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The Determinants of the Money's Worth of Participation in Collective Pension Schemes

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Abstract

In many countries, employees have implicit or explicit options to opt-out of collective pension schemes. The contributions to such schemes are often set uniformly, irrespective of age, gender, or education level. We quantify the incentives for individuals that participate in such systems. The indexation quality of pension funds introduces additional incentives to opt-out of schemes with inadequate funding.

We show that young highly educated males have a strong incentive to opt-out of the collective system in case of uniform pricing, since their contribution is high relative to the benefit obtained. This incentive is enforced by the fact that the switching costs for young individuals are relatively low. Moreover, it turns out that the indexation quality of the scheme is a non-negligible determinant of the incentives provided to participants.

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1 Introduction

In many countries employees have implicit or explicit options to opt out of collective pension schemes. The option can be to participate in a collective pension scheme or to receive a lump-sum contribution via an individual defined contribution scheme, but it is often also more implicit. Employment with a specific firm might imply mandatory participation in the collective pension scheme of this firm, which can be avoided by switching jobs to another firm, sector or country. In this paper we analyze the economic value or money's worth of the annuity contracts that are typically offered by collective pension schemes. Collective pension schemes are often funded via a uniform contribution, determined as a fraction of the wage earned. Therefore, the premium paid is invariant to the individual characteristics of the employee, like age, gender, and education level. For instance, the economic value of identical annuity contracts is substantially lower for young employees than for employees close to retirement due to the time value of money and a lower likelihood of surviving up to retirement. The money's worth of participation in uniformly priced pension schemes depends on other individual characteristics than age, that determine the survival probabilities of employees. It is well-known that women live longer than men in expectation, and the life expectancy of highly educated groups substantially exceeds that of lower education groups, see e.g. Brown (2003) and Huisman et al. (2004, 2005). This discrepancy between the money's worth and what employees pay introduces incentives that are analyzed in this paper.

The collective schemes considered in this paper can be characterized by obligatory participation, forced annuitization, collective asset allocation decisions and uniform pricing. The schemes can be either defined benefit (DB) or collective defined contribution (DC). The rights in the defined benefit plans purely depend on the labor history of the participant. In collective DC plans the asset returns and future premium rules play an important additional role, and e.g. determine whether or not the rights will be indexed against inflation. Both of the collective plans (either DB or collective DC) generate the same incentives of the participation in the schemes. Occupational earnings related collective pension plans with flat contribution rates as studied in this paper are common in the Netherlands, UK, US, Switzerland and Canada, just to name a few.

Apart from individual heterogeneity, differences in the extent to which inflation protection is provided can also have a very significant impact on the money's worth of participation in a pension scheme. Obviously real annuities are more valuable than nominal ones, but the relative valuation in money terms will depend on the current inflation and interest rates. More importantly, many collective schemes do not offer

straight nominal or real pension rights but target to provide inflation indexation if sufficient funding is available. The extent to which such schemes offer indexation protection will be referred to as the indexation quality of the scheme. In our numerical calculations we will focus on the specific indexation rules that have recently been adopted by many Dutch pension funds. In the schemes offered by these funds indexation of pension rights will typically be only partial if the funding ratio of the fund is insufficient[1]. Moreover, insufficient funding implies that subsequently the pension premium will be increased. As a consequence, employers covered by a pension fund with a currently high funding ratio are offered a more attractive pension scheme than another fund where the rules that determine the entitlements are identical but the funding is worse. Therefore, apart from the impact of individual characteristics, there is heterogeneity at the level of pension funds, which introduces additional incentives. Furthermore, we address the interplay between the individual incentives and the incentives provided by the current financial situation of the pension fund.

Analysis of the welfare effects of a pension scheme is required whenever pension schemes are evaluated or redesigned. Deviations between the cost and the market value of participation in the scheme require solidarity of groups of individuals that is not Pareto improving[2]. If the deviation between costs and ex ante benefits of participation would get too large, the net contributors in a voluntary scheme will not participate and the scheme may become unsustainable. Many studies have analyzed the cost and benefits of life-time participation in specific DB schemes (see e.g. Cui et al., 2005; Gollier, 2005) and considers sustainability of the scheme relative to DC schemes. This paper in contrast focuses on costs and benefits of participation in a collective scheme for a single year.

Even if participation in a DB pension scheme is legally obliged, the net contributors will try to avoid participation and e.g. switch jobs to another firm or industry or to another country for that reason, which also makes the scheme unsustainable. Of course, the cost of switching jobs (including loss of job or sector-specific human capital) can be substantial, and the incentive to leave the fund becomes relevant only if the differential between contribution and economic value exceeds the switching costs.

Participation for one year in a DB scheme generates an annuity payment as of the

[1] A similar indexation policy is used in Switzerland.

[2] This should be distinguished from the risk solidarity that is often imposed by the premium and indexation rules in schemes which are welfare improving if equivalent financial instruments are not available.

retirement age. This paper focuses on the economic value of this annuity.[3] An extensive literature has analyzed the welfare effects of annuities (see for example Brown, 2002, 2003), of holding indexed bonds (see among others Campbell and Viceira, 2001, 2002; Campbell et al., 2003; Brennan and Xia, 2002) and of other investment strategies. These papers assume specific initial assets and decision rules for investors and make a utility comparison. In line with the literature on the money's worth of annuities we restrict ourselves to the case of a fully rational optimizing agent and complete markets. We assume that the agent can and will costlessly unwind the portfolio strategy of the fund as well as of the annuities imposed by the scheme. In this setting the investment strategy of the fund, the precise form of the utility function of the agent or any additional asset holdings that the agent might have, does not have any relevance for the value of participation in the scheme. A comparison of the money's worth of the participation in the scheme and the cost of participation captures all incentives to participate.

It is well known in the literature that for subgroups of the population the money's worth of the annuities that are imposed can differ from the costs charged by uniform contributions. These differences are often referred to cost solidarity between subgroups. For Dutch schemes for instance, Kune (2005) has listed cost solidarities between men and women, between younger and older workers, between singles and couples, between workers and disabled, between low and high educated groups, etc. Some of these solidarities might be intentional and desirable (e.g. the solidarity between workers and disabled persons), others might be non-intentional and undesirable. The aim of this paper is to quantify the solidarities imposed by uniform pricing, not to consider which solidarities would or would not be desirable.

In our analysis we will assume that the costs of switching from one pension scheme to another are small. In reality, switching jobs can be costly and switching between funds can have a significant impact on the accumulated retirement wealth[4]. For instance, in the US and in the UK pension rights that are not yet vested can evaporate. In final wage schemes it can be rather unattractive to leave pension rights with one pension fund and switch to another since the final wage in the first scheme will no longer be adjusted. Insufficient transferability of pension rights to another scheme can imply that transfer of the funds to another scheme is likewise not too attractive. In this paper we assume in contrast that the accumulated retirement wealth is not affected by a change

[3]The product offered by many compulsory pension schemes also contains disability insurance and partner pension. These are not considered in the analysis.

[4]See Cocco and Lopes (2004) for a detailed description of the rules to transfer retirement wealth between different funds in the UK.

of pension schemes. This is at least approximately the case in the Netherlands, where transferability of retirement funds at actuarially fair prices is a legal right. Note however, that the formulas that are currently used in Dutch pension transfers are approximations to the market values that we analyze. Moving between funds with a different indexation quality e.g. generally does have an impact on the value of retirement wealth, because the indexation quality is not taken into account. Note also that for young workers the switching costs are likely to be small in all cases, so that they have the strongest incentives to find optimal pension schemes.

Our comparison of the cost and benefits of participation in a pension scheme is closely related to generational accounting (see Auerbach et al., 1999). Ponds (2003) emphasizes the need to have ex ante fair pricing of the pension contract for each cohort. We extend his analysis to a comparison of the costs and benefits of other subgroups (men versus women and differences in education level) and to differences in indexation quality. Moreover, we analyze the incentives for every cohort year by year rather that we sum them up to an overall number.

Our emphasis in this study is on the determinants of the money's worth of participation in a collective pension scheme. We focus on the implications of the money's worth for possible options to opt out. The money's worth of participation in a pension scheme is also important for a variety of other reasons. First of all, the fair price for new entrants to a scheme (e.g. because they were recently hired by the firm that sponsors the scheme) equals the market value taking their characteristics into account. The analysis also clarifies the incentives for insurers to offer annuity products in specific segments of the population. While very few insurers do so explicitly, this can be done implicitly by focusing the marketing efforts on specific subgroups. The value of participation in a scheme is also required to have transparent labor markets where agents react to the incentives implied by the scheme. The premium to be paid for nontransparent obligatory pension schemes can easily be perceived as taxation rather than a contribution to personal income during retirement, which would imply that the net wage that is offered is underestimated and the labor market is distorted. Finally, the market value of the liabilities to all participants is also an important element in the accounting of the firm as of the introduction of the International Financial Reporting Standards. Market valuation requires that for valuation of the liabilities, the heterogeneity in the population of the pension fund is properly accounted for.

The main results of this paper are the following. Participants in a scheme with primarily older and highly educated workers have strong incentives to opt out of a

uniformly priced collective pension scheme if they have access to annuity markets at risk-based pricing. Assuming the economic conditions of January 2004 the money's worth of one year participation in a nominal scheme for a 25-year-old man is estimated to be 1.5% of his annual salary on average, while the money's worth for a 64-year-old man is 18.7%. The money's worth moreover depends significantly on gender as well as on level of education. For conditionally indexed schemes the money's worth moreover depends on he current funding ratio and on the asset mix of the fund. The money's worth of participation of a 25-year-old man in a fund with funding ratio of 100% and 100% bond investment is 1.7% of his salary, while the value of participation in an identical fund with with a funding ratio of 140% is 3.3%. We moreover show how the money's worth of participation in the scheme depends on the assumptions on improvements in life expectancies.

The set-up of this paper is as follows. In Section 2 we present a review of the extensive literature on differences among groups in the welfare effects and pricing of annuities. In Section 3 we determine the money's worth for different groups of individuals of a year of participation in a nominal or fully indexed DB plan. The results indicate that the economic value of the annuity rights that are obtained can be substantially different, while the cost of participation is typically the same for all. We discuss the drawbacks for the cost solidarity between groups that is imposed by an obligatory collective scheme that is based on uniform pricing. Throughout, we ignore the effect of premium adjustments if the funding ratio of a fund would drop and assume that the individual could avoid this increase by switching to another employer or to third pillar products. In Section 4 we focus on the use of models of the nominal and real term structure similar to the ones proposed by Brennan and Xia (2002) and Campbell and Viceira (2001) to determine the market value of a year of participation in conditionally indexed schemes. Section 5 restates the main conclusions of the analysis.

2 A survey of the literature on money's worth of annuities

An extensive recent literature outlines elements of the optimal individual financial decision making related to retirement. Two important risk factors are longevity risk and inflation. The financial instruments that can be exploited to hedge these risk are annuities, see e.g. Poterba and Wise (1998), and real bonds, see e.g. Campbell and Viceira (2001). Participation in a pension scheme usually does provide coverage against

longevity risk and aims for inflation indexation and will therefore usually have substantial added value for a naive investor.[5]

Life expectancies are different among people, which have a welfare effect on individuals participating in a mandatory pension plan. Brown (2002, 2003) documented unequal expected lifetimes for groups with different characteristics. Women live longer than men, and there are significant differences in life expectancies along racial/ethnic lines. Brown (2003) documented 6 years longer life expectancy for women than for men at the age of 22 in the total US population. However these differences vary along ethnic lines. 22-year-old white men live 6.5 years longer in expectation than black, and the difference is 4 years in favor of white women. Life expectancy varies with education. White men at the age of 22 with college education live 5.2 years longer in expectation than white men with less than high school education. This difference is 7.6 years for black men, 3 years for white women and 4.4 years for black women. The differences mentioned above are still present, but slightly smaller for people with higher attained age. For instance, the expected lifetime is 3.7 years longer for women than for men with the age of 67. White men at the age of 67 have 2.1 years longer life expectancy than white women, etc.

Differences in life expectancies are also present in Europe. Kunst (1997) found the effect of different educational levels on life expectancy in several European countries. Huisman et al. (2004) also documented mortality differences among cohorts with different educational levels in 11 European populations. A recent report by Herten et al. (2002) documents similar findings to Brown (2003) in the Netherlands. On the basis of a social economic survey between 1995 and 1999, women at the age of 20 are expected to live 5.4 years longer than men, while this difference between women and men slightly decreases to 4.7 for people who attained the age of 65. The difference in expected lifetime is present among cohorts with different educational level. A 20-year-old highly educated man is expected to live 5 years longer, than a man at the same age with the lowest education. This difference shrinks to 3.7 years as soon as he reaches the age of 65. A 20-year-old woman with high education lives 2.6 years longer in expectation than a woman with the lowest education, and this difference becomes 2.1 years as a woman gets 65 years old.

Differences in life expectations induce wealth transfers among different cohorts, distinguished along gender, educational level or ethnic lines. If cohorts with different char-

[5]This is not only true for a naive investor, but in many countries the markets for both annuities and inflation-linked securities are underdeveloped. Pension funds therefore complete the market with respect to these two risk factors.

acteristics are pooled and participate in a collective pension or annuity plan that does not take into account cohort-specific differences, people with worse survival prospect subsidize groups with higher expected lifetime. This statement equally holds for pension funds setting premium or for insurance companies selling annuity products.

For instance, Brown (2002, 2003) examined the distributional implications of compulsory annuitization in the US by using the money‘s worth of annuity framework. The money‘s worth measure is the expected present value of annuity payments per dollar spent to purchase the annuity. If annuities are qualified (payment received each month from a qualified annuity is taxable as income) annuity rates are generally unisex, which implies that the monthly annuity payments are constrained to be the same (uniform) for all individuals. Brown (2002, 2003) report the money‘s worth of the uniform annuities for individuals with the age of 67, by taking into account cohort-specific (gender, educational, race) survival characteristics. In expectation men pay 6.6% more and women pay 5.6% less than the present value of the nominal annuity they are expected to receive. Black men pay 12.9%, however white men pay 5.9% more than the fair value of the annuity. Black women pay 1.1%, while white women pay 6.3% less. Moreover, highly (college) educated white men pay only 2% more and low educated white men pay 12.5% more than the market value of their nominal annuity. On the other hand, highly educated white women pay 7.9% less, while low educated white women pay 3% less than the market value of their annuity they are expected to receive. Similar patterns can be observed while looking at the effect of educational differences for black men and women.

Brown (2002, 2003) calculated the money's worth of annuity values for real annuities as well. Cohorts which gain in the nominal plan will also gain with the real annuity. Similarly, the same is true for losses. In the whole population, the losses suffered by men are 8.7%, and the gains for women are 7.1% in real terms. The losses and gains are higher than in nominal plans, and this statement holds for all race- or education-specific cohorts in almost all cases.

Feldstein and Liebman (2002) calculated the net present value of the lifetime participation for different cohorts in the US population in a funded pension system. Participants pay 9% payroll taxes to a personal retirement account (PRA) which earns 5.5% return and the balance is fully annuitized when the individual reaches retirement. PRA annuities are calculated by using a single uniform unisex mortality table; age-, sex-, race-, and education-specific differences are ignored. The results are sensitive to the choice of the discount rate which is used to calculate the net present values, however, most of the conclusions are robust to its size. The net present value of the lifetime contribution is

higher for women than for men, and white people benefit more than black. The results related to differences in the educational level depend on the size of the discount rate. If the discount rate is 1% or 3%, then higher education groups benefit more than cohorts with low education. However, if the discount rate is assumed to be 5%, then cohorts with the highest and lowest education benefit almost the same, however the group with middle level education benefits the most.

Many of the results discussed in this paper can also be applied to pricing annuity products that are offered to individuals. The potential important additional complication there is that of adverse selection. A well-known stylized fact in the annuity literature is that those that choose to buy an annuity have a life expectancy that exceeds that of the population at large (see Mitchell et al., 1999; Finkelstein and Poterba, 2002, 2004). In this paper we ignore potential information asymmetries.

3 The money's worth of participation in collective pension schemes

In this section we consider the economic value of participation in a collective pension fund which offers purely nominal or purely real pension benefits to the participants. The participation in a pension fund is compulsory, i.e. all employees have to participate collectively in a fund. We consider three types of funds. The nominal fund offers guaranteed (DB) nominal benefits after retirement. The real fund offers benefits that do not deteriorate in real terms, i.e. are protected against inflation. Subsequently we will also consider conditionally indexed funds which offer inflation indexation if the funding ratio is sufficiently high. We restrict our attention to defined benefit pension systems with uniform pricing, i.e. all participants of the pension plan pay the same fraction of the salary. This means the contribution rate is set by the fund uniformly among participants in percentage terms, and each member contributes the same percentage of his/her yearly salary irrespective of the individual characteristics.

The current institutional setting in the Netherlands is such that pension rights are built up for every year worked. For a one year participation in the average-wage[6] pension system, the employee earns the right to receive 1.75% of the current yearly wage after

[6]In case of final wage, there is a phenomenon called back service, which means that additional money has to be contributed if people have steep career patterns. If the average wage scheme is used, it can be shown that back service is no longer important.

retirement[7]. The amount of pension benefit is maximized at 70% of the average wage over the career, and possibly corrected for inflation.

Differences in survival rates, income profiles and in particular deferred time of the annuity imply differences in the value of participation among cohorts. We distinguish the participants along age, gender and educational level. We assume that people can contribute to the fund from the age of 25 till age 64. In addition, we distinguish 5 educational groups[8], such as people only with basic, lower secondary, higher secondary, high education, and education level which reflects the population average.

Survival probabilities for different socioeconomic groups in the Netherlands are not easily available. In Appendix A we explain how estimated survival probabilities per group have been constructed using Dutch data for the population at large and Belgian data on socioeconomic survival probabilities.

There are sizable differences in expected lifetime[9] among cohorts with different educational levels. First, we calculated these differences by assuming that survival probabilities do not improve in the future, and the calculations have been made on the basis of the latest Dutch life table observed in 2003. However, due to improvements in health care or in living standards etc., survival probabilities may change over time. A parsimonious model to capture the dynamics of the survival probabilities is the Lee-Carter (LC hereafter) model as introduced by Lee and Carter (1992). The details on the model and estimation are provided in Appendix B. Therefore, as an alternative to the no improvement in survival assumption, we accommodate the projected improvement in survival probabilities and recalculated the differences in expected lifetime between cohorts. Table 1 shows the educational-, age-, and gender-specific expected lifetime and probability of survival with constant mortality and with mortality improvement.

If survival rates are assumed to be constant, a 25-year-old man with high education

[7]In average wage systems, the participant receives 70% of the average wage after 40 years of participation. This is equivalent with the fact, that one year participation yields the right to receive 1.75% of the current wage after retirement.

[8]Basic education means the primary level education, which is 8 years of school. The second educational group, the lower level of secondary education is defined as the level of education reached after three years of primary education. The third group with the higher level of secondary education consists of people who have 6 years education after primary school. The fourth group has higher or university degree. In addition, we create a group which portrays the average education of the total population of the Netherlands.

[9]Survival data for the Netherlands is downloaded from the Human Mortality Database. University of California, Berkeley (USA), and Max Planck Institute for Demographic Research (Germany). Available at www.mortality.org or www.humanmortality.de (data downloaded on 01.12.2004).

	Constant Mortality						Projected Mortality Improvement					
	men			women			men			women		
Age	Exp. Life	Prob of surviving age 65	Prob of survivin g age 80	Exp. Life	Prob of surviving age 65	Prob of survivin g age 80	Exp. Life	Prob of surviving age 65	Prob of surviving age 80	Exp. Life	Prob of surviving age 65	Prob of surviving age 80
	Average education											
25	76.6	0.857	0.469	81.1	0.903	0.648	78.9	0.880	0.543	85.1	0.930	0.759
35	76.9	0.863	0.472	81.3	0.907	0.650	78.6	0.878	0.530	84.4	0.925	0.738
45	77.4	0.875	0.478	81.7	0.916	0.657	78.6	0.883	0.520	84.0	0.926	0.722
55	78.4	0.907	0.496	82.6	0.942	0.675	79.2	0.909	0.522	84.1	0.945	0.717
64	80.1	0.986	0.539	83.8	0.992	0.711	80.5	0.986	0.553	84.7	0.992	0.733
	High education											
25	78.7	0.888	0.538	82.0	0.912	0.672	80.9	0.906	0.607	86.0	0.937	0.776
35	78.8	0.892	0.540	82.1	0.915	0.674	80.6	0.904	0.594	85.3	0.932	0.757
45	79.2	0.900	0.545	82.5	0.923	0.680	80.4	0.906	0.584	84.8	0.932	0.741
55	80.0	0.925	0.560	83.3	0.947	0.697	80.8	0.927	0.585	84.9	0.950	0.736
64	81.4	0.988	0.598	84.5	0.992	0.731	81.9	0.988	0.612	85.4	0.992	0.751
	Higher Sec. education											
25	76.4	0.854	0.462	81.2	0.904	0.653	78.6	0.877	0.537	85.3	0.931	0.762
35	76.7	0.859	0.465	81.4	0.908	0.656	78.4	0.875	0.523	84.6	0.926	0.743
45	77.2	0.871	0.471	81.8	0.917	0.662	78.4	0.880	0.513	84.2	0.927	0.726
55	78.3	0.904	0.489	82.7	0.943	0.681	79.0	0.907	0.516	84.3	0.946	0.722
64	79.9	0.985	0.533	84.0	0.992	0.716	80.4	0.985	0.547	84.9	0.992	0.738
	Lower Sec.education											
25	75.7	0.842	0.438	80.8	0.900	0.643	77.9	0.867	0.516	84.8	0.927	0.755
35	76.1	0.850	0.442	81.0	0.904	0.646	77.7	0.866	0.502	84.2	0.923	0.735
45	76.7	0.863	0.449	81.5	0.914	0.653	77.8	0.872	0.493	83.8	0.924	0.719
55	77.8	0.899	0.468	82.4	0.942	0.673	78.5	0.902	0.496	83.9	0.945	0.715
64	79.5	0.985	0.512	83.7	0.992	0.709	79.9	0.985	0.527	84.6	0.992	0.731
	Low education											
25	73.9	0.805	0.375	78.8	0.878	0.573	76.1	0.835	0.456	82.9	0.911	0.702
35	74.5	0.816	0.380	79.2	0.884	0.576	76.1	0.835	0.442	82.3	0.907	0.680
45	75.2	0.833	0.388	79.7	0.896	0.584	76.3	0.844	0.433	82.0	0.909	0.660
55	76.5	0.877	0.408	80.8	0.928	0.605	77.2	0.880	0.437	82.2	0.932	0.654
64	78.5	0.981	0.457	82.2	0.990	0.645	78.9	0.981	0.473	83.0	0.990	0.671

Table 1: Educational-specific expected lifetime and probability of survival for selected age groups. The table gives the educational-, age-, and gender-specific expected lifetime and probabilities with the constant mortality (at the level estimated for 2003), and the time-varying future mortality assumptions.

is expected to live 4.8 years longer than a man with basic education, while this difference is 3.2 years for a woman with the age of 25. The differences in life expectations between high and basic educational groups decrease to 2.9 for men and 2.3 for women at the age of 64.

The model with time-varying survival rates predicts further increases of life expectancy.[10] The projected life expectancy at age 25 of a man with average education increases with 2.3 years if the assumption of constant mortality rates is dropped and similar differences hold for other educational levels as well. The corresponding difference for women is 4. The differences decrease to 0.4 years in the case of men, and to 0.9 in the

[10]Note that life expectancy does not increase monotonically with age if the projected mortality improvements are incorporated. This is due to the fact that survival rates are expected to drop considerably which at young ages dominates the effect that people have already reached a certain age.

case of women at the age of 64. Since the methodology we use (see Appendix A on estimating socioeconomic life tables for the Netherlands) to calculate educational-specific survival rates makes sure that the relative differences in gender-specific life expectations between educational cohorts do not change, or at least do not decrease in the future (for details, see Pappas et al., 1993; Preston and Taubman, 1994; Mackenbach et al., 2003), the educational-specific differences in life expectancy with mortality improvement are similar to the results based on the constant mortality assumption.

The present value of a nominal (real) annuity contract depends on mortality rates as well as on the nominal (real) term structure. If tax considerations are ignored, the present value $V^i_{x,t}$ at time t for an individual i with age x of an A dollar nominal annuity as of the age of 65 can be written as (compare e.g. Mitchell et al., 1999):

$$V^i_{x,t} = \sum_{s=65-x}^{\infty} \mathbb{E}_t \left(p^i_{x,t} \right) \frac{A}{\left(1 + R^{(s)}_t\right)^s}, \tag{1}$$

where ${}_sp^i_{x,t}$ is the probability at time t that person i at the age of x is going to live at least for another s years, and $R^{(n)}_t$ is the nominal interest rate at time t for payment in n periods from now. The same expression applies for the value of a real annuity $V^{i,R}_{x,t}$ if the nominal interest rate is replaced by the corresponding real interest rate $R^{R(n)}_t$. In our case A denotes the right to receive 1.75% of the current yearly wage after retirement either in a nominal or real indexed scheme, and $\mathbb{E}_t \left({}_sp^i_{x,t} \right)$ can be calculated by using either the no improvement or improvement assumption in future survival rates. For a detailed derivation of (1), see Appendix B and C, where it is moreover shown that this expression can be extended to the case of conditionally indexed schemes.

The real ($R^{R(n)}_t$) and nominal ($R^{(s)}_t$) term structures used to calculate the present value of the one year participation in a pension fund are presented on Figure 1, which corresponds to a nominal 10-year rate of 4.2% and inflation rate of 1.2% p.a. in January 1, 2004, in the Netherlands.

The money's worth of a fixed (in nominal or real terms) annuity in a collective scheme is presented in Tables 2 and 3. These tables generalize Table 2 in Brown (2002) by providing evidence not only of the annuity value at retirement, but also of the money's worth in the accumulation phase of the life-cycle, thereby adding the age dimension. The figures in the tables show the present value of one year participation for age-, gender-, and educational-specific cohorts.

Table 2 gives the money's worth of participation if survival rates are constant and the term structure of interest rates in Figure 1 is used. A 35-year-old woman with

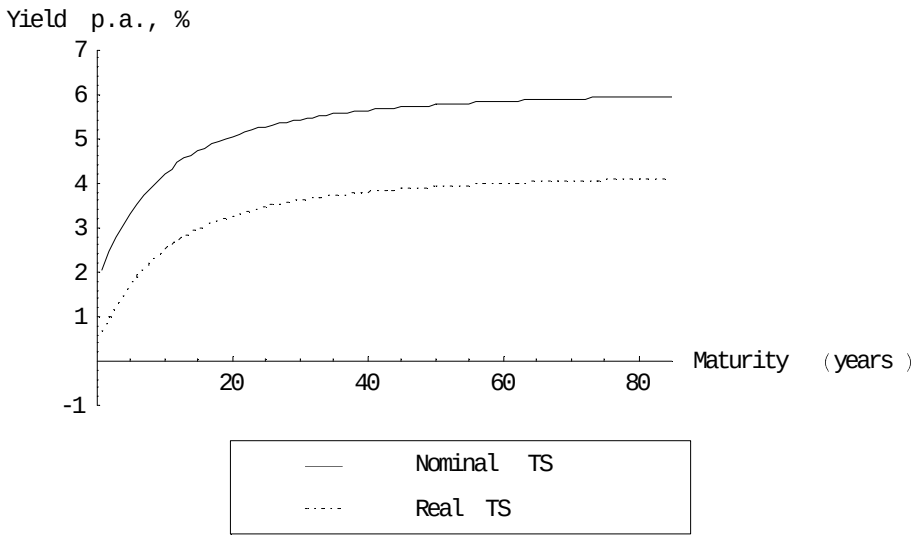

Figure 1: Nominal and real term structures that are used to determine the money's worth of participation in a collective scheme, January 1, 2004. The figure illustrates the nominal and the real term structure of interest rates, which corresponds to a nominal 10-year rate of 4.2% and inflation rate of 1.2% p.a. in January 1, 2004, in the Netherlands.

low education earns 3.21% of the current yearly salary if she participates in a nominal pension scheme, while she earns 6.41% of the current salary in a real pension plan, twice as much as in the nominal case.

The numerical results on the money's worth of one year participation in a nominal fund with different characteristics of people can be summarized as follows. For a given age, the money's worth of a single year participation is decreasing as educational level decreases. For instance, a 25-year-old man earns 1.68% of the annual salary if he is highly educated, and 1.37%, if he attained basic education only. A 25-year-old highly educated man earns 22.6% more pension right than a man with basic education at the same age, which is due to different survival prospects. The difference between the money's worth of participation due to different level of education shrinks to 12.7% for men at the age of 64. The differences in the money's worth measure among the highest and lowest educated groups with the same age are also present for women, however they are somewhat smaller. A 25-year-old woman with the highest education earns 10.5% more than someone with the lowest education, while the difference decreases to 7.2% at

Constant mortality			Low	Lower Sec.	Higher Sec.	High	Average
Nominal Annuity	Men	At 25 yrs	1.37%	1.49%	1.54%	1.68%	1.55%
		At 35 yrs	2.59%	2.81%	2.89%	3.15%	2.91%
		At 45 yrs	4.92%	5.32%	5.45%	5.91%	5.50%
		At 55 yrs	9.52%	10.17%	10.40%	11.18%	10.47%
		At 64 yrs	17.39%	18.23%	18.54%	19.60%	18.65%
	Women	At 25 yrs	1.71%	1.83%	1.85%	1.89%	1.84%
		At 35 yrs	3.21%	3.43%	3.47%	3.54%	3.45%
		At 45 yrs	6.06%	6.46%	6.52%	6.65%	6.49%
		At 55 yrs	11.55%	12.24%	12.35%	12.55%	12.28%
		At 64 yrs	20.22%	21.23%	21.40%	21.68%	21.30%
Real Annuity	Men	At 25 yrs	3.23%	3.54%	3.65%	4.03%	3.69%
		At 35 yrs	5.06%	5.53%	5.69%	6.26%	5.75%
		At 45 yrs	7.98%	8.67%	8.91%	9.75%	8.99%
		At 55 yrs	12.82%	13.77%	14.11%	15.29%	14.23%
		At 64 yrs	19.94%	21.02%	21.44%	22.85%	21.58%
	Women	At 25 yrs	4.11%	4.44%	4.50%	4.61%	4.47%
		At 35 yrs	6.41%	6.90%	6.99%	7.15%	6.94%
		At 45 yrs	10.03%	10.77%	10.90%	11.13%	10.83%
		At 55 yrs	15.86%	16.95%	17.12%	17.45%	17.02%
		At 64 yrs	23.66%	25.04%	25.28%	25.68%	25.14%

Table 2: The present value of participation in collective pension funds with no mortality improvement. The table gives the educational-, age-, and gender-specific money's worth of participation as a percentage of the yearly salary in a nominal and real pension scheme if survival rates are constant over time at the level estimated for 2003.

the age of 64. Due to different gender-specific survival rates a 25-year-old man with low education is expected to earn 1.37% of the yearly salary, a woman with the same age with the same level of education earns 1.71% of the salary, 24.8% more.

Gender- and education-specific survival differences at a given age are reflected in real pension plans as well. The pattern of differences is similar, but the differences in percentage terms are higher for real pension plans, that are analyzed in the lower part of Table 2. This is due the fact, that real interest rates are smaller than the nominal ones, therefore the differences in real pension rights due to different survival characteristics of people with the same age are less affected by the effect of discounting than in the case of nominal pension schemes. The difference between the present value of participation earned by low and highly educated men is 24.8% at the age of 25 and it is 14.6% with the age of 64. For women the corresponding numbers are 12.2% and 8.5%.

Table 2 clearly shows that the age of the participants has a very important role in determining the present value of nominal annuity earned by a one year participation in the fund. In the nominal scheme, the present value of participation for a woman

with low education is 20.22% of her yearly salary at the age of 64, and it is 1.71% of the salary at the age of 25, which is 91.5% less. This is caused by two effects. One is the probability of death, and the other one is the time value of money. Table 1 clearly shows the uncertainty effect in survival. A 25-year-old woman with low education has an 88.7% (see Table 1, column 11, 0.878/0.99) probability to survive till the age of 64. This makes the present value of the annuity decrease by 11.3%. However, the discounting makes the present value of annuity decrease further by another 80.2%. The effect of discounting dominates the differences in the money‘s worth of annuities among groups with different ages. If the term structure shifts downwards, the effect of discounting is obviously less strong. Consequently, the differences in the present value of participation in the real pension plan is smaller (the corresponding number for 91.5% in nominal the plan becomes 82.6% in the real plan) due to the lower real yields.

In Table 2 we assumed that the survival probabilities as observed in 2003 will not improve further. Table 3 presents similar results, but assumes projected mortality improvements as discussed in Appendix B and presented in Table 1. Table 3 shows the value of participation if survival probabilities are time-varying, the nominal 10-year rate is 4.2% and the 1-year inflation rate is 1.2%.

The benefits are obviously higher compared to the case with time-invariant survival probabilities, because the probability of surviving increased for all age groups. Adjustment for mortality improvement has an impact of up to 50-85 bp on the money's worth for some age groups, depending on the scheme. In a nominal pension scheme, a 25-year-old highly educated man earns 19.1% more pension right than a man with basic education at the same age, if he participates in the fund for a year. This difference shrinks to 12.7% for men at the age of 64. A 25-year-old woman with the highest education earns 8.2% more than someone with the lowest education, while the difference decreases to 7.0% at the age of 64. A 25-year-old woman with low education earns 27.6% more than a man of the same age. The money's worth of one year participation is 90.6% lower for a 25-year-old low educated woman, than for a woman with the same characteristics with the age of 64. Similar conclusions hold for real pension schemes.

In pension funds with uniform contribution rates the cost-effective contribution rate is set to cover the market value of the rights assigned to the participants.[11] The deviation between this cost-effective rate for the fund as a whole and the percentage of the wage that reflects the money's worth of the annuity indicates the cost solidarity, that is imposed by the fund; i.e it indicates whether an individual is a net contributor or

[11] This contribution rate varies typically from 12.5% to 17.5% of the yearly salary in the Netherlands.

Mortality Improvement			Low	Lower Sec.	Higher Sec.	High	Average
Nominal Annuity	Men	At 25 yrs	1.52%	1.63%	1.67%	1.81%	1.68%
		At 35 yrs	2.79%	3.00%	3.08%	3.33%	3.10%
		At 45 yrs	5.18%	5.56%	5.69%	6.15%	5.74%
		At 55 yrs	9.80%	10.44%	10.67%	11.44%	10.74%
		At 64 yrs	17.62%	18.46%	18.78%	19.85%	18.88%
	Women	At 25 yrs	1.94%	2.04%	2.06%	2.10%	2.05%
		At 35 yrs	3.55%	3.74%	3.78%	3.84%	3.76%
		At 45 yrs	6.51%	6.87%	6.94%	7.05%	6.90%
		At 55 yrs	12.06%	12.72%	12.83%	13.03%	12.77%
		At 64 yrs	20.69%	21.68%	21.85%	22.13%	21.75%
Real Annuity	Men	At 25 yrs	3.62%	3.91%	4.02%	4.38%	4.05%
		At 35 yrs	5.52%	5.96%	6.12%	6.68%	6.18%
		At 45 yrs	8.46%	9.13%	9.38%	10.21%	9.46%
		At 55 yrs	13.26%	14.21%	14.55%	15.74%	14.67%
		At 64 yrs	20.28%	21.36%	21.79%	23.22%	21.92%
	Women	At 25 yrs	4.77%	5.05%	5.10%	5.20%	5.08%
		At 35 yrs	7.19%	7.64%	7.72%	7.87%	7.68%
		At 45 yrs	10.90%	11.60%	11.72%	11.95%	11.66%
		At 55 yrs	16.71%	17.77%	17.95%	18.27%	17.84%
		At 64 yrs	24.35%	25.72%	25.97%	26.38%	25.82%

Table 3: The present value of participation in collective pension funds, mortality improvement. The table gives the educational-, age-, and gender-specific money's worth of participation as a percentage of the yearly salary in a nominal and real pension scheme if survival rates are allowed to be time-varying.

beneficiary of the scheme in the year under consideration.[12] Our analysis quantifies the solidarities in the typical Dutch pension deal (referred to e.g. by Kune, 2005): from young to old, from men to women, from lower educated to higher educated.

If the costs of switching to another fund are low, and the additional assumptions made in the analysis apply (in particular the assumption that agents unwind the positions imposed by the fund so that only market values are relevant), some groups are better off if they leave the compulsory fund. The uniform premium creates an incentive for young cohorts to avoid the compulsory scheme, e.g. by reducing their labor market participation, by moving abroad or by becoming self-employed. This finding is valid both for nominal and real schemes. Leaving the age effect aside, individuals, whose money's

[12]If the uniform contribution rate is 12.5% in the nominal plan and mortality improvement is taken into account (Table 3) for instance, then a highly educated 25-year-old woman is a net contributor. She pays 12.5% of her salary as a contribution and the present value of her yearly participation is 2.1% of the yearly salary, implying a 10.4% net contribution. However, a woman with the same education level but with the age of 64 is a net beneficiary. She pays 12.5% of the salary and receives 22.13% in exchange for the participation; she benefits 9.63% of her yearly salary. A similar cost-benefit analysis can be carried out based on a given uniform contribution rate for funds in Table 2 or in Table 4.

worth of the imposed annuity is lower than the uniform premium that is charged, have an incentive to buy a tailor-made annuity on the private insurance market if such annuities are offered. If opting out of the compulsory system is feasible at relatively low cost, the sustainability of the compulsory system of course becomes questionable.

4 The money's worth in conditionally indexed collective pension schemes

Pension plans in many countries are typically neither nominal nor real in nature, instead, they are hybrid constructs. Pension funds typically offer inflation indexation of accumulated pension rights if the current state of the fund is good. However the rules of indexation are often not specified explicitly in the contract.

In the Netherlands many pension funds have recently made more or less explicit their indexation promise. If the nominal funding ratio drops below a certain threshold no indexation is granted and the premium is increased. If the nominal funding ratio is above a certain threshold, full indexation is granted and contributions are decreased. In between the thresholds, partial indexation is granted and contributions decrease as the funding ratio increases.

In the sequel we determine the fair value of such a contract, and indicate how it depends on the funding ratio and the asset mix that is selected. We assume that pension rights are fully indexed against inflation if the nominal funding ratio is larger than 1.36, they are partially indexed if the nominal funding ratio is between 1.05 and 1.36, and no indexation occurs below the 1.05 level.

The fund we consider is a large[13] closed-end fund[14] where no premium inflow takes place and no new benefits are built up in the future on the basis of the discontinuity perspective. The liabilities of the fund are used to pay future pension liabilities, and the pension benefits are indexed according to the indexation rules discussed in the previous paragraph. We assume a specific linearly decreasing expected nominal liability stream with duration of 13.4 years, illustrated on Figure 2:

The variable annuities that are offered by participation in a conditionally indexed

[13]Micro-longevity risk, which results from nonsystematic deviations from an individuals expected remaining lifetime is negligible in the case of a large fund.

[14]Alternatively, we could set up a running fund and policy rules are applied for setting the size of the contribution in each year. This would yield a more realistic liability stream but would complicate the analysis.

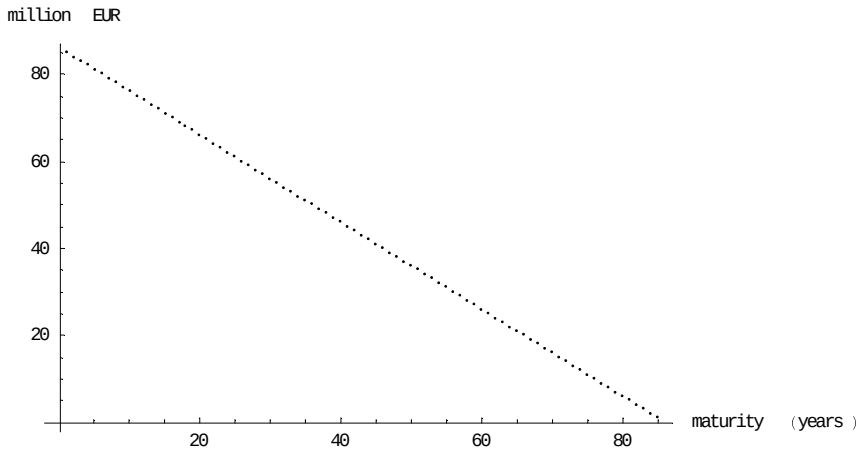

Figure 2: Liability stream of the collective fund. The figure shows the expected liability stream of a large pension fund up to the maturity of 85 years, where no premium inflow takes place and no new benefits are built up in the future on the basis of the discontinuity perspective.

scheme can be priced by using the pricing kernel. The pricing kernel is a particular random variable so that the price of any asset at time t satisfies $P(t) = E(P(t+1)M(t+1)|F(t))$[15]. This implies that the risk premium of any asset is determined by its covariation with the pricing kernel. If markets are complete, i.e. every contingent claim can be replicated by a self-financing portfolio, this pricing kernel is uniquely given. In this case, the price obtained via valuation using the pricing kernel can be shown to be equivalent to the cost required to set up the replicating portfolio. In incomplete markets, however, we cannot identify all risk premia on the basis of the traded assets. In our context, market incompleteness is for instance caused by macro-longevity risk[16], if no annuities are traded, and inflation, if no inflation-linked securities are available. In such situations, one has to make assumptions regarding the pricing kernel specification. We assume that neither longevity risk nor inflation risk is priced.[17] For a detailed

[15]The existence of a pricing kernel is ensured when the financial market is free of arbitrage, which we will assume throughout. For more details on pricing kernels, we refer to Campbell et al. (1997) and Cochrane (2001).

[16]Macro-longevity risk results from the fact that survival probabilities change over time.

[17]There is an empirical evidence that investors demand risk-premium for holding inflation-sensitive assets (see the UK nominal and inflation-linked gilt market for instance in Evans, 1998), however, for simplicity we assume that inflation risk is not priced, because we do not observe inflation-linked

specification of the pricing kernel and a discussion regarding the assumptions made, we refer to Appendix C and to Nijman and Koijen (2006).

		FR=1			FR=1.4			Nominal Plan	Real Plan
		0% Stocks	50% Stocks	100% Stocks	0% Stocks	50% Stocks	100% Stocks		
Constant Mortality									
Men	At 25 yrs	1.74%	1.98%	1.91%	3.28%	2.63%	2.21%	1.55%	3.69%
	At 35 yrs	3.08%	3.50%	3.44%	5.21%	4.41%	3.89%	2.91%	5.75%
	At 45 yrs	5.61%	6.24%	6.23%	8.37%	7.46%	6.85%	5.50%	8.99%
	At 55 yrs	10.53%	11.29%	11.37%	13.66%	12.75%	12.15%	10.47%	14.23%
	At 64 yrs	18.66%	19.26%	19.40%	21.24%	20.57%	20.13%	18.65%	21.58%
Women	At 25 yrs	2.08%	2.36%	2.28%	3.95%	3.16%	2.65%	1.84%	4.47%
	At 35 yrs	3.67%	4.16%	4.09%	6.26%	5.28%	4.64%	3.45%	6.94%
	At 45 yrs	6.65%	7.40%	7.38%	10.02%	8.89%	8.16%	6.49%	10.83%
	At 55 yrs	12.37%	13.30%	13.39%	16.25%	15.11%	14.37%	12.28%	17.02%
	At 64 yrs	21.33%	22.11%	22.27%	24.65%	23.77%	23.20%	21.30%	25.14%
Mortality Improvement									
Men	At 25 yrs	1.90%	2.15%	2.08%	3.61%	2.89%	2.41%	1.68%	4.05%
	At 35 yrs	3.29%	3.72%	3.67%	5.59%	4.74%	4.14%	3.10%	6.18%
	At 45 yrs	5.86%	6.52%	6.52%	8.78%	7.82%	7.16%	5.74%	9.46%
	At 55 yrs	10.80%	11.58%	11.67%	14.04%	13.10%	12.47%	10.74%	14.67%
	At 64 yrs	18.90%	19.51%	19.67%	21.55%	20.86%	20.41%	18.88%	21.92%
Women	At 25 yrs	2.35%	2.65%	2.56%	4.51%	3.57%	2.98%	2.05%	5.08%
	At 35 yrs	4.03%	4.56%	4.48%	6.94%	5.81%	5.09%	3.76%	7.68%
	At 45 yrs	7.10%	7.90%	7.89%	10.79%	9.54%	8.71%	6.90%	11.66%
	At 55 yrs	12.88%	13.86%	13.96%	17.03%	15.79%	14.97%	12.77%	17.84%
	At 64 yrs	21.80%	22.61%	22.78%	25.30%	24.36%	23.74%	21.75%	25.82%

Table 4: The value of conditionally indexed rights. The table gives the value of the age- and gender-specific conditionally indexed rights as a percentage of the yearly salary for an annuity population with an average education. We consider a fund with a starting nominal funding ratio of 1, and alternatively, we consider another fund with identical characteristics, except, that the starting nominal funding ratio of the latter is 1.4. We allow for alternative investment strategies. The assets are invested into 10-year nominal bonds and stocks. The results of the conditionally indexed schemes are compared to the purely nominal and real plans. In the upper part of the table survival probabilities are constant over time at the level estimated for 2003, while in the lower part survival probabilities are time-varying.

Table 4 shows the value of the conditionally indexed rights for constant mortality at the level estimated for 2003, and for time-varying mortality rates. The value of the conditionally indexed pension rights are bounded by the value of a nominal and a real plan. If the current funding ratio is low, the likelihood of indexation is low and as a consequence, the conditionally indexed scheme resembles a nominal scheme. If the

instruments in the Netherlands.

current funding ratio is high, the pension scheme is highly comparable to a real pension scheme.

Note that the asset mix influences the money's worth of participation in a conditionally indexed scheme. Since the nominal rights are guaranteed, the money's worth of participation in a scheme with low funding ratio increases if more risk is taken. If the fraction of assets invested into stocks reaches a certain threshold, the money's worth starts decreasing for some age groups. Likewise, a large fraction invested in stocks reduces the value of participation if the funding ratio is high. The reason of the decline in the money's worth if the risk increases is the fact that the sponsor gets the up-side above the full indexation.

For the constant mortality case, if the initial nominal funding ratio is 1, a 64-year-old woman earns 21.33% of the salary if 100% of the assets are invested into 10-year nominal bonds, while the present value of participation increases to 22.11% with a 50% 10-year nominal bonds and 50% stock asset mix, and it further increases to 22.27% with an asset mix of 100% stocks. If the initial funding ratio is 1.4, the increase in the risky assets in the asset mix yields a lower value for the value of participation, because the probability of ending up in the bad state is higher with more volatile stock investments. The pattern is similar if mortality rates are time-varying, and the present value of participation is higher due to the improvement in expected lifetime.

The main conclusion is that all members, regardless of the age of the participants, have an incentive[18] to opt out of a fund with a low funding ratio to a fund with a higher funding ratio. The value of participation is lower and the cost of participation is the same or higher than in a fund with a higher funding ratio. In the case of time-varying mortality, the value of participation is 2.15% of the yearly salary for a 25-year-old man if he participates in a fund with the initial funding ratio of 1 and the asset mix of 50%-50%. This person has an incentive to change pension fund, because the identical fund with a higher funding ratio is more appealing. The value of the one year participation increases to 2.89% of the yearly salary. A 64-year-old man with earns 19.51% of the yearly salary in a fund with funding ratio of 1 and 50% of stocks. However, if the funding ratio increases to 1.4 and all the other characteristics of the fund remain the same, then this person earns 20.86%.

Besides the indexation quality, the other motives generated by the gender-, education- and age-specific survival differentials still play an important role for opting out of the

[18]Transfer value of pension rights is often calculated with actuarial valuation, which might distort the incentives induced by the indexation quality and the asset mix.

collective pension plan, either individually or collectively. However, we do not want to replicate those arguments again in this section.

5 Conclusions

The money's worth of participation in a collective fund is different among age groups and socioeconomic groups. Young participants have an incentive to opt out if uniform pricing over age groups is applied. Similar but much smaller differences occur between socioeconomic groups and male/female participants. Generally, the money's worth of participation for lower educated is lower than for higher educated cohorts, and the value of participation for men is less than for women. Young, lower educated and males have the incentive to leave the collective fund, and to switch to another job, sector, where the characteristics of the participants of the new fund are closer to the characteristics of the people with the incentive to opt out. Alternatively, they can reduce their labor supply or try to obtain access to insurance products that are priced on the basis of their individual characteristics.

Indexation quality is another factor that affects money's worth. The money's worth of participation in a fund with a low funding ratio is less, therefore participants of a fund with a low funding ratio might have an incentive to switch to a fund with a higher funding ratio. It should be noted that switching from fund to another may have implications for the rights built up. If so, then the incentives provided should be balanced with the switching costs. We find that especially young individuals might have the incentive to opt out, and this is exactly the group for which the rights built up are the lowest.

If people start opting out on large scale, a pension scheme is not sustainable. Consequently, the arguments underlying uniform pricing should be carefully reconsidered.

Note that a transition from uniform pricing in collective pension schemes to pricing on the basis of market value conditional on age (and possibly also on other individual characteristics) generates a substantial transitional problem, not unlike that of a transition from pay-as-you-go (PAYG) to funded systems. Young generations that have received the implicit promise that their money's worth would be more than their contribution during the last part of their working life, will have to be compensated if uniform pricing over age groups would be abolished.

In this paper we made a number of strong assumptions. Only old-age pension has been considered and part of the money's worth differentials that have been identified can be compensated by the partner pension arrangements that are usually also included

in actual pension products. Moreover we made the strong assumption that only the market value of what is received is relevant because agents can and will unwind all product features that are imposed by the pension fund. Subsequent analysis will have to consider the question to what extent these assumptions dominate the analysis.

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A Socioeconomic life tables

This section addresses the methodology that has been used to construct the educational-specific cohort life tables. The main complication is that these data are not publicly available for the Netherlands. Deboosere and Gadeyne (2002) calculated educational-specific cohort life tables for Belgium for the period of 1991-1996. We use their results to estimate survival probabilities per socioeconomic group for the Netherlands. Deboosere and Gadeyne (2002) distinguish 4 educational levels, namely low education (L), lower secondary (SL), higher secondary (SH), and high (H).

Following Brown et al. (2002), we construct educational-specific cohort life tables, assuming that relative discrepancies in mortality rates between different socioeconomic groups are constant over time. Although it is hard to verify this assumption, Pappas et al. (1993), Preston and Taubman (1994) and Mackenbach et al. (2003) document that differences in mortality rates between socioeconomic groups are not shrinking in the late 20th century, instead, there may have been widening. If the latter is true, then we underestimate the differences between educational groups. Therefore, our assumption, if not satisfied, seems to result in conservative estimates of the differences between different educational groups. Secondly, we assume that the differences between socioeconomic groups in Belgium provide a reasonable representation of the Dutch population.

First of all, we construct the relative discrepancies from the average mortality rates for all socioeconomic groups:

$$\Delta_x^i = \frac{\hat{q}_{x,t^*}^i}{\hat{q}_{x,t^*}}, \tag{2}$$

on the basis of the Belgian data, where $\hat{q}_{x,t^*}^i$ indicates the 1-year mortality rate at time t^* for a person of age x that is within socioeconomic group i, $i = L, SL, SH, H$. $\hat{q}_{x,t^*}$ is the weighted average of all Belgian mortality rates at time t^*, where the weighting

occurs via the number of people present in the socioeconomic group in the Netherlands:

$$\hat{q}_{x,t^*} = \frac{\sum_i N_{i,t^*} \hat{q}^i_{x,t^*}}{\sum_i N_{i,t^*}}, \tag{3}$$

where N_{i,t^*} indicates the number of people with age x present in socio-economic group i in the Netherlands[19]. Secondly, we apply these ratios to the Dutch population in order to calculate educational-specific mortality rates at all future points in time

$$q^i_{x,t} = q_{x,t} \Delta^i_x, \tag{4}$$

where $q^i_{x,t}$ is the probability that an individual with age x at time t who is within socioeconomic group i, dies in the next year. The weighting in (2) is important since the composition of the Belgian population may be different than the Dutch population.

Finally, we fit cubic polynomials to smooth the ratios for different ages, and we use the smoothed ratios in order to calculate educational-specific cohort life tables for the Netherlands.

B Modeling survival probabilities

A crucial element in the determination of the money's worth of participation in a collective pension scheme is $\mathbb{E}_t \left(I_{(\tau<S)}(x,\tau) \right)$, see (30). In this section we discuss a convenient way to model mortality rates, as has been introduced by Lee and Carter (1992).

Let L denote the realization of the uncertain life table. By using the law of iterated expectations we can rewrite $E_t \left(I_{(\tau<S)}(\tau) \right)$ as follows:

$$\mathbb{E}_t \left(I_{(\tau<S)}(x,\tau) \right) = \mathbb{E}_t \left(\mathbb{E}_t \left(I_{(\tau<S)}(x,\tau) \mid L \right) \right) = \mathbb{E}_t \left({}_\tau p_{x,t} \right), \tag{5}$$

where ${}_\tau p_{t,x}$ is the probability that an individual with age x at time t is going to survive at least till year τ.

In the subsequent model we assume that the probability distribution of survival is uncertain. However, instead of modeling of ${}_\tau p_{x,t}$ directly, we model the time series of the log of the force of mortality $\mu_{x,t}$ [20] to calculate $\mathbb{E}_t \left(I_{(\tau<S)}(x,\tau) \right)$. In the sequel, we

[19] The educational distribution of the Dutch active population for year 2002 used in the calculations were downloaded from CBS Netherlands (http://statline.cbs.nl)

[20] The force of mortality, at time t, of an individual with age x is defined as: $\mu_{x,t} = \frac{f_t(x)}{1-F_t(x)}$, where f_t (F_t) denotes the pdf (cdf) at time t of the lifetime of a newly born. For the estimation and more details, see e.g. Gerber (1997).

will assume that for any integer age x, and any time t, it holds that:

$$\mu_{x+\tau,t} = \mu_{x,t}, \qquad \text{for all } \tau \in [0,1). \tag{6}$$

Then, one can verify that

$$_{\tau}p_{x,t} = \exp\left(-\sum_{i=t}^{t+\tau} \mu_{x+i,i}\right). \tag{7}$$

An important property of a model is to allow for a trend in mortality rates as has been observed historically due to improvements in medical care. A parsimonious model to capture the dynamics of the mortality rates is the Lee and Carter (LC hereafter) model. We assume that the log force of mortality is affine in a latent factor u_t, which captures the trend in mortality rates. Formally,

$$\ln \mu_{x,t} = a_x + b_x u_t + \varepsilon_{x,t}, \tag{8}$$

where the coefficients, a_x and b_x are age-dependent. An additional error term, $\varepsilon_{x,t}$, which is time- and age-specific, is added to capture particular age-specific influences that are not properly accounted for by the general trend. If $D_{x,t}$ denotes the number of death at time t in a cohort aged x, and $E_{x,t}$ is the number of person years, the so-called exposure, then the force of mortality can be approximated as $\mu_{x,t} \approx \frac{D_{x,t}}{E_{x,t}}$[21].

The estimation procedure of the LC model has been done in several steps. First of all, a singular value decomposition is used to retrieve an estimate of the series of the latent factor, $\hat{u}_t$. Subsequently, OLS regression are run to estimate the age-specific α_x and β_x, resulting in $\hat{\alpha}_x$ and $\hat{\beta}_x$. Once this procedure is applied, observed death numbers are generally not exactly equal to the model-based death numbers. Therefore, a correction step is made, the estimate for the latent factor at a certain point in time is adjusted so that observed number of death at time t equals the one implied by the model, i.e. $\tilde{u}_t$ solves

$$\sum_x D_{x,t} = \sum_x E_{x,t} \exp(\hat{\alpha}_x + \hat{\beta}_x \tilde{u}_t), \tag{9}$$

Finally, the Box-Jenkins method has been used to identify the dynamics of the latent factor u_t. The resulting specification for the latent factor driving the trend in mortality rates is

$$u_{t+1} = \mu + u_t + \eta_{t+1}, \tag{10}$$

[21]For more details on estimating the force of mortality by the exposure and the death number, see Gerber (1997).

with $\eta_{t+1} \overset{i.i.d.}{\sim} D(0, \sigma_\eta^2)$, i.e. the latent factor follows a random walk with drift.

In order to calculate the $\mathbb{E}_t \left(I_{(\tau<S)}(x,\tau) \right)$, we need to determine $\mu_{x,t+s}$, $s > 0$ in (7). We find

$$\begin{aligned} \ln \mu_{x,t+s} &= a_x + b_x u_{t+s} + \varepsilon_{x,t+s} && (11) \\ &= \ln \mu_{x,t} + b_x \left(u_{t+s} - u_t \right) + \varepsilon_{x,s} - \varepsilon_{x,t} && (12) \\ &= \ln \mu_{x,t} + b_x \left(\sum_{i=1}^{s} \eta_{t+i} + s\mu \right) + \left(\varepsilon_{x,s} - \varepsilon_{x,t} \right). && (13) \end{aligned}$$

Lee and Carter (1992) calculate the different sources of uncertainty in the age-specific log mortality rate, and find that the disturbance term of the latent process dominates the error of the overall forecasted mortality rates. Lee and Carter (1992) report that in long-term forecasts about 95% of the variance is generated by innovations of the latent variable process. Therefore, we abstract from all other uncertainties. Then we find

$$\mu_{x,t+s} = \mu_{x,t} \exp \left(b_x \left(\sum_{i=1}^{s} \eta_{t+i} + s\mu \right) \right), \tag{14}$$

and we calculate $\mathbb{E}_t \left({}_\tau p_{x,t} \right)$ with simulation by using (7).

We use 100 yearly observations of number of death and exposure in the Netherlands, from 1904 till 2003[22]. We calculated force of mortality rates from the data provided by The Human Mortality Database (available at www.mortality.org or www.humanmortality.de and the data was downloaded on 01.12.2004) for 61 age groups: 25, 26, ...84, 85+. The last category denoted by 85+ refers to the average mortality rate of people with the age of 85 or older. The reason that we did not create age-groups over the age of 85 is the following. The number of people exposed to risk is relatively low in age-groups above 85 (e.g. 85-89, or 90-94 etc.) in the early 20th century. In order to get the time-series of mortality rates of elderly people calculated from sufficiently large number of observations we merged all the age-groups above year 85.

The parameter estimates of the model, which are used to forecast survival rates are as follows. As far as the latent process is concerned in (10), the drift term is -0.128 and

[22]The trends in the age-specific log mortality rates with the random walk with drift specification depend on the first and the last observations (see for instance Girosi and King, 2005), therefore the forecasted log mortality rates are going to be sensitive to the sample period applied in the estimations. The CBS Netherlands was established in 1899, and it became the main institution responsible for collecting statistical (including population) data. Since then, data collection on births and deaths became better organized and standardized, and the data is less susceptible to measurement problems. Consequently, the sensitivity of the estimation results to the sample period suggests to use the data starting at the beginning of the 20th century.

the standard deviation of the disturbance term is 0.5 for women. The corresponding estimates for men are $-.095$ and 0.57 respectively.

The estimates of the age-specific sensitivity coefficients (i.e. the sensitivity of the log death rates to the change in the latent process) for women and men are illustrated on the following figures:

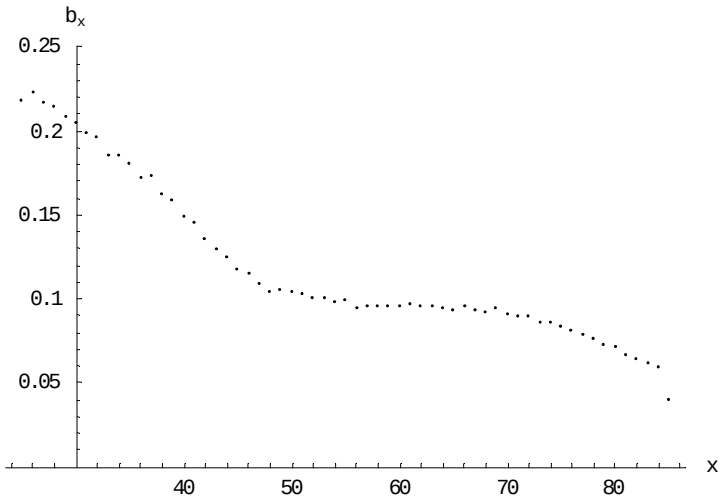

Figure 3: The age-specific sensitivity coefficients b_x for women. They represent the sensitivity of the age-specific log death rates to the change in the latent process.

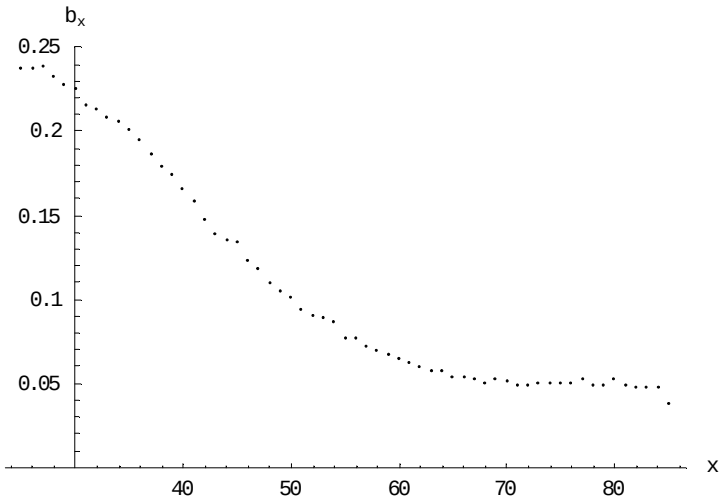

Figure 4: The age-specific sensitivity coefficients b_x for men. They represent the sensitivity of the age-specific log death rates to the change in the latent process.

C Money's worth of participation in collective pension schemes

In this appendix we outline the valuation approach that has used throughout this paper to determine the money's worth of a single year participation. The institutional setting that has been adopted is a so-called average wage system. This implies that the average wage during the individual's working life serves as the metric to determine the pension benefit. Moreover, since inflation potentially erodes the investments of participants, pension funds often aim to provide some form of inflation indexation, which is accounted for. We assume that participation in the pension scheme implies compulsory annuitization of the accumulated benefits at the retirement date.

Formally, suppose that the individual participates l years in the collective pension scheme and receives at the beginning of each year the nominal wages $w_t, .., w_{t+l-1}$. Suppose, the individual retires at time T, with $T > t + l - 1$. At that point in time, the accumulated retirement benefit is converted into an annuity. The pension benefits are paid at the beginning of each year starting in year T, and the annual payment equals a fraction α of the average wage if one has participated for 40 years and scaled proportionally otherwise, accounted for the indexation policy of the pension fund. In the numerical applications, α has been set to 70%. Then the total payoff of l years participation equals

$$\sum_{\tau=T}^{\infty} I_{(\tau<S)}(x,\tau)\left(\alpha\frac{l}{40}\right)\left(\frac{1}{l}\sum_{i=0}^{l-1} w_{t+i}\mathcal{I}(t+i,\tau)\right), \tag{15}$$

where S is the year in which an individual dies, and is therefore a stopping time. The indicator function $I_{(\tau<S)}(x,\tau)$ in (15) equals 1 if the individual with age x survives year τ, and it is zero otherwise. The second term in (15), $\left(\alpha\frac{l}{40}\right)$, accounts for the number of years that an individual has participated. If $l = 40$, the individual receives a fraction α of the (possibly indexed) average wage. The last term in (15) determines the average wage, accounting for the indexation granted by the pension scheme. $\mathcal{I}(t+i,\tau)$ denotes the inflation indexation provided from time $t+i$ to time τ. If we denote the (commodity) price level at time t by Π_t, then $\mathcal{I}(t,\tau)$ can be defined as

$$\mathcal{I}(t,\tau) = \prod_{s=t+1}^{\tau}\left(1 + h\left(FR_s\right)\frac{\Pi_s}{\Pi_t}\right), \tag{16}$$

where $h\left(FR_s\right)$ represents the indexation policy of the pension fund, depending on the funding ratio at time s[23]. Examples are $h\left(FR\right) = 0$ for a nominal pension scheme and

[23]The funding ratio has been defined in this paper as the ratio of assets to the nominal value of the

$h(FR) = 1$ for the real counterpart. On the other hand, in this paper we consider the conditionally indexed scheme as well. This implies that full indexation is given if the funding ratio is sufficiently high (i.e. $FR > U$), but no indexation is granted if the funding ratio is too low (i.e. $FR < L$). In between the fraction of inflation indexation is determined proportionally. Formally,

$$h(FR) = \begin{cases} 0 & , \ FR < L \\ \frac{FR-L}{U-L} & , \ FR \in [L, U] \\ 1 & , \ FR > L. \end{cases} \tag{17}$$

Before determining the value of the payoff, it is important to realize that the total payoff is additive in the payoffs of a single year participation in year $t+i$, i.e.

$$\frac{\alpha}{40} \sum_{\tau=T}^{\infty} I_{(\tau<S)}(x,\tau) \left[w_{t+i} \mathcal{I}(t+i,\tau) \right]. \tag{18}$$

This property is natural within the average wage system, but is no longer valid in a final wage system, in which back service issues come into play. i.e. the decision to participate an additional year is dependent on the previous wages earned. In the average wage system, these considerations are irrelevant and therefore, we can focus in this paper on a single year participation of an individual within a collective pension scheme.

In order to value the payoff in (18), we specify a pricing kernel that is consistent with a simple financial market[24]. The relevant economic factors are assumed to the real interest rate ($R_t^{R(1)}$), inflation (π_t), and stock returns in excess of the nominal short rate (r_t). The dynamics are captured by a VAR(1) - model, in which we assume that the process for inflation and the real interest rate move independently. Finally, we assume that inflation and the real interest rate are independent from excess stock returns. Formally,

$$R_{t+1}^{R(1)} = \mu_R + \phi_R \left(R_t^{R(1)} - \mu_R \right) + \varepsilon_{t+1}^R \tag{19}$$

$$\pi_{t+1} = \mu_\pi + \phi_\pi (\pi_t - \mu_\pi) + \varepsilon_{t+1}^\pi \tag{20}$$

$$r_{t+1} = \mu_r + \varepsilon_{t+1}^r, \tag{21}$$

liabilities.

[24]The model is similar, but not identical to the market specified in Campbell and Viceira (2001) or in Brennan and Xia (2002). They model the realized inflation as the sum of an expected and an unexpected inflation component, where the expected inflation is characterized by an AR(1) process. Instead, we model the realized inflation without decomposition, similar to Ang and Bekaert (2005). Ang and Bekaert (2005) modeled the inflation process as an ARMA(1,1), however, we did not find evidence for that. The yearly realized inflation was found to be best characterized by an AR(1) process based on Dutch yearly inflation data, provided by the CBS Netherlands.

with

$$\varepsilon_{t+1} \equiv \left(\varepsilon_{t+1}^{R}, \varepsilon_{t+1}^{\pi}, \varepsilon_{t+1}^{r}\right) \overset{i.i.d.}{\sim} N\left(0_{3\times 1}, diag\left(\sigma_R^2, \sigma_\pi^2, \sigma_r^2\right)\right), \tag{22}$$

and inflation is defined as $\pi_{t+1} \equiv \log \Pi_{t+1} - \log \Pi_t$. For the specification of the real pricing kernel (M_{t+1}), we postulate

$$-\log M_{t+1} = \alpha + \delta R_t^{R(1)} + \beta_R \varepsilon_{t+1}^{R} + \beta_r \varepsilon_{t+1}^{r} + \eta_{t+1}, \tag{23}$$

where $\eta_{t+1} \overset{i.i.d.}{\sim} N\left(0, \sigma_\eta^2\right)$ and η_s and ε_t mutually independent for all t and s. We refer to Campbell et al. (1997) for a motivation of such a pricing kernel. Note that an assumption underlying the kernel specification in (23) is that the 1-period inflation risk is not priced in real terms, in line with Ang and Bekaert (2005) and Campbell and Viceira (2001).[25] Stated differently, ε_{t+1}^{π} does not appear in the pricing kernel. In order to value nominal liabilities, we make use of the link between the nominal and the real pricing kernel

$$m_{t+1}^{\$} = m_{t+1} - \pi_{t+1}. \tag{24}$$

It is well-known that the affine nature of these models translates in affine nominal and real yields at all maturities and therefore, the corresponding bond prices are exponentially affine in the real rate and inflation, see for instance Campbell and Viceira (2001). Formally, we obtain

$$P_t^{(n)} = \exp\left(-A_n - B_{n,1} R_t^{R(1)} - B_{n,2}\pi_t\right), \tag{25}$$

for the price of a nominal bond at time t with time to maturity n and

$$P_t^{R(n)} = \exp\left(-A_n^R - B_{n,1}^R R_t^{R(1)}\right), \tag{26}$$

for the price of a real bond at time t with time to maturity n.

Using the nominal pricing kernel, the price of any nominal payoff X_{t+1} at time t can be obtained via

$$P_t = \mathbb{E}_t\left(M_{t+1} X_{t+1}\right). \tag{27}$$

In the same spirit, the value of a single year participation within a collective pension scheme can be determined as

$$\frac{\alpha}{40}\mathbb{E}_t\left(\sum_{\tau=T}^{\infty} I_{(\tau<S)}(x,\tau) M_{t,\tau}^{\$}\left[w_t \mathcal{I}(t,\tau)\right]\right), \tag{28}$$

[25] There is an empirical evidence from the UK indexed guilt market that investors demand risk-premium for holding inflation-linked assets (see, for instance Evans, 1998), however, for simplicity we assume that inflation risk is not priced, because we do not observe inflation-linked instruments in the Netherlands.

with

$$M^{\$}_{t,T} = \prod_{s=t+1}^{T} M^{\$}_{s}. \tag{29}$$

Important elements in calculating this expectation are the dependencies between S, the time at which the individual dies, the financial markets, and the pricing kernel. In doing this, we make the common assumption that S and the financial market are independent. Secondly, we assume that the time of death, S, is independent of the nominal pricing kernel. This assumption is somewhat more subtle. If we consider a large collective pension scheme, then idiosyncratic risks in the individual life times will be negligible as an application of the law of large numbers. However, when survival probabilities of the participants as a whole increase due to improvements in medical care, this does constitute an important risk factor for the pension fund. Since we are not able to identify the 'price of mortality risk', we assume throughout that mortality risk is not priced, implying that the conditional expectation for in (28) factorizes into

$$\frac{\alpha}{40} \sum_{\tau=T}^{\infty} \mathbb{E}_t \left(I_{(\tau<S)}(x,\tau) \right) \cdot \mathbb{E}_t \left(M^{\$}_{t,\tau} \left[w_{t+i} \mathcal{I}(t+i,\tau) \right] \right). \tag{30}$$

Appendix B discusses in detail how we model $\mathbb{E}_t \left(I_{(\tau<S)}(x,\tau) \right)$. The second part can be valued using the specification of the financial market as presented before, in conjunction with the pricing kernel. When the pension schemes are straight nominal or real, the second conditional expectation can be determined easily using the closed-form solutions that result from the affine term structure model. In case of conditionally indexed pension schemes, we use Monte Carlo techniques to determine this value. In this simulation procedure, standard variance reduction methods, like control variate and antithetic variables, turn out to be useful to reduce the Monte Carlo error.

We calibrated the parameters of the financial market in such a way that reflects the main stylized facts in the observed nominal yield, inflation and stock market return data in the Netherlands. The 1-year nominal yield between 1975 and 2004 is proxied by the the Dutch 1-year euro (previously guilder) interest rate swap middle rate[26] downloaded from Datasteam, and the yearly inflation rate between 1975 and 2004 was supplied by the CBS

[26]The zero-coupon yield data are available for the period starting only from year 1997, which is very short to estimate its time-series properties. The euro/guilder interest rate swap market might contain some counterparty risk, however, the depth and the quality of the market in London is likely to make the counterparty risk limited. The comparison of the zero-yield with the swap rate in the period between 1997 and 2004 yielded a deviation of at most 0.1% point, also suggesting, that the swap rate is likely to be a good proxy.

Netherlands. The observed nominal term structure has the following characteristics. The first order autocorrelation of the nominal 1-year rate is about 0.8 if the yearly inflation is also included in the regression as an explanatory variable, which has a coefficient of -0.08. The autocorrelation coefficient of the yearly inflation was estimated to be in the magnitude of 0.75. The standard deviation of the nominal annual 1-year interest rates and inflation rates are 1.8% and 1.1% respectively, with a correlation coefficient of about 0.6. The mean of the 1-year nominal yield is in the order of magnitude of 5.6%, and for the yearly inflation the corresponding value is about 2.3%. Moreover, we estimated a 1.2% term premium on a nominal bond with a maturity of 50 years by using a single factor affine Gaussian term structure model driven by the 1-period nominal rate[27].

In order to match the above mentioned characteristics of the observed nominal term structure to a large extent, the autocorrelation coefficient of the real 1-year yield was chosen to be 0.85 with a mean of 3.3% and a standard deviation of 1.6%.

The excess stock return is about 6%[28] with a standard deviation of 24% p.a., based on the total return index of the Dutch market downloaded from Datastream for the period between 1983 and 2004.

Because of the long-term nature of the pension claims, the correct representation of the long end of the term structure is far more important than that of the short end. The model-implied long rates at the beginning of 2004 were below the observed long rates at that time. To fix this problem the factors have been rotated where the nominal 10-year rate takes the role of the 1-year rate and the observed 10-year rate is taken as input for the analysis.

[27]The market price of risk parameter was calibrated by observing the 10-year nominal yield with error, and the 10-year rate between 1975 and 2004 was proxied by the 10-year benchmark yield provided by Datastream. For more methodological details, see Ang and Piazzesi (2003) for instance.

[28]Fama and French (2002) suggest that the equity premium estimated from fundamentals (for instance, the dividend or earnings growth rates) can be much lower than the equity premium produced by the average stock return. For simplicity, to calculate the excess return we used the average stock return in the sample from 1983 and 2004 and no fundamentals.