The Role of the Annuity's Value on the Decision (Not) to Annuitize: Evidence from a Large Policy Change

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Abstract

This paper presents new evidence on how the annuitization decision is affected by changes in the annuity's value. We take advantage of an unprecendented change in policy, which in 2004 moderated the super-mandatory Swiss occupational pension scheme: The 20 percent reduction in the rate at which retirement capital is translated into a life-long annuity equates to a net present value loss of approximately 20'000 SFR (20'000 US$) for the average retiree. Using administrative data and correcting for anticipation effects, we show that due to the change in policy there was an approximately 8 percentage point change in the share of men choosing to annuitize their savings. We also show that the estimated responsiveness of the cash-out decision to variations in a utility based measure for the annuity value is comparable to results of previous studies, which employed completely different sources of variation in the annuity's value.

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Keywords: annuity puzzle, occupational pension, policy change.

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1 Introduction

Economic theory predicts that most people should annuitize, but when international numbers are analyzed it is apparent that when given a choice, only a minority do so voluntarily. This raises questions with regard to the adequacy of income provided in old age in many countries: As a consequence of unfavorable demographic and financially imbalanced social security systems, public pension annuity payments are declining and fully funded systems which typically do not mandate annuitization are playing a greater role.

The lack of voluntary annuitization is also puzzling given the numerous theoretical findings following the seminal paper of (Yaari (1965)), which all suggest sizeable benefits to annuitization. A great amount of literature has attempted to shed light on the "annuity puzzle" (see Brown (2007) for an excellent review of this literature), but it has failed to present a convincing general explanation. While adverse selection, administrative loads, the existence of first-pillar annuities, intra-family risk-sharing, bequest motives,[1] and a desire to insure against expenditure spikes can rationalize the preference for the lump sum to some degree, the low annuitization rates remain hard to reconcile with economic theory. Recent work on the determinants of individual cash-out behavior includes not fully rational behavior.[2]

Given the very low level of voluntary annuitization observed in the real world, the puzzle's solution is even more difficult due to the lack of suitable data. Empirical evidence that clearly identifies the determinants of the cash-out decision at retirement are thus scarce. Notable exceptions are Hurd et al. (1998) on pension cash-outs when changing jobs or retiring, Brown (2001) analyzing ex ante intentions to annuitize, and Bütler and Teppa (2007) on the annuitization decision in the context of the mandatory Swiss occupational pension scheme.

This paper analyzes the annuitization decision at retirement taking advantage of an unprecedented and sudden change in conversion rates recently implemented in Switzerland. In 2004, most large insurance companies (covering approximately 10% of the work force) reduced their conversion rates by almost 20%. These conversion rates are used to calculate a life- long annuity based on the

[1] In a recent working paper Lockwood (2008) presents microsimulations that demonstrate that the bequest motive might be able to explain low annuitization rates.

[2] See, for example, Brown et al. (2008) who find that people are more likely to annuitize when the choice is presented to them in a consumption framework than when it is presented in an investment framework.

accumulated retirement capital in the super-mandatory part of the second pillar Super-mandatory pensions represent a considerable part of retirement income in Switzerland, providing approximately 20 percent of total retirement income for an average retiree. Reducing the conversion rate as was done here is equivalent to a loss in present value of 20'000 SFR (≈$ 20'000) for an average individual.

To the best of our knowledge, our paper is the first to analyze such a large and truly exogenous variation in the annuity price. With this we establish a causal relationship between annuity values and cash-out behavior. The annuitization decision is analyzed using the administrative records of several Swiss insurance companies that provide occupational pension plans to small and middle-sized firms that represent all branches in the economy. The data includes approximately 6'000 individual annuitization decisions made between the years of 2001 and 2005, and provides reliable and complete information on pension plan details. Our setup evades the issue of individual self-selection into firms/plans based on unobserved characteristics, as well as the problem of omitted variables, which would impede the identification of the price effect.

A fraction of the individuals who had planned to retire after the policy change obviously anticipated the change and retired earlier in order to profit from the higher annuity. This is evident in the sharp increase in the number of annuitants in the months just preceding the conversion rate reduction. Due to this anticipation effect, a simple comparison of the cash-out behavior before and after the policy change is likely to overestimate the effect of the policy change. In order to address this problem of potential anticipation we make use of the fact that the cost of retiring earlier is related to the difference in age between that of just before the policy change and the statutory retirement date. This allows us to construct a good proxy for the likelihood an individual proactively anticipates the change in policy. Once this anticipation is accounted for, we observe a constant annuitization rate before the policy change and an approximately 8 percentage point drop in annuitization in the year after the change. The annuitization rate also remains constant in the periods preceding the reform and following the change, respectively, when the anticipation component is filtered out.

While the approach to regress the annuitization decision on individual controls and a dummy indicating the policy change is straightforward, it has two shortcomings. First, it does not include any measure for the annuity's value, which would allow us to calculate the value's elasticity in the annuitization decision. Second, the only source of identification comes from the variation in

annuity value over time. Yet, the annuity value also varies across individuals. To address these points, we construct a utility based measure of the annuity value (as pioneered in Brown (2001)), and include it as an additional regressor. We find that a one percentage point increase in annuity equivalent wealth increases the annuitization rate by approximately 0.8 percentage points, after controlling for anticipation. This estimate is remarkably similar to estimates derived under completely different conditions (Brown (2001) and Bütler and Teppa (2007)).

The paper proceeds as follows. Section 2 describes the key features of the Swiss social security system and the change in policy. Section 3 describes the data set and analyzes the macroeconomic environment. Section 4 outlines the theoretical and empirical strategy to analyze the effects of the policy change. Section 5 presents the results and the conclusion is presented in section 6.

2 The Swiss social security system and pension plan details

2.1 The importance of the second pillar

Switzerland's pension system is made up of three pillars, in which the first and second are more or less equally important.[3] The first pillar 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. To some degree, benefits depend on the number of years contributed and the average working income, although in reality the benefit structure is relatively flat. The statutory retirement age is 65 for men and 64 for women. At the earliest men may claim first piller benefits at 63 and women at 62 (at actuarially fair reductions in pension benefits).[4]

The second pillar, introduced on January 1, 1985, is an employer-based, fully funded occupational pension scheme and is mandatory for all employees whose annual income exceeds a certain minimum. The third pillar is earmarked and

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

[4]Retirement at 65/64 is not mandatory by law, but rather the age of 65 for men or 64 for women is the condition of eligibility for claiming public pension benefits. Most labor contracts specify a retirement age that coincides with the age of eligibility.

benefits from tax-preferential treatment. For the case in which the combined pension income is not enough to cover basic needs in old age, means-tested supplemental benefits may also be claimed. These additional benefits usually result in an income that is well above the poverty threshold. In 2004 89.4 percent of working men and 71.1 percent of working women were involved in an occupational pension plan, which corresponds to 81.2 percent of the total workforce. An employer can choose between different organizational structures for its occupational pension plan. These range from setting up a completely autonomous pension fund, which covers all risks to outsourcing the scheme entirely to an insurance company. The latter is relatively common particularly for small and medium sized companies.

The main goal of the occupational pension system is to maintain pre-retirement income in addition to benefits from the first pillar. It insures income above that level that is covered by the first pillar.[5] Most pension funds aim at a total replacement rate of approximately 50-60 percent of insured income, including income from the first pillar. The net replacement rate after taxes often amounts to 70-80 percent for an uninterrupted career even for higher categories of income, and can reach 100 percent for beneficiaries with dependent children.

The insured income above the lower threshold and below the upper threshold[6] is called the *mandatory component*, and the income above the upper threshold is called the *super-mandatory* component of the second pillar. Pension insurers are required by law to insure the mandatory share. They are free to provide insurance for the super-mandatory share, but virtually all do. This is mainly a consequence of the second pillar being thought of as an important attribute in attracting a well-educated workforce to Switzerland's tight labor market.

The mandatory part of the second pillar is subject to stringent regulation with respect to minimum contribution rates, minimum interest rates and the rate at which the accumulated pension wealth is translated into an annuity. Although most pension plans are set up on a defined contribution base, the income guarantees (minimum accrual and interest rates, conversion factor) mandated by law make the defined benefit and contribution schemes in practice very similar. The federal law is less restrictive with respect to the contract conditions offered by

[5] In 2004 the threshold was approximately 20'000 SFR (≈ $ 20'000). The threshold explains the lower rate of coverage for women, who often work part-time and thus on average earn less.

[6] The upper threshold is equal to three times the yearly maximum single first pillar pension (i.e. 75'960 SFR or $ 56'220 in 2004).

the insurance companies in the super-mandatory part. The minimum interest rate in the super-mandatory part, which is set by the pension fund, is normally lower than the minimum interest rate in the mandatory part and can be adjusted according to fund performance.

The accrued capital is fully transferable when the individual changes employers. By law, when an employee goes from one company to another he receives all of the accumulated contributions (including the employer's part), but the full sum has to be paid into a new fund. The total amount of assets at retirement has thus been accumulated over the entire working lifetime and is a good proxy for lifetime income.

2.2 Options to withdrawal of the Swiss second pillar

The accumulated retirement capital can be withdrawn either as a monthly lifelong annuity (including a 60 percent survivor benefit), a lump sum or a mix of the two options. In some plans the cash-out limit is equal to 50 or 25 percent (the legal minimum) of accumulated capital. The individual must declare his choice between three months and three years prior to the effective withdrawal date depending on insurer regulations. Many pension insurers define a default option for the case when the beneficiary does not make an active choice.

Occupational pension annuities are strictly proportional to the accumulated retirement assets (contributions made during the working lifetime plus accrued interest). The capital K is translated into a yearly pension B using the so-called *conversion rate* γ: $B = \gamma K$. The conversion rate is independent of marital status, but depends on retirement age and gender. The law stipulates a minimum conversion rate, which is currently 7.1 percent but was equal to 7.2 percent until 2004 and 7.15 percent in 2005.[7] In the super-mandatory part, pension insurers are free to set the conversion factor.

A majority of retirees covered by Swiss occupational pension plans choose the annuity, despite the first pillar already providing an annuity stream covering subsistence needs in old age and lump sums profiting from preferential tax treatment in many cantons (the Swiss states). In a related paper, Bütler and Teppa (2007) analyze the annuitization decision at retirement, making use of variations in the annuity's value caused by differences in company pension plans. They show that the annuity's value is the most important determinant in the cash-out

[7]The conversion rate will be lowered continuously to 6.8 percent in 2015.

decision, but that there are also great differences between companies. Individuals often choose the standard option offered by the company or seem to follow their peers. Small stocks of old age capital are much more likely to be withdrawn as lump sums. This is probably a consequence of the fact that for small amounts of capital, it may be optimal to spend all resources in order to qualify for means-tested social assistance. As in other studies, Bütler and Teppa (2007) find mixed evidence for the existence of a bequest motive. This in large part is due to the limited amount of information they had on individual backgrounds in their data.[8]

In light of some recent behavioral explanations for the low demand for annuities (such as Brown et al. (2008)), the unusually high degree of annuitization in Switzerland may be attributed to the prevailing framing in the scheme: By law, pension insurers must provide each year all of their insurees with a statement in which not only the accumulated capital to date is reported, but also the expected anticipated annuity stream (based on an extrapolation of current earnings and interest rates). Although the accumulated capital is reported on the statement, the space dedicated to annuity streams (including survivor benefits and disability benefits) is much bigger. The statement thus comes close to what Brown et al. (2008) call a consumption frame.

3 Data

Our data set consists of administrative records of individuals covered by occupational pension plans provided by large Swiss insurance companies. The private firms which outsource their second pillars to these insurance companies are typically small and are active in several branches of the economy. The sample is restricted to defined contributions (DC) plans. The data is a repeated cross-section, each individual is only observed once at retirement. For the companies in our sample, we were given information about all employees who retired between 2001 to 2005. We use the records for the years 2001 to 2003 as the "before" period and those for 2004 and 2005 as the "after" period.

In the analysis we will focus on men for two reasons. First, the retirement age for women was raised twice over the period of interest (in 2001 from 62 to 63 and

[8] Bütler and Teppa (2007) report that divorced and widowed men (who are much more likely to have children) have a higher propensity to cash-out than singles at the margin, although the former have a higher remarriage rate and should therefore choose the annuity more often. This finding may be interpreted as an indicator for the existence of a bequest motive.

in 2005 from 63 to 64). Second, women generally have a much lower amount of capital stock in the super-mandatory part and are thus affected much less by the decrease in the annuity's value. We also exclude from our sample 44 individuals with a total capital stock more than 1'500'000 SFR. Men with a higher capital stock are likely to be covered by supplementary insurance for managers. The chosen threshold corresponds to approximately four times the average second pillar capital stock at retirement as calculated in Bütler and Teppa (2007). We also drop 37 individuals who retired before reaching the age of 60. Retirement before age 60 is very unusual for the companies in our sample and is often an indication of poor health or a difficult employment situation. The final sample thus consists of 5'855 men, with retirement ages between 60 and 70.

For each individual we have information on their date of birth, retirement date, annuitization decision, amount of accumulated capital stock, name of employer, earnings in the last year before retirement, as well as the individual specific conversion factor of the mandatory and super-mandatory amounts. For individuals retiring in 2003 and before, we only know their total capital stock, but not how it's split up between the mandatory and super-mandatory parts. However, the imputation of the amount in the super-mandatory part is straightforward and not sensitive to the chosen method. [9]

Nonetheless, the data suffer from a number of shortcomings. In particular, the insurance companies do not collect any information on non-pension wealth, marital status, education, health, occupation, or other indicators of socioeconomic status, even though such factors are likely to be related to the risk of mortality. As long as we have no reason to believe that these factors differ in a significant way from one year to the next, neglecting them should not critically affect our results.

3.1 Pension plan details and the policy change

The insurance companies in our sample provide insurance for both the mandatory and the super-mandatory parts of pension plans. The statutory age of retirement is set at 65 for men. Early retirement, starting at age 55, as well as working

[9] To calculate the amount of super-mandatory capital we used the estimated amount of super-mandatory capital as a function of the retiree's age and his total capital stock with the estimates taken from the 2004/5 retirees. Details about this imputation strategy can be found in the appendix. We also experimented with other ways to impute the super-mandatory share, but experienced little sensitivity as to the method of imputation.

beyond planned retirement is possible. With regard to early retirement, the conversion rate is reduced by approximately 3 percent for each month, while it is raised by around 2 percent per month if retirement occurs after the statutory retirement age. At retirement the individual can choose between an annuity, a full lump sum or a mixture of the two. The insurance companies in our sample require three months advance notice in making a decision and force individuals to make a timely decision (no defaulting).

Until the end of 2003 conversion rates of the mandatory and super-mandatory parts coincided with the statutory proposed conversion rates, despite the fact that the insurance companies were free to set the conversion rate of the super-mandatory part. According to the insurance companies, these high conversion rates threatened to create losses in the future. As a consequence, in 2003 most large insurance companies decided to drastically reduce their conversion rates of the super-mandatory part by January 2004, arguing that due to the demographic change and the increase in life expectancy, annuities have become too generous compared to the individual contributions. For a man retiring at age 65 the conversion rate was reduced from 7.2 percent to 5.835 percent and for a woman retiring at age 62 from 7.2 percent to 5.454 percent.[10] The reduction in the conversion rates was almost the same for all insurance companies. According to the insurance companies, this is because they based their calculations on the same mortality table.[11] The policy change was announced 7 months before it became effective, i.e., in June 2003, and was extensively discussion in the media.

3.2 Descriptive analysis

Table 1 provides summary statistics for the most important variables in our data set. Prior to the policy change, i.e. for 2001 to 2003, the conversion rate in the mandatory and super-mandatory part match and are practically constant over time. In 2004 and 2005 the conversion factor in the super-mandatory part is on

[10]The difference in conversion rates between men and women is the result of a lower female retirement age not a difference in longevity. As the survivor component of the second pillar is free, men and women cost approximately the same for the pension insurer provided they retire at the same age.

[11]Calculations of the adjusted conversion rates were based on a life expectancy of 83 years for men and 89 years for women.

average 18 percentage points lower than in the mandatory part.[12] The average retirement age remained fairly constant across the different years: 72.1 percent of all men in our sample retire at the statutory age of 65.

Approximately two thirds of the total accumulated capital is in the mandatory part and one third in the super-mandatory part. A large fraction of the beneficiaries chose a polar option (full lump sum or full annuity) and did not distinguish between the mandatory and super-mandatory part of the insurance, although the implicit annuity prices are dramatically different after 2003. Between 56.6-72.7 percent of all men chose a lump sum depending, another 25.5-39.4 percent chose full annuitization, whereas only 1.9-4.0 percent opted for a mixed option. The percentage of annuitants fell considerably in 2004, thereby providing the first evidence that the policy change had an effect on cash-out behavior.

Table 1

The average stock of capital in the super-mandatory part is much higher in 2003 prior to the policy change. This jump in super-mandatory capital is accompanied by a greater number of retirees, a slightly lower than average retirement age, an upward jump in pre-retirement wage and a higher annuitization rate. Taken together, the numbers indicate that the policy change may have triggered not only a higher cash-out rate, but also a shift in retirement to take advantage of the advantageous conditions prior to the change for those most affected by the reform.

Figure 1 below provides additional evidence for potential anticipation effects. Not less than 26 percent of all retirees in 2003 retired in December 2003, and December 2003 is the only month, in which the annuity is chosen more often than the lump sum. Individuals thus seem to have been aware of the policy change and possibly adjusted their behavior accordingly. Not taking this *anticipation effect* into account could potentially cause an upward bias of the impact of the policy change on cash-out behavior.

[12]It should be noted that the data is not representative for the Swiss second pillar. The individuals in the sample are on average poorer as measured by their accumulated pension wealth, retire approximately 2 years later and choose to cash-out much more often. These statistics mirror the fact that insurance companies in general provide second pillar plans for smaller firms with a less educated workforce. Unlike many other countries, well educated and high income workers tend to retire at an earlier age than their less educated and lower paid colleagues. This most likely is a consequence of very high replacement rates in the second pillar for high income workers (Bütler et al. (2004)).

Figure 1

3.3 The environment of the policy change

An important concern in our setup is that other institutional and macroeconomic factors relevant for the annuitization decision might have changed about the same time as the change in policy. The insurance companies in our sample did not report any other changes to retirement plans for men during the period. Neither did the occupational pension law for men change.

As for the macroeconomic environment, things are more complicated. Unfortunately, data from other insurance companies that did not reduce their conversion rates and could potentially be used as a control group to account for changes in the economic environment are not available. The implied conversion factor of the most popular (two-life) annuities offered in the very small unregulated market stayed practically constant over the entire period. However, these annuities constitute a poor benchmark, due to the very thin market in which they operate. Under these circumstances, our strategy is to control for all macroeconomic factors that are relevant for the annuitization decision. In the final analysis interest rates are the only macroeconomic variables we kept. Other macroeconomic indicators such as the unemployment rate varied very little over the period of analysis or were not available at monthly frequencies.

If people expect future interest rates to rise, (risk less) investment opportunities yield a higher payoff and hence the lump sum option becomes more attractive.[13] To capture changes in the yield curve we compute the mortality adjusted present value of an annual income stream for a 65 year old single man until the end of his life for each month in our dataset. Multiplied by the conversion factor, this measure — denoted by PV(income) in the tables — yields the money's worth ratio of a constant annuity for a single men.[14]

[13] Milevsky (1998) argues that one of the reasons for the weak demand for voluntary annuities is the individual draw towards investing in risky assets in order to receive higher expected returns.

[14] We use nominal yields on Treasury bonds with maturities of 1, 2, 3, 4, 5, 7, 8, 10, 20 and 30 years reported by the Swiss National Bank (SNB (2007)) to calculate the expected nominal short rate in each future period. Data on mortality rates are based on the mortality tables created by the Swiss Federal Statistical Office (Kohli (2005)) over the period 1998-2003. We have also experimented with other summary measures for potential investment yields. However, the estimation results were not sensitive as to the yield measure employed.

Figure 2 presents the conversion rate at retirement age 65 (right scale), the conversion rate at age 65 adjusted for the yield-curve in the super-mandatory part (right scale) together with the monthly share of individuals taking a lump sum (left scale). The conversion rate in the super-mandatory part fell from 7.2 percent in December 2003 to 5.845 percent in January 2004. Over the same period the fraction of individuals taking the full or partial lump sum increased from 46.2 percent to 84.7 percent. From mid-2004 to mid-2005 interest rates fell by almost one percent on average, representing an increase in the yield-adjusted conversion rate. The recovery in the annuitization rate as shown in Figure 2 and also evident from Table 1 (over the same period the percentage of annuitants rose by around 8 percentage points) can be seen as informal evidence for the relevance of interest rates for the annuitization decision.

Figure 2

4 Theory and empirical strategy

The (money) value of an annuity is determined essentially by the conversion rate and the relevant interest rates at the time of retirement. The 19 percent reduction in the conversion rate of the super-mandatory part is thus accompanied by an equal net present value loss of the annuity relative to the lump sum. For the average male retiree at age 65, the decrease in the net present value of the benefit stream amounted to approximately 20'000 SFR (≈$ 20'000). Before the policy change a conversion rate of 7.2 percent would have generated a yearly annuity payment of 7'920 SFR in the super-mandatory part; afterwards the applicable conversion rate of 5.835 percent only yielded a yearly pension of 6'420 SFR.

Individuals with a higher capital stock in the super-mandatory part are more affected by the policy change in absolute terms. Therefore, we can expect the impact of the policy change on the cash-out behavior to be strongest among those individuals. We should observe no reaction to the policy change among individuals with zero or negligible amounts in the super-mandatory part of the system.

Based on the results of previous studies, we also expect a higher probability to cash-out retirement wealth for those with low capital stocks. The main reason for that behavior lies in the availability of social assistance in the form of supplemental benefits. This is true because an annuity, even a small one, is detrimental

to the eligibility for income support.

Prior to its adoption, the policy change had been subject to extensive discussions in the media. As a consequence, the policy change and its consequences became public knowledge and we can expect that some people may have chosen to retire earlier to benefit from the higher conversion rate. This is particularly true for individuals that had planned to retire not too long after the policy change and wanted to annuitize their accumulated assets at the pre-reform conversion rate. Ignoring the anticipation effect will lead to biased estimates of the effects of the policy change; strategies to deal with anticipation effects are outlined below.

4.1 Before and after comparison

A straightforward way to analyze the effects of the policy change is to compare the behavior before with the behavior after the policy change. This comparison is done estimating a regression in the form

$$LU_i = f(\alpha + \beta AFTER_i + x_i'\gamma + \epsilon_i) \tag{1}$$

where LU_i is a dummy, which takes the value of 1 if the individual chooses the lump sum or a mixed option and 0 in the case of an annuity.[15] $AFTER_i$ indicates whether the individual was affected by the policy change and x_i is a set of covariates for person i (accumulated total capital stock in 100'000 SFR, square of accumulated capital stock, last wage, and age dummies). To account for changes in the economic environment (in particular, the fact that alternative investment opportunities potentially influence the annuitization decision), we include the summary measure for bond yields at the time of retirement, PV(income). In our regression framework, the coefficient of interest β therefore measures the effect of the policy change on the cash-out behavior of men in our sample.

In the tables we report the coefficients of a linear probability model. We also run regressions with probit and two-sided Tobit models. The resulting effects from these latter models are very close to the coefficients estimated with a simple linear probability model.

[15]The reasoning behind this strategy is that in the case of a mixed option the insurance companies pay out the annuity from the mandatory capital and the lump sum on the remaining capital. Therefore, choosing a mixed option is equivalent to choosing a lump sum in the supermandatory part.

4.2 Anticipation effects

Potential anticipators are individuals who will turn 65 shortly after the policy change, and have some of their pension wealth in the super-mandatory part of the occupational pension plan. Anticipating retirement comes at the cost of losing additional contribution months and years that would have led to both a higher amount of capital stock and a higher annuity. Taken together and neglecting the disutility of labor, both the utility value of the lump sum and annuity are increasing with the age of retirement, as depicted schematically in Figure 3. The decrease in the conversion rate leads to a discontinuous fall in the utility value of the annuity, but not of the lump sum.

The individual foresees an earlier retirement if the utility value of the annuity just before the policy change exceeds the utility value of the chosen payout option at the statutory retirement age. A differentiation can be made between three cases, of which two are presented in Figure 3. First, individuals whose valuation of the annuity before the policy change does not significantly exceed that of the lump and who would cash out after the policy change (see upper panel in Figure 3) These individuals can be expected to retire earlier provided they are close to the statutory retirement age: The higher the pre-reform ratio between the value of the annuity and the lump sum, the more likely is an earlier retirement. The second case is similar to the first, but deals with individuals who would have annuitized even after the loss in the annuity's value (see lower panel in Figure 3). Again these individuals are more likely to retire earlier. The third case concerns individuals who would have chosen the lump sum even before the change and are thus clearly not susceptible to an anticipation effect.

Figure 3

4.2.1 Ad hoc correction of anticipation effects

The first strategy to deal with potential anticipation effects is to use an ad-hoc correction of potential anticipators' behavior: Instead of taking their chosen retirement date (which corresponds to the one or two months immediately preceding the policy) for the year-by-year comparisons, we make them retire either at the statutory retirement age of 65 or in the year 2004 (the year after the policy change). This procedure should provide us with reasonable bounds as to the effect of the policy change. A drawback of this procedure is that although *potential*

anticipators can easily be identified — individuals retiring early with a positive amount in the super-mandatory part —, some of those who retired early in late 2003 might have done so even in the absence of the conversion rate change.

To illustrate the implications of the proposed strategy let LU_0 and LU_1 be the fraction of individuals taking the lump sum under the retirement conditions before and after the policy change. Without any loss of generality we assume that the number of retirees in the absence of anticipation is equal in the two periods and normalized to 1. Had the reform come as a surprise, we could simply compute $\Delta_{\text{true}} = LU_1 - LU_0$ to quantify the impact of the reform on the annuitization decision. As outlined above, however, there are two groups of people likely to retire earlier to benefit from a higher annuity value before the policy change (see also Figure 3): Let AA and AL denote the anticipators (measured as a fraction of all retirees after the policy change) who would have chosen the annuity and the lump sum, respectively, after the policy change. If anticipation occurs, we therefore observe

$$\Delta_{\text{bias}} = \frac{LU_1 - AL}{1 - AL - AA} - \frac{LU_0}{1 + AL + AA},$$

which clearly shows that this measure is biased.

If we were able to correctly identify anticipators ($AA + AL$), but unable to distinguish between the two groups, we could compute upper and lower bounds for the policy effect as follows: An upper bound for the impact can be found by letting all anticipators choose the lump sum after the policy change, $\overline{\Delta} = (LU_1 + AA) - LU_0 > \Delta_{\text{true}}$, a lower bound by letting the anticipators choose the annuity, $\underline{\Delta} = (LU_1 - AL) - LU_0 < \Delta_{\text{true}}$. We define an upper bound as the greater impact of a change in policy (choosing a lump sum instead of a reduced annuity) , and a lower bound of impact as no reaction to the change in policy (allowing the anticipators to be satisfied with the reduced annuity). The higher the fraction of anticipators, the wider the bounds will be. The same strategy would also deliver bounds for other year-by-year comparisons of annuitization rates.

Unfortunately, this ad-hoc strategy is susceptible to individuals who are falsely identified as anticipators. It can be shown that if there are not too many false anticipators, the bounds still contain the true effect,[16] but the gap between the bounds widens and the strategy delivers worse bounds for other year-by-year comparisons.

[16]Taking into account false anticipators F reduces the upper bound to $\overline{\Delta_F} = \frac{(LU_1 + AA + F)}{1+F} - \frac{LS_0}{1-F}$, and the lower bound to $\underline{\Delta_F} = \frac{(LS_1 - AL)}{1+F} - \frac{LS_0}{1-F}$.

4.2.2 Proxy for the likelihood to anticipate

Ideally, if we knew who anticipated (retired earlier) as a reaction to the future decrease in the annuity's value, we would simply add the information a as an additional regressor in the estimation. But, unfortunately, this information is unavailable because not all of those individuals retiring earlier just before the change do so because of the policy change. The second strategy is to construct a suitable proxy z for the *unobserved* anticipation and earlier retirement decision a. Such a proxy has to satisfy two conditions:[17]

1. z must be *redundant*, i.e., in a conditional sense z is irrelevant in explaining the choice of the pay-out option, once anticipation a and other covariates x have been controlled for.

2. z's relation to the anticipation and earlier retirement decision a should be close enough so that once z is included in the estimation equation, the covariates x are not partially correlated with a.

From the graphical analysis drawn in Figure 3 it follows immediately that the cost of not anticipating and retiring earlier increases with the age of potential anticipators at the time of the policy change. The cost is 0 for people turning 65 before the policy change, jumps to a maximum for people who turn 65 just after the change and then decreases with the increasing time span between the policy change and the date of the statutory retirement age of a potential anticipator. The individual has to balance the advantages of anticipating the policy change and retiring earlier (thus benefiting from a higher pre-change conversion rate) against the cost of extra contribution years that are forgone when retiring early. Following the visual presentation in Figure 4, we take as a base proxy a maximal cost (assumed to be 10) minus the number of years between the policy change and the date of retirement at the statutory age 65, i.e., a maximal cost minus the number of years for which additional contributions are forfeited for those who are affected by the reform (0 otherwise).

The chosen proxy satisfies the first requirement. Given the value of the annuity relative to the lump sum and the age of the individual, the time span between the date of the policy change and the individual's 65th birthday is redundant for the choice of the option. For the second requirement — the correlation between

[17] See Wooldridge (2002) (Section 4.3.) for details concerning the properties of the estimators when using proxies for omitted variables.

the omitted anticipated earlier retirement decision a and each of the covariates x proxy is zero once we partial out for z — it can be argued that all true anticipators ($a = 1$) are picked up by the proxy variable. However, anticipated early retirement is not only driven by the length of the anticipated early retirement period, but also by the magnitude of the anticipated benefit per period. The latter is directly related to the size of the capital stock in the super-mandatory part. Taken alone, the proxy z as depicted in Figure 4 might not be fully informative.

To capture the impact of the super-mandatory capital stock on the anticipated earlier retirement decision, we also interact the proxy with the size of the capital stock in the super-mandatory part. To account for the fact that the utility value of the annuity does not increase linearly with the super-mandatory capital, we also include its square interacted with the proxy. Note that if the base proxy does not have a zero mean in the population (which is obviously not the case here), OLS does not consistently estimate the size of the policy impact. As we do not know its population mean, we demean the proxy variable before interacting it with the super-mandatory capital stock. Once interaction effects with the capital stock are included, we would expect the base proxy to no longer have an impact on the choice between the lump sum and the annuity.

Figure 4

4.3 The responsiveness to changes in the annuity's value

The regression approach taken above allows us to clearly identify the effects of the policy change, as long as there are no unobservable factors that are correlated with the cash-out behavior and that may have changed with the policy change. However, this has two shortcomings. First, it makes it impossible to measure the elasticity of the annuitization decision with respect to the change in the annuity value, since we have measured only the numerator (the change in the cash-out behavior) but not the denominator (the change in the value of the annuity). The second drawback is that the value of the annuity depends in a nonlinear way on the price of the annuity, the individual's retirement age and his/her accumulated retirement capital, especially in the presence of first-pillar annuities.

To address these points, we calculate the annuity equivalent wealth (AEW) for each individual in our data set. The AEW is a utility-based measure of annuities that has been used in previous studies by Friedman and Warshawsky (1988), Mitchell et al. (1999), Brown (2001), and Mitchell and McCarthy (2002).

Using this measure we then estimate regression models of the form:

$$LU_i = f(\alpha + \beta \text{AEW}_i + \gamma year_t + x_i'\delta + e_i) \tag{2}$$

where AEW_i is equal to the annuity equivalent wealth of individual i, $year_t$ is a set of retirement year dummies, and x_i consists of a set of other covariates, such as retirement age, wage and a measure for bond yields in the month of retirement. The coefficient of interest is β, which can be interpreted as the responsiveness of the demand for the annuity to changes in the annuity's value.

5 Results

5.1 Before-after comparisons

Simple year by year comparisons are almost certainly plagued by anticipation decisions. We therefore first present some results estimating equation (1) for different quartiles of the wealth distribution separately to account for their different exposures to the policy change. Recall that individuals with no or very little capital in the super-mandatory part should hardly be affected by the change in policy. The higher the capital stock in the super-mandatory part, the greater we can expect the impact to be. Table 2 clearly confirms these conjectures. The change in the cash-out behavior is strongest among those at the top of the accumulated super-mandatory capital distribution. Among the wealthiest quartile, the share of men taking a lump sum increases by not less than 35 percentage points from 2003 to 2004. The increase is only 13 percentage points for the second quartile and drops to around seven percent for the lowest quartile. The share of those who choose the lump sum decreases by roughly 8.5 and 14 percentage points from 2002 to 2003, in the third and fourth quartile, respectively, but does not change in a significant way for the lower half of the capital distribution. Taken together, these two comparisons may indicate the presence of anticipation.[18]

All quartiles show a decrease in the cash-out rate by approximately 5 to 10 percentage points from 2004 to 2005. During that period interest rates fell by almost one percent, increasing the relative attractiveness of the annuity from both

[18]Even without controlling for anticipation, a higher second pillar capital stock significantly increases the probability of full annuitization (not shown here, but see Table 4). These findings are in line with previous empirical studies by Hurd et al. (1998), Brown (2001), and Bütler and Teppa (2007).

the mandatory and super-mandatory part alike. It is thus not surprising, that there is little difference in the response of people in different super-mandatory capital quartiles. Changes in interest rates as well as a changing composition of retirees in the years compared in Table 2, not least as a result of people advancing their retirement to avoid the lower conversion rate, call for an inclusion of additional regressors to isolate the effect of the policy change.

5.2 Correcting for anticipation

As outlined in the previous section we adopt two strategies in dealing with anticipated earlier retirement effects. The first is an ad-hoc correction in which we treat potential anticipators as if they retired at their statutory retirement age or simply one year after the change. The results are shown in Table 3, columns (A1) to (A6). The outcomes of the proxy variable OLS estimations are shown in Table 4. All specifications in Tables 3 and 4 include retirement age dummies, capital, the capital squared, the last wage and a summary measure for interest rates.

5.2.1 Ad-hoc correction

Table 3 shows that the obtained bands for the true effect are greater when potential anticipators are treated as if they retired the year after the policy change (columns (A3) and (A4)) and when we also include November 2003 early retirees as potential anticipators (columns (A5) and (A6)).[19] Although prima facie there is little reason to retire in November 2003 (instead of December 2003), the uncertainty accompanying the policy change might have induced some individuals to choose the earlier exit out of the labor market despite the loss of annuity income. Including November early retirees however renders the strategy more susceptible to falsely identified anticipators. This is confirmed by a lower upper bound in the 2003/4 comparison and a higher upper bound in the 2002/4 comparison. Moreover, columns (A5) and (A6) show a significant increase in the lump sum withdrawal from 2002 to 2003 (contrary to a fall in the cash-out rate

[19]Moving potential anticipators to the year 2004 probably delivers less reliable estimates for the lower bound of the policy effect. Once an individual has decided against anticipating retirement to benefit from the higher conversion rate, it is unlikely she/he chooses to retire early *and take the annuity before the statutory retirement age. The move to 2004 is included as a robustness check*

without controlling for anticipation), potentially showing an over-correction of the anticipation effect.

Recall that if we assume people stick to the annuity when they can not anticipate the change in policy, the effect of the policy change on the annuitization decision is eased (this gives us a lower bound of the effect of the policy change). On the other hand, if people instead cash-out their retirement balances, the effects of the policy change on cash-out behavior is amplified (this is the upper bound). If individuals would otherwise retire at age 65 and still take the annuity (columns (A2) and (A6)), the lower bound for the policy change is found to be small or even insignificant. However, if we assume that individuals switch to a lump sum, the effect of the policy change remains significant and relatively large. This is also confirmed by the bounds for the impact of the policy change comparing the years 2002 and 2004 (see in particular the estimates in which anticipators are shifted to age 65).

All in all, the estimates shown in Table 3 suggest a sizeable effect of the policy change, about 5 to 10 percentage, points and a relatively constant cash-out behavior in the periods preceding and following the policy change. Then again in some cases the bounds are very wide. This is a consequence of both the large number of anticipators and falsely identified anticipators. Not all retirees in December (or November) 2003 who take an annuity and retire before the statutory retirement age have retired earlier as a consequence of the policy change.

5.2.2 Proxy for anticipation decision

Table 4 reports the estimation results from a linear probability model with and without a proxy correction for anticipation for the entire duration of our data. We also experimented with more sophisticated empirical models instead of the linear probability model. The marginal effects of the policy change derived from differ little from the linear probability model.

Table 3, column (A0) displays the results of estimating equation (1) for different combinations of years without controlling for anticipation. These estimates show a decrease in the cash-out rate prior to the change (represented by a positive coefficient before the dummy year 2002), a highly significant 12.9 percentage point increase in the probability to take the lump sum as a consequence of the policy change and a mild recovery of the annuitization rate one year after the policy was implemented. As outlined before, these estimates overstate the true

effect of the policy change due to the option to retire early.

The results of a simple proxy-variable OLS estimation (Table 4, estimation P1) do not differ significantly from the more sophisticated proxy estimation in which the time cost is interacted with the amount of capital in the super-mandatory part of the second pillar and its square (Table 4, estimation P2). In the latter, the impact of the simple proxy vanishes as expected, as the cost of non-anticipation is approximately proportional to the length of the anticipation period multiplied by the magnitude of the benefit.

The proxy variable OLS estimations show a robust increase in the cash-out rate by nearly 8 percentage points as a consequence of the policy change, regardless of whether the simple or the interacted proxy is employed. The estimated coefficient lies within the bounds derived by the simple ad-hoc strategy (Table 3) if only early retirees in December are considered potential anticipators. The effect exceeds the upper bound if the ad-hoc procedure also includes early retirees from November, but might also be the consequence of the latter strategy's higher exposure to falsely identified anticipators.

Both the period preceding the policy change and the period following the change do not display any significant tendency in the probability to annuitize retirement wealth, once the impact of the interest rate on the money value of an annuity have been taken into account (captured by variable PV(income)). Again this finding is consistent with the results of the ad-hoc strategy.

Other covariates in our estimations show the expected sign and magnitude: The probability to annuitize is always lowest for individuals with a low capital stock (presumably as a consequence of the availability of means-tested income support), then increases and reaches a maximum at around 700'000 Swiss francs and declines with even higher capital stocks. Holding constant the capital stock, higher income men are more likely to take the lump sum. A possible explanation for this finding is that individuals with greater income (and thus higher productivity) are more likely to take up another job after retirement, which would lessen their demand for an annuity, not least for tax reasons.

The summary measure for interest rates, PVincome, has the expected negative sign, once the possibility to retire early is taken into account. An increase in the mortality adjusted present value of an income stream of 1 for a 65 year old single man increases the annuitization probability by approximately five percentage points. Given that the downward shift in the yield curve from 2004 to 2005 led to an increase in PVincome by approximately 0.8, a sizeable fraction of the

recovery in annuitization rates from 2004 to 2005 can be explained by the increase in the annuity's value.

In the uncorrected estimation, men who retire at age 64 have a much higher probability to take the annuity than those who retire at age 65. A possible interpretation for this finding is that early retirement (to take advantage of the higher conversion rate) is particularly interesting for men who are close to age 65 at the time of the policy change. The second and third columns of Table 4, (P1) and (P2), which take into account anticipation, confirm this inference. The likelihood to cash-out retirement balances is higher for individuals under the age of 63. A possible interpretation for this finding is that these retirees have liquidity constraints and take the lump sum to bridge the time until the earliest first-pillar benefits are available (from age 63 at an actuarially fair reduction of the benefit).[20] Cash-out rates are still somewhat higher for 63 and 64 year old retirees, but the difference between 63 and 65 year old behavior is considerably smaller and barely significant for the difference between 64 and 65 year old behavior.

5.3 Parameterized model

To estimate the responsiveness of the annuitization decision with respect to the change in the annuity value, we regress the variable Option Lump Sum on the *AEW* (for different levels of risk aversion), a series of year dummies and other covariates. The results for three coefficients of relative risk aversion are shown in Table 5. The first column (estimation V0) presents the baseline case. The estimate implies that a one percentage point increase in the *AEW* increases the annuitization rate by 0.86 to 1.12 percentage points, depending on the level of risk aversion. The policy still has a significant positive effect on the annuitization rate, but the coefficient is less than half of that compared to specification without *AEW*.

However, provided that *AEW* and the Option Lump Sum are certainly correlated with the unobserved anticipation decision, one concern is that these values are upward-biased. In fact, some retirees can choose to increase their AEW by withdrawing from the work force before the end of 2003. To correct for anticipation effects, we present the results of the proxy-variable OLS estimation (second column of Table 5, estimation V1) . Including a proxy-variable reduces

[20]Recall that the individuals in our sample usually receive a higher benefit from the first pillar than from the second.

the estimated impact of the AEW. The implied responsiveness of the annuitization decision with respect to the change in the annuity value is between 0.8 (CRRA=4) and 0.99 (CRRA=0). Furthermore, the effect of the policy change on the cash-out behavior becomes insignificant, consistent with the idea that the time cost is good proxy for the anticipation decision: The AEW completely picks up the effect of the policy change, once the self-selection into early retirement is fully accounted for.

Finally, column 3 (estimation V2) reports the results of the OLS regression in which the proxy is interacted with the total capital stock and its square. The effect of the *AEW* is somewhat smaller than in column 2, but still very sizeable. For both the risk neutrality (CRRA=0) case and medium risk aversion (CRRA=2) a one percentage point increase in the AEW results in a 0.88 percentage point increase in the annuitization rate. For the higher level of risk aversion (CRRA=4) is amounts to a 0.78 percentage point increase. These estimated values are close to previous studies despite the fact that these use completely different sources of variation in the AEW, namely differences across company pension plans and individual characteristics: They are somewhat lower than the one percent increase in the probabability to annuitize for a one percent increase in the AEW as reported by Brown (2001) (men and women pooled), but higher than the 0.44 percent estimate for men only reported in Bütler and Teppa (2007).

6 Conclusions

To the best of our knowledge, our paper is the first to analyze the effects of a truly exogenous variation in the annuity price on the cash-out behavior at retirement. The annuitization decision is analyzed by exploiting a recent policy change in which several Swiss insurance companies reduced the conversion rate at which the retirement capital in the super-mandatory part of the second pillar is translated into a life-long annuity by almost 20 percent. Despite the fact that the administrative data made available by Swiss insurance companies contain no information on non-pension wealth and limited individual information, they offer many advantages over existing empirical studies. The data set is made up of real rather than planned annuitization decisions. Individual decisions involve very large amounts of money (approximately $ 240'000 on average) and some individuals had to make retirement choices subject to great differences in prices. We therefore believe that individuals spend more time in the decision-making

process than answering a questionnaire on hypothetical choices.

The empirical results highlight the importance of the annuity price for the annuitization decision. As a consequence of the policy change the fraction of individuals choosing a lump sum increases by approximately 8 percentage points. This estimate already takes into account that some individuals do not only change their cash-out behavior as a reaction to the change in policy, but also retire earlier without taking the policy change into account. The two strategies to control for the anticipation effect, an ad-hoc correction of shifting potential anticipators to their statutory retirement age and an appropriate proxy for the decision to anticipate retirement deliver consistent results. Once anticipation is properly taken taken into account, we also find constant cash-out rates in the periods before and after the reform.

An annuity provides insurance against the risk of outliving one's assets in old age. To accurately measure the impact of the annuity's value, we compute the annuity equivalent wealth (AEW) within a life-cycle framework previously applied by Brown and Poterba (2000) and Brown (2001). We find that a one percentage point increase in the AEW raises the probability to annuitize by around 0.8 percentage points. This responsiveness is slightly lower than in previous empirical studies by Brown (2001) (for men and women) and somewhat higher than the corresponding estimate for men only in Bütler and Teppa (2007). Nonetheless, given that the variation in the annuity's value differs so much between these studies, the effect of the annuity's value on the decision (not) to annuitize seems remarkably robust across different countries, pension plan details and the manner in which the choice is elicited.

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A Annuity Equivalent Wealth

We use the methodology of Brown and Poterba (2000) to calculate the AEW, but adapt it to the Swiss pension system. The AEW approach calculates the additional wealth an individual would need to be equally well off without actuarially fair annuities in comparison to with actuarially fair annuities. According to Brown (2001), four main factors are essential to compute the AEW: risk aversion, fraction of pre-annuitized total wealth, mortality risk, and marital status. Unfortunately, we have no information on the marital status of the individuals in our data set. In the calculation we are assuming that all individuals are married. This is a reasonable assumption given that in 2003 76 percent of all Swiss men in the age group 40-64 were married. Moreover, there is no reason to believe that the composition of our sample with respect to the marital status should differ systematically in 2003 and in 2004, suggesting that the error in the calculation of the AEW is similar in both years.

In Brown and Poterba (2000), the couple chooses total consumption, C_t^c, optimally to maximize an additively separable lifetime utility function. The total consumption of the couple consists of a weighted combination (with parameter δ) of the husband's consumption (superscript m) and that of the wife's (f). The utility function exhibits constant relative risk aversion and is a weighted combination (with parameter ψ) of the utility function of the husband, $u^m(\cdot)$, and that of the wife's, $u^f(\cdot)$. We compute the AEW with three different coefficients of relative risk aversion equal to 0, 2 and 4. Since the insurance component of annuities against the risk of longevity is more pronounced the more risk-averse individuals are, the AEW is increasing in the coefficient of relative risk aversion.

Formally, the Bellman equation (only shown for the state in which both members of the couple are still alive) can be stated as:

$$\begin{aligned}
\max_{C_t^m, C_t^f} V_t^c(W_t) \quad = \quad & \max_{C_t^m, C_t^f} u^m(C_t^m + \delta C_t^f) + \psi u^f(C_t^f + \delta C_t^m) + && (3)\\
& + \frac{1}{1+\rho} \Pr[m \text{ alive}, \ f \text{ dead}](t+1) V_{t+1}^m(W_{t+1}) + && (4)\\
& + \frac{1}{1+\rho} \Pr[m \text{ dead}, \ f \text{ alive}](t+1) V_{t+1}^f(W_{t+1}) + && (5)\\
& + \frac{1}{1+\rho} \Pr[m \text{ alive}, \ f \text{ alive}](t+1) V_{t+1}^c(W_{t+1}) && (6)
\end{aligned}$$

subject to the following constraints:

$$\begin{aligned} &\text{i)} && W_0 \text{ given} \\ &\text{ii)} && W_t \geq 0, \forall\ t \\ &\text{iii)} && W_{t+1} = (W_t - C_t^c + S_t + B_t)(1+r), \end{aligned} \tag{7}$$

where V_{t+1}^m and V_{t+1}^f represent the value functions for the states in which the male or the female, respectively, is the surviving spouse. Here, $\Pr[\cdot](t+1)$ are the conditional transition probabilities to the different states from period t to period $(t+1)$. As we do not have any information on non-pension wealth, we assume that the initial wealth, W_0, is equivalent to the accumulated second pillar pension wealth in the mandatory and super-mandatory parts $K = K_{mand} + K_{sup}$. S_t denotes first pillar benefits, which are paid out from the time of retirement, adjusted in an actuarially fair manner in the case of early retirement. B_t represents the sum of occupational benefits if any from the mandatory ($B_{t,mand} = \gamma_{mand} K_{mand}$) and super-mandatory part ($B_{t,sup} = \gamma_{sup} K_{sup}$). We use a constant annual interest rate of 3 percent and a rate of time preference of $\rho = 0.03$.

To compute the AEW we compare two scenarios: In the *full annuitization* case the couple's eligible person annuitizes the entire pension wealth K and reaches a maximum utility of V^*. The initial capital W_0 is equal to zero. In exchange for K, (s)he gets a lifelong nominal annuity, $B_t = B_{t,mand} + B_{t,sup}$. If the couple's main Claimant should die, the surviving spouse gets a reduced lifelong annuity, $B_t = \lambda(B_{t,mand} + B_{t,sup})$ with $\lambda = 0.6$. In the *full lump sum* scenario, the entire capital stock is cashed out, leaving the couple without any additional benefit from the second pillar, i.e., $B_t = 0$. The initial capital is equal to the entire pension capital, $W_0 = K$. The corresponding utility is V. We then calculate the amount of additional wealth ΔW the individual must receive in the full lump sum case to reach the same utility level V^* as in the full annuitization scenario.[21]

$$V(K + \Delta W | B_t = 0,\ \forall t) = V^*$$

The resulting AEW is then

$$AEW = \frac{K + \Delta W}{K}$$

[21] As a robustness check we recalculate the AEW assuming that in the full lump sum scenario individuals take a lump sum in the super-mandatory part only and choose an annuity for the mandatory part.

A.1 Imputation of Super-mandatory capital

One difficulty in computing the AEW is given by the fact that we do not observe the mandatory and super-mandatory part separately for individuals that retire before 2004. We rely on mean imputation as well as regression-based imputation to estimate the super-mandatory part for those individuals. In the mean imputation approach we split the accumulated capital stock of individuals retiring after the policy change in bins of 20'000 SFR. For each bin we compute the mandatory and super-mandatory capital as a fraction of the total capital stock. These estimated shares can then be used to divide the total accumulated capital stock of men retiring before the policy change into the mandatory and super-mandatory part. In the regression-based imputation we regress the super-mandatory capital on the retirement age, the total capital stock and the earnings in the year prior to retirement. Since the accumulated capital stock is left-censored, we apply a one-sided Tobit-model.

	period not affected				*reform announced*		*reform implemented*			
	2001		**2002**		**2003**		**2004**		**2005**	
Variable	**Mean**	**S.D.**	**Mean**	**S.D.**	**Mean**	**S.D.**	**Mean**	**S.D.**	**Mean**	**S.D.**
Age at retirement	64.7	1.1	64.5	1.2	64.3	1.3	64.5	1.3	64.5	1.2
Conversion rate										
Mandatory Part	7.136	0.210	7.104	0.231	7.059	0.256	7.071	0.288	7.029	0.280
Supermandatory Part	7.136	0.210	7.104	0.231	7.059	0.256	5.774	0.153	5.780	0.139
Last Wage	76,232	56,462	78,050	70,853	88,673	86,457	79,687	65,699	79,922	65,133
Capital at retirement	238,850	200,431	235,799	189,342	282,198	225,018	228,548	178,182	238,914	181,249
Mandatory Capital	127,306	52,506	128,695	55,417	137,681	58,344	132,047	63,161	134,952	69,028
Super-mandatory Capital	111,515	173,218	107,088	158,011	144,705	193,272	96,380	154,462	103,962	156,075
Annuity	0.353	0.478	0.330	0.470	0.407	0.491	0.251	0.434	0.334	0.472
Lump Sum	0.616	0.487	0.632	0.482	0.548	0.498	0.730	0.444	0.632	0.482
Mixed	0.031	0.173	0.038	0.191	0.045	0.208	0.019	0.138	0.033	0.180
Observations	976		1,104		1,678		1,080		1,017	

Table 1: Summary statistics for men.

Variable	comparison 2004-2003 (W1)	comparison 2004-2002 (W2)	comparison 2005-2004 (W3)	comparison 2003-2002 (W4)
Sample: all men				
LU_i, 0-25 perc.	0.072**	0.041	-0.106**	-0.031
	(0.035)	(0.038)	(0.041)	(0.034)
LU_i, 25-50 perc.	0.057	0.002	-0.053	-0.054
	(0.036)	(0.039)	(0.041)	(0.037)
LU_i, 50-75 perc.	0.131***	0.046	-0.100**	-0.085***
	(0.038)	(0.040)	(0.042)	(0.038)
LU_i, 75-100 perc.	0.352***	0.212***	-0.086**	-0.140***
	(0.037)	(0.041)	(0.042)	(0.037)
cut-off values	2002	2003	2004	2005
min	5,603	5,663	0	0
25th	29,013	31,631	8,896	11,514
50th	42,536	45,506	37,968	47,068
75th	116,945	181,553	116,870	130,675
max	1,224,305	1,278,186	1,282,853	1,172,339

Table 2: Before-after comparisons by super-mandatory retirement capital for men aged 60 and above (no covariates). For 2002 and 2003, super-mandatory retirement capital has been imputed (see text). Significance levels: *** = 1%, ** = 5%, * = 10%.

potential anticipaters	*none*	*retire in dec 2003*		*retire in dec 2003*		*retire in nov/dec 2003*	
strategy: retirement date	as observed	shifted to age 65		shifted to year 2004		shifted to age 65	
strategy: imputed choice	No correction	lump-sum	annuity	lump-sum	annuity	lump-sum	annuity
Variable	(A0)	(A1)	(A2)	(A3)	(A4)	(A5)	(A6)
No. pot. anticipaters	241	241	241	241	241	303	303
No. shifted 2003 → 2004	0	92	92	241	241	115	115
No. shifted 2003 → 2005	0	53	53	0	0	66	66
$\overline{\text{LU}}_{2004} - \overline{\text{LU}}_{2003}$	0.119***	0.091***	0.034*	0.113***	-0.023	0.069***	0.011
	(0.019)	(0.018)	(0.018)	(0.018)	(0.018)	(0.018)	(0.018)
$\overline{\text{LU}}_{2004} - \overline{\text{LU}}_{2002}$	0.099***	0.120***	0.059**	0.140***	-0.011	0.125***	0.052**
	(0.025)	(0.024)	(0.025)	(0.024)	(0.025)	(0.024)	(0.025)
$\overline{\text{LU}}_{2003} - \overline{\text{LU}}_{2002}$	-0.037	0.018	0.018	0.018	0.018	0.059**	0.059**
	(0.024)	(0.025)	(0.025)	(0.025)	(0.025)	(0.024)	(0.024)
$\overline{\text{LU}}_{2005} - \overline{\text{LU}}_{2004}$	-0.020	0.045	-0.162***	-0.020	-0.027	0.054	-0.183***
	(0.043)	(0.035)	(0.036)	(0.042)	(0.045)	(0.033)	(0.035)

Table 3: Non-parametrical correction of anticipation effects: Potential anticipaters (annuitants with postive retirement capital in the super-mandatory part retiring early in December (and November) 2003) are shifted to the statutory retirement age or to the year after the policy change. The lower bound for the effect is found by letting them annuitize even after the policy change, the upper bound by assuming that the anticipaters would have taken the lump sum after the change. Additional controls: retirement capital and its square, summary measure for interest rates, retirement age dummies. Significance levels: *** = 1%, ** = 5%, * = 10%.

Option LS	*uncorrected* Coef.	(Std.)	p	*proxy* Coef.	(Std.)	p	*proxy interacted* Coef.	(Std.)	p
		(P0)			(P1)			(P2)	
Capital (100k)	-.088	(.008)	***	-.086	(.008)	***	-.085	(.008)	***
Capital sq	.005	(.001)	***	.005	(.001)	***	.006	(.001)	***
Wage (100k)	.029	(.020)	***	.029	(.010)	***	.029	(.010)	***
R.A. 60	.033	(.037)		.064	(.037)	*	.060	(.037)	
R.A. 61	.062	(.035)	*	.105	(.034)	***	.106	(.033)	***
R.A. 62	.031	(.029)		.086	(.030)	***	.078	(.029)	***
R.A. 63	-.006	(.025)		.062	(.026)	**	.048	(.025)	*
R.A. 64	-.106	(.024)	***	-.013	(.028)		-.014	(.028)	
R.As. 66-70	YES			YES			YES		
Post2003	.129	(.019)	***	.076	(.020)	***	.077	(.020)	***
Y01	.011	(.027)		-.051	(.029)	*	-.052	(.029)	*
Y02	.043	(.022)	*	-.022	(.024)		-.022	(.024)	
Y05	-.058	(.028)	**	-.037	(.028)		-.038	(.028)	
PV(Income)	-.021	(.024)		-.047	(.024)	**	-.049	(.024)	**
Proxy Anticip				-.026	(.004)	***	-.003	(.005)	
Proxy*Sup65							-.016	(.003)	***
Proxy*Sup65sq							.001	(.000)	***
Annuity max	808k			786k			739k		
R squared	0.048			0.057			0.064		
No. Obs	5677			5677			5677		

Table 4: Linear probability estimates of the lump-sum / annuity decision. Anticipation effects are corrected using a proxy (last two estimations). Significance levels: *** = 1%, ** = 5%, * = 10%.

	Option LS	*uncorrected* Coef. (Std.) p (V0)	*proxy* Coef. (Std.) p (V1)	*proxy interacted* Coef. (Std.) p (V2)
none	Post2003	.128 (.018) ***	.076 (.020) ***	.077 (.020) ***
	Y01	.011 (.027)	-.060 (.029) **	-.062 (.029) *
	Y02	.043 (.022) *	-.023 (.024)	-.024 (.024)
	Y05	-.058 (.028) **	-.037 (.028)	-.037 (.028)
	AEW0	-1.117 (.170) ***	-.993 (.171) ***	-.883 (.172) ***
CRRA=0	Post2003	.049 (.022) **	.011 (.023)	.018 (.023)
	Y01	.008 (.027)	-.056 (.029) *	-.057 (.029) **
	Y02	.040 (.022) *	-.020 (.024)	-.021 (.024)
	Y05	-.071 (.028) **	-.051 (.028) *	-.049 (.028) *
	AEW2	-1.013 (.141) ***	-.925 (.141) ***	-.877 (.142) ***
CRRA=2	Post2003	.050 (.021) **	.010 (.022)	.013 (.022)
	Y01	.010 (.027)	-.055 (.029) *	-.056 (.029) *
	Y02	.041 (.022) *	-.019 (.024)	-.020 (.024)
	Y05	-.072 (.028) **	-.052 (.028) *	-.051 (.028) *
	AEW4	-.860 (.122) ***	-.795 (.122) ***	-.775 (.122) ***
CRRA=4	Post2003	.061 (.021) ***	.018 (.022)	.020 (.022)
	Y01	.011 (.027)	-.055 (.029) *	-.056 (.029) *
	Y02	.043 (.022) *	-.019 (.024)	-.019 (.024)
	Y05	-.071 (.028) **	-.051 (.028) *	-.051 (.028) *

Table 5: Coefficients for measures of Annuity equivalent wealth (AEW with coefficients of relative risk aversion (CRRAs) of 0, 2, and 4) and dummies for policy change and other years in a linear probability model. Dependent variable is Option Lump Sum, other covariates include capital and its square, the individual's last wage and retirement age dummies. Anticipation effects are corrected using a proxy (last two columns, as in Table 4). Significance levels: *** = 1%, ** = 5%, * = 10%.

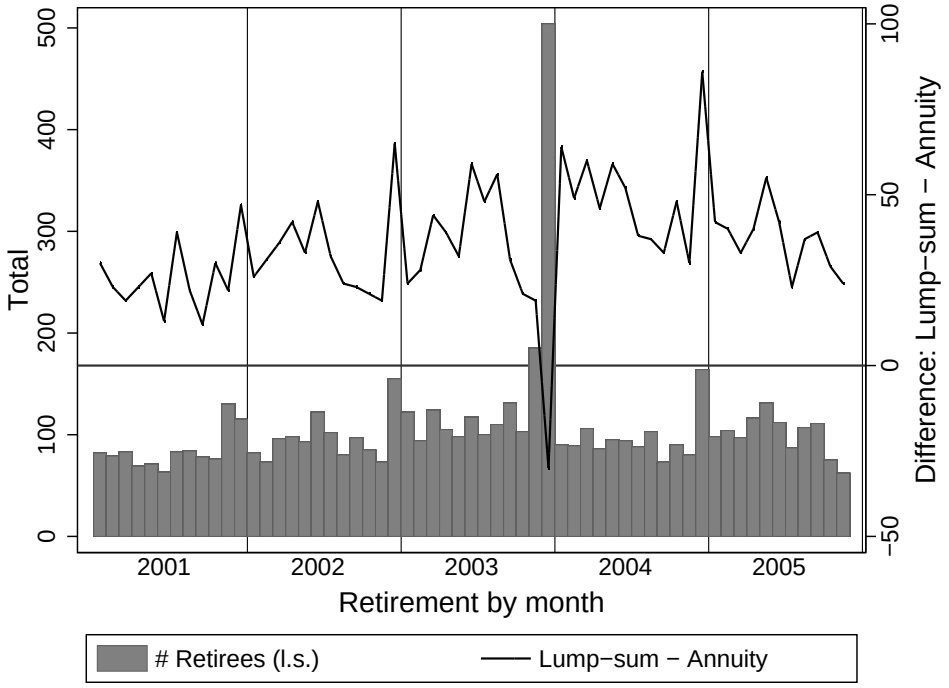

Figure 1: Monthly retirement numbers for men, years 2001-2005 (left scale), and the difference between lump sum and annuity (in percentage points, right scale).

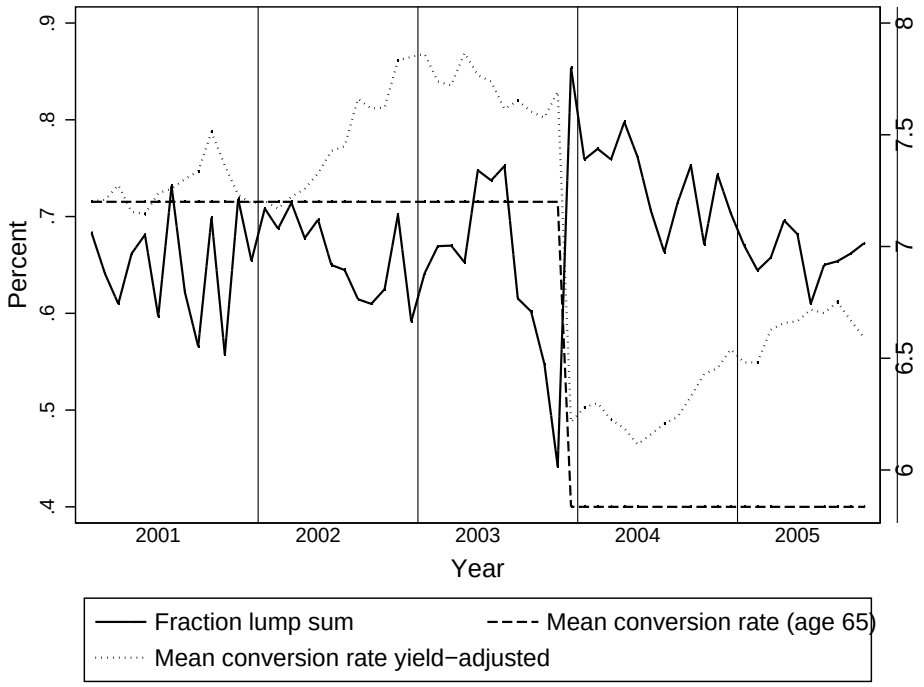

Figure 2: Conversion rate (dashed line, right scale) and yield-adjusted conversion rate (normalized to equal the conversion rate in Jan 2001, dotted line, right scale) both in the super-mandatory part, and the percentage of individuals taking the lump sum (solid line, left scale) over time.

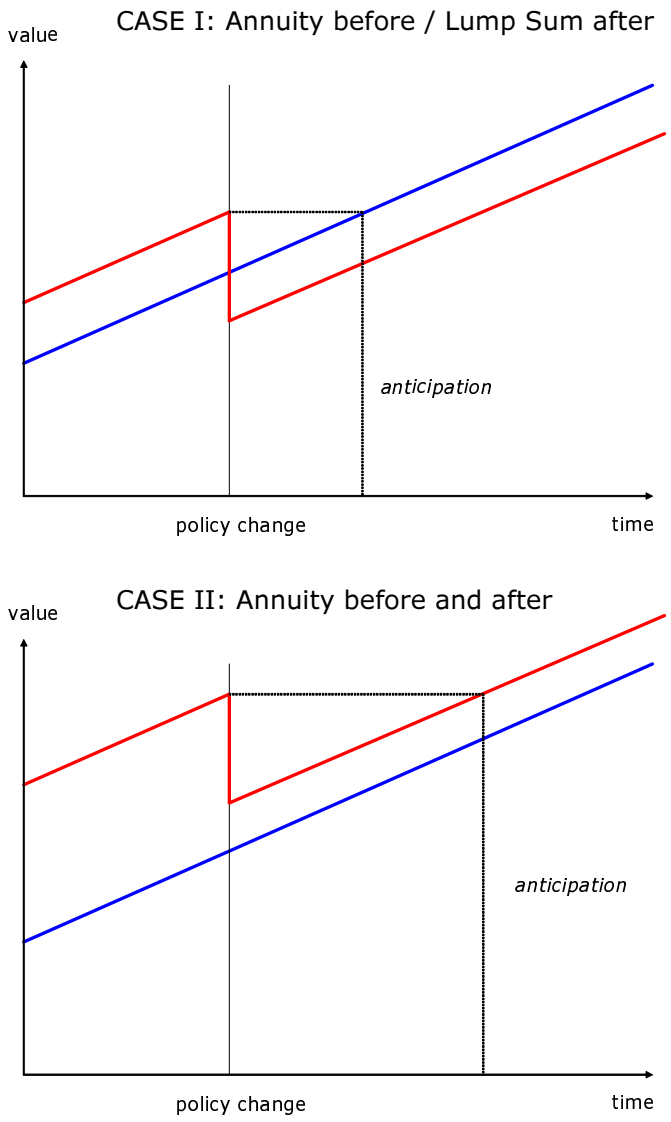

Figure 3: Potential anticipators forseeing a decrease in the annuity's value. The blue line represents the individual's utility valuation of the lump sum, assumed to increase with age due to additional contributions. The red line represents the annuity's value. Early retirement is anticipated if the annuity's value just before the policy change exceeds the annuity's value or the lump sum, respectively, at the statutory retirement age.

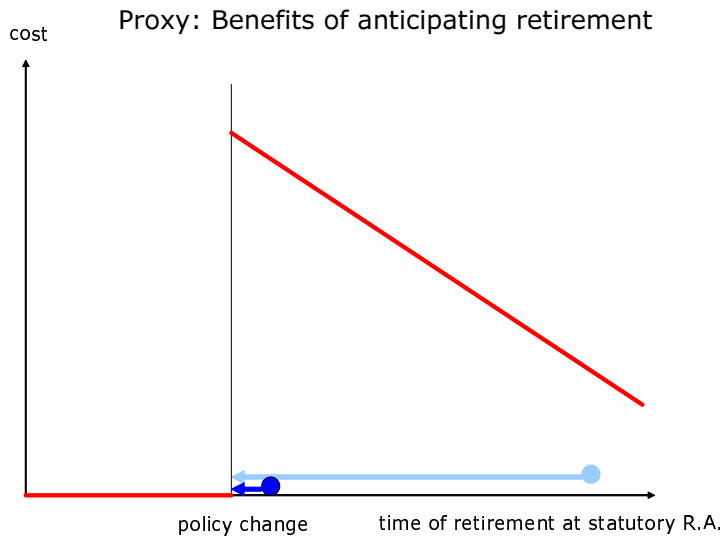

Figure 4: Construction of a proxy c for the likelihood of anticipating early retirement in light of the decrease in the annuity's value. The base proxy c is zero for individuals who reach the statutory retirement age before the reform and is maximal for those who turn 65 just after the policy change.

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