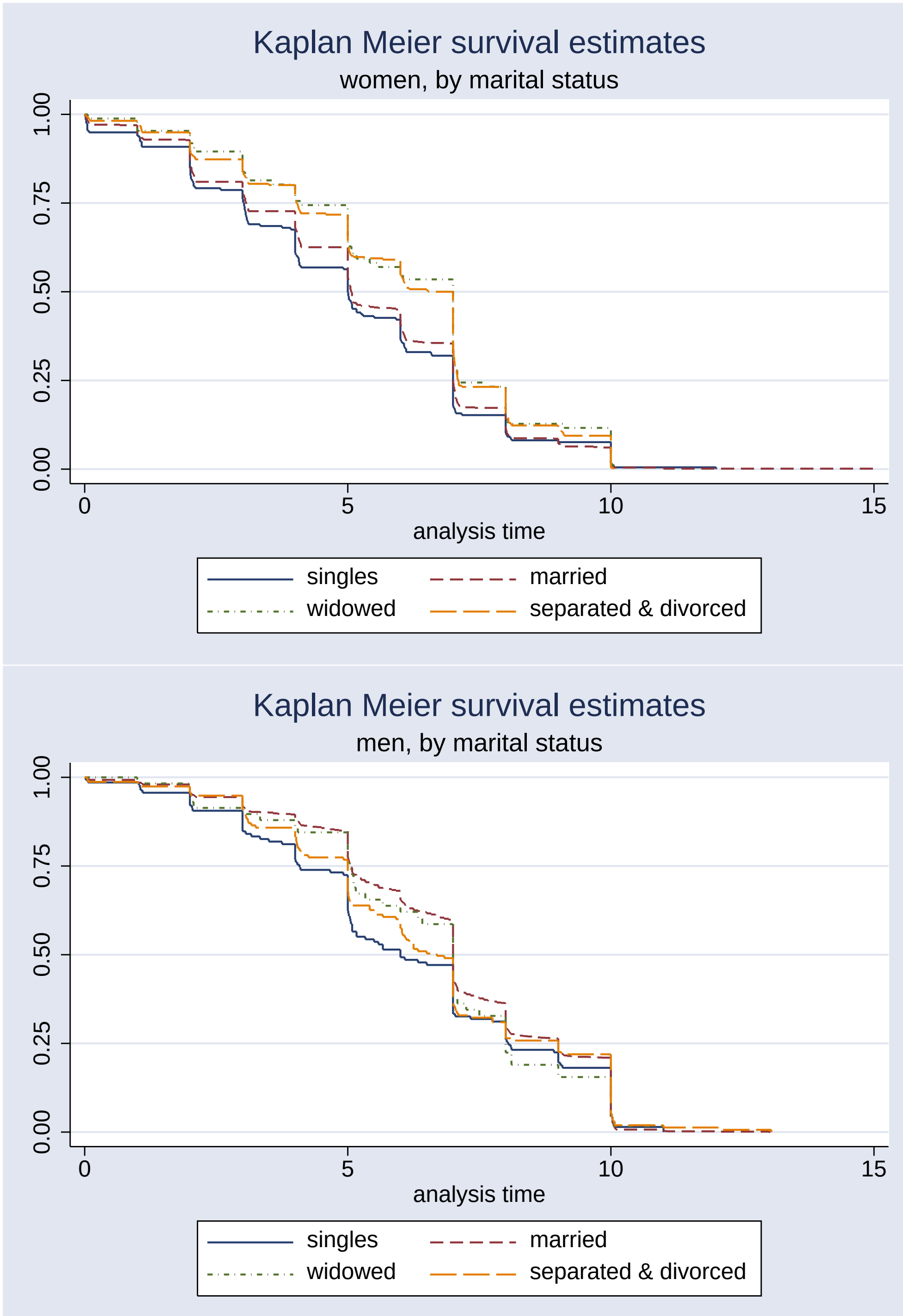

Figure 3: *Kaplan-Meier estimator without censoring for the period 2000-2003 by marital status for women (upper panel) and men (lower panel). The numbers on the horizontal axis denote the years after age 55.*

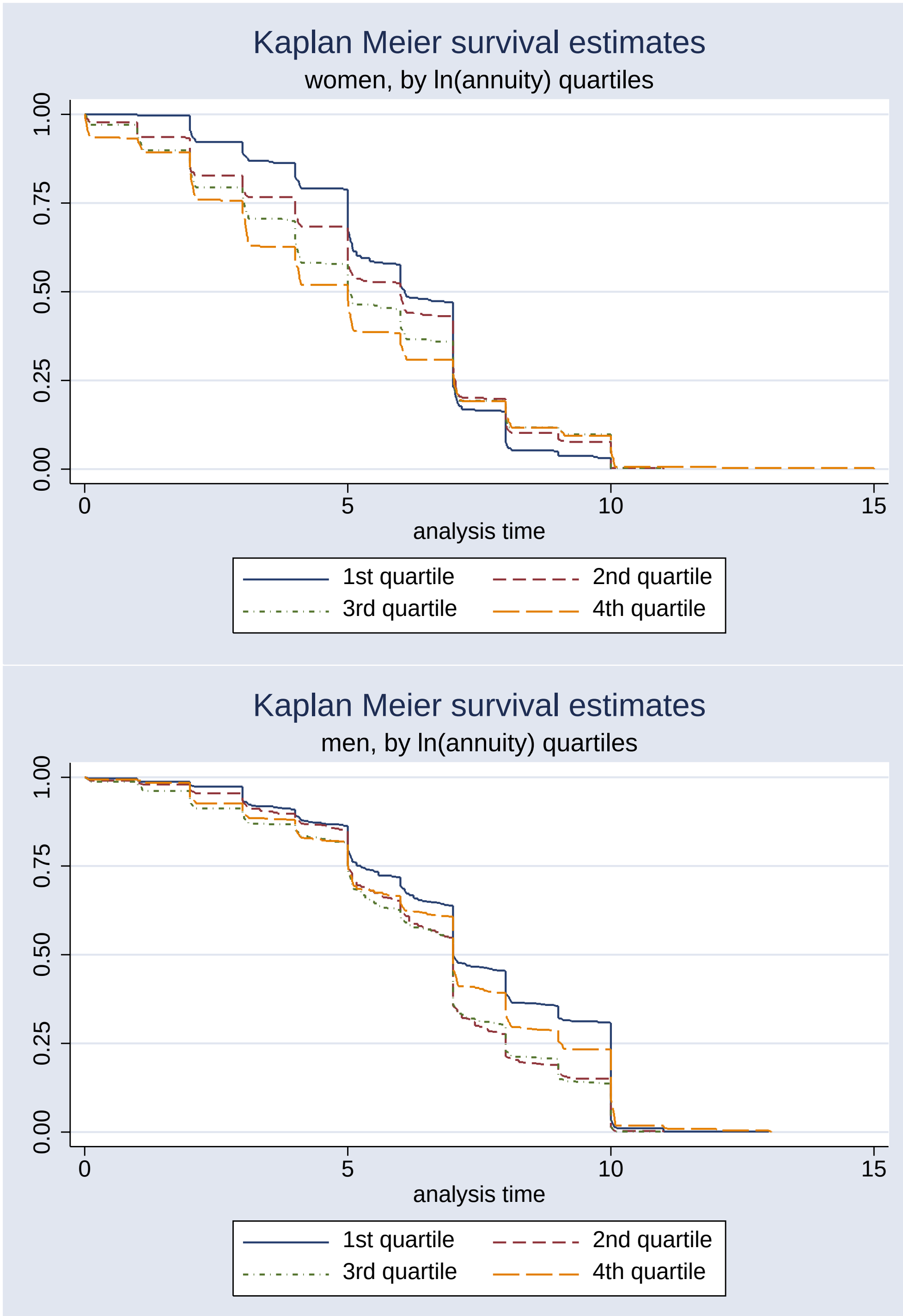

Figure 4: *Kaplan-Meier estimator without censoring for the period 2000-2003 by income quartiles for women (upper panel) and men (lower panel). The numbers on the horizontal axis denote the years after age 55.*

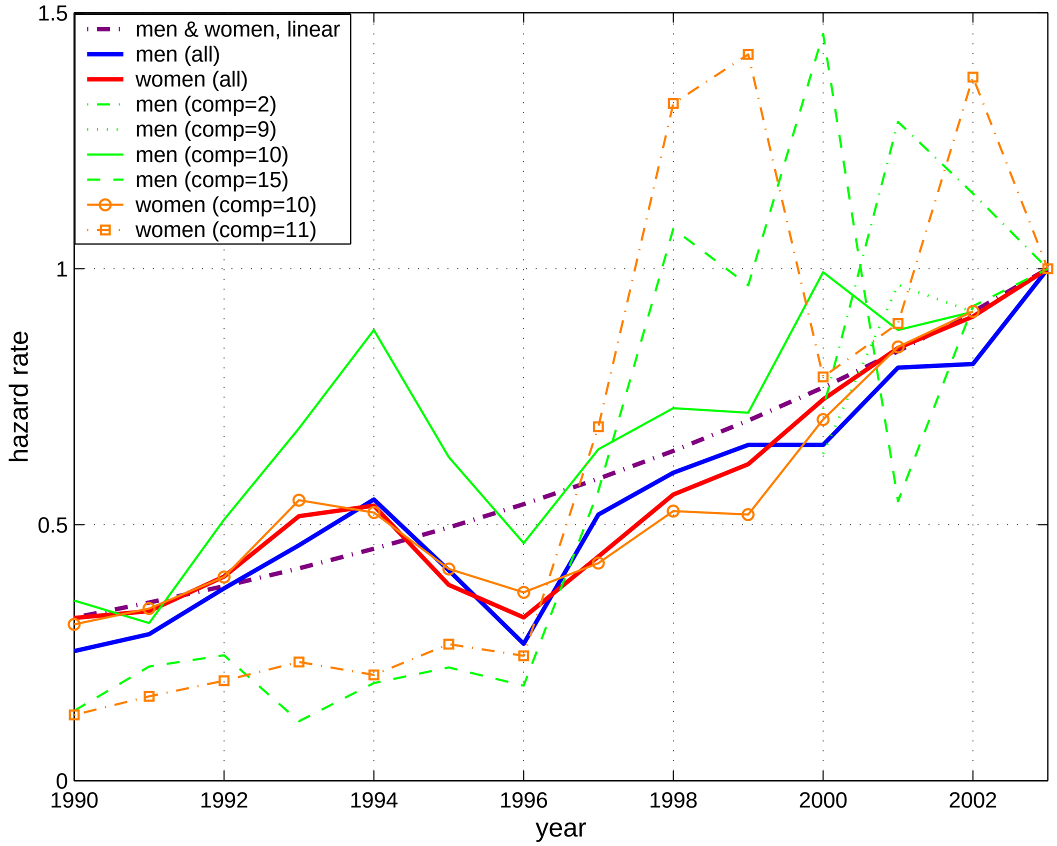

Figure 5: *Relative hazard rates for year of retirement (base year 2003).*

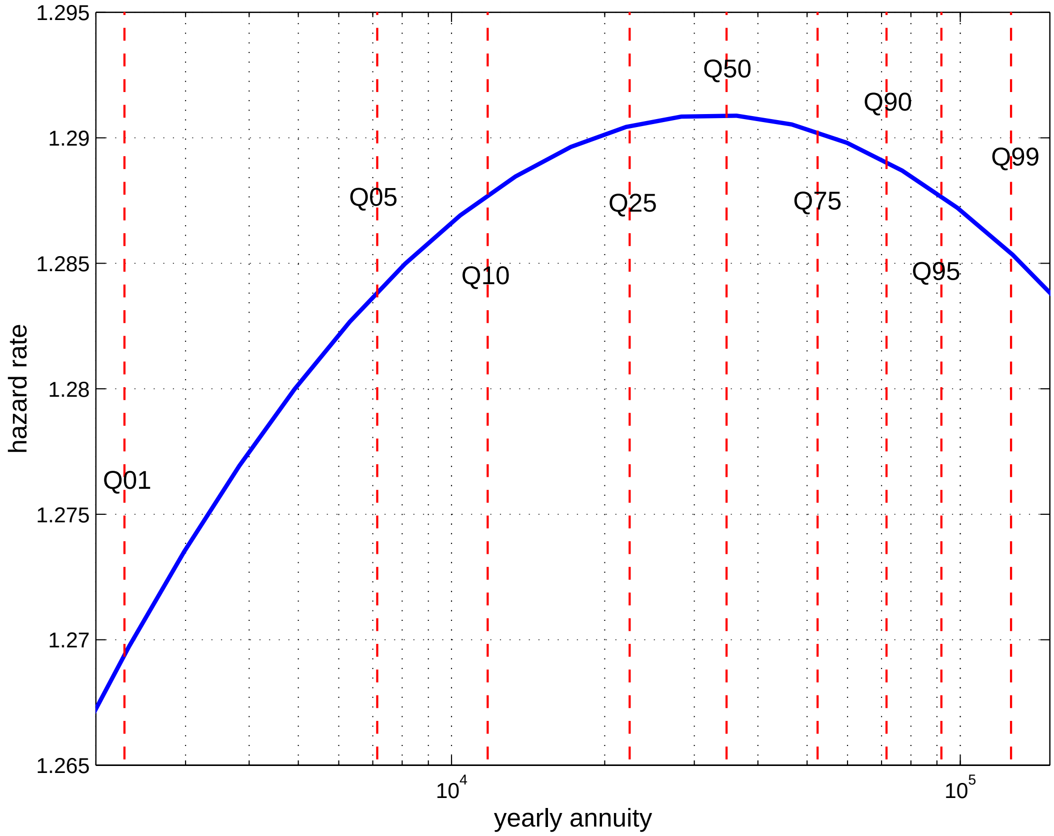

Figure 6: *Relative hazard rates for estimation IV(m) as a function of yearly deflated annuity (base = annuity of 1 SFR). 'Qx' denotes the xth quantile of the annuity distribution for men.*