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ABSTRACT

In Sweden, employers pay non-wage costs for their workforce in the form of legislated employment tax and collective fees. For parts of the workforce, the collective fees are progressive with respect to the employee's age and wage. The objective of this paper is to examine how non-wage costs affect voluntary early retirement. To this end we use a large longitudinal employer–employee matched data set with administrative records of the private sector in Sweden. We exploit the variation in collective fee costs across companies to identify employer incentives to encourage early retirement. The results from the instrumental variable estimator suggest that a 1 percentage point increase in non-wage costs in relation to wage costs increases retirement by 6 percent. Further, given the wage sum and workforce structure, large firms spend more on non-wage compensation than small firms. The share of non-wage costs in relation to the wage sum is also positively linked to net employment growth.

Keywords: Early retirement, non-wage labor costs, pensions, labor demand, collective fees.

JEL codes: J26, J23, J21, J32.

1 Introduction

In the economic literature, retirement is usually seen as a choice of the individual, with a focus on the individual's economic incentives, health status and earnings (see e.g., Gruber & Wise, 2004; Hurd, 1990, and the references therein). However, it is also highly likely that the performance of firms and firms' economic incentives to hire or give older employees early retirement offers are important for the individual decision. Withdrawal from the labor market may in fact be a joint decision on the part of the employee and the employer (the relevant literature is presented below). For the demand of older workers the labor costs are often seen as the main disadvantage relative to that of younger workers (see Munnell et al, 2006; and references cited in Skirbekk, 2008). In particular, older workers have higher wages on average but are also linked to substantial non-wage costs such as pension premiums and health insurance expenses that on average are higher than those for younger workers.

In Sweden the non-wage costs, or social costs, consist of legislated employment tax and collective fees, the latter being stipulated by collective agreements between unions and employers' federations. Almost all employees are covered by an agreement but the agreements vary across different parts of the labor market. The collective fees consist mainly of contributions to the second-pillar pensions (the occupational pensions) but also – to a lesser extent – other collectively agreed insurances. While the employment tax is proportional to the wage sum, the collective fees are, interestingly, for some groups of workers, progressive in relation to the employee's age and wage. This causes an age-related variation in workforce non-wage costs, which implies that the wedge between the wage costs and the gross wage becomes much broader for the older workforce.

Given the higher costs, employers may have financial incentives to offer early retirement to their older workers and substitute them with younger personnel. Furthermore, a solution involving an early retirement pension for an older worker is in many cases financially favorable for the employer compared with continued employment (described further below). Thus, although older employees in Sweden are protected via seniority rules ('last-in first-out'), employers have in the case of redundancy great financial possibilities to persuade an older worker to retire 'voluntarily' by the use of early retirement packages.[1] However, younger workers may be imperfect substitutes for older workers. There may also be positive educational and experience spillover effects of older workers. The composition with respect to age and education may be a result of an intricate dynamic pattern that differs with respect to the specific surroundings and context in which a firm operates.

In this paper, we study the impact of increased labor costs on early retirement. In particular, we examine how the *tax rate* – defined shortly – affects voluntary early retirement. By 'voluntary early retirement' we mean

[1] Naturally, independently of the (short-run) economic gain of granting early retirement, the employer might see other, long-run benefits in restructuring the workforce. Older workers may hinder a necessary renewal of the workforce, and early retirement may permit the continued employment of younger staff members. Investments in older staff might be less attractive as they have fewer years until expected retirement. As for the individual, she might feel forced by social norms to make room for younger workers. In addition, a settlement that involves 'voluntary' early retirement may be more socially acceptable than downright dismissal.

the transition from work to old age pension.[2] By the tax rate, which is a firm-level variable, we mean the firm's non-wage costs (compulsory social security contributions, payroll taxes and other charges, and pension costs, including pension provisions and premiums to the occupational pensions) divided by the wage sum (salaries and other forms of compensation).[3, 4] The tax rate thus measures the added-on costs of the workforce and varies with the characteristics of the workforce. As was explained above, the tax rate is expected to rise considerably with an older and a well-paid workforce.

Our primary interest in this study lies in whether a worker is more likely to retire if he or she is employed by a company with a high tax rate compared with a worker employed by a company with a low tax rate, all else being equal. We argue that the policy implication of this effect largely concerns the incentives of the employer to push for early retirement and less so the employee's own pull for early retirement. For instance, it is reasonable that the financing design of occupational pension plans, with employer-paid pension premiums allocated disproportionally late in the career, is likely to be of significant importance to the employer's incentives to offer early retirement but irrelevant to the work and retirement incentives of the employee herself.

To study these issues we make use of an extremely detailed longitudinal employer–employee matched data set with administrative records, covering the private sector from 1997 to 2005, in which we observe a wide range of characteristics of both the employee and the employer. For the individual we observe earnings, education, sector of employment, industry, etc. The business statistics for companies give us reliable measures of labor costs (wage costs and non-wage costs), value-added, capital investments, etc., as well as the size and the age, gender and educational composition of the workforce in the firm.

There is a methodological problem since the unobserved retirement propensity of the individual ('the distaste for work') may well be negatively correlated with the earnings levels and, thus, with the social costs (Coile & Gruber, 2000; Krueger & Pischke, 1992, among others). This will on average spill over to the social costs at the firm level, leading to a downward bias in the ordinary least square (OLS) estimate of the tax rate effect. To make an attempt to identify the causal implication of the tax rate on early retirement behavior we will employ an instrument variable estimator and assume that the workforce age and education composition at the firm level is exogenous to the unobserved retirement propensity at the individual level (holding other variables in the model fixed, including the worker's age and wage, industry, year and the key characteristics of the employer).

[2] We do not study 'involuntary' transitions out of the labor market that are health-related (i.e., transitions via sickness insurance or disability pension) or job losses because of unemployment.
[3] Henceforth non-wage costs and social costs are used exchangeably.
[4] To our knowledge empirical data on social costs only exist at the firm level. Hence, we are unfortunately not equipped with a measure of the social costs at the individual level.

Hence, the workforce composition is used as instruments in a two-stage least-square (2SLS) estimation.[5]

The empirical results show that the probability of early retirement grows with the tax rate, i.e., with more social costs in relation to the wage sum at the company level. The 2SLS estimate suggests that a 1 percentage point reduction in the tax rate would reduce retirements by about 6 percent. The OLS estimate is also significantly positive, but the coefficient is of a smaller size. The results also imply, given the workforce structure, that the tax rate is positively correlated with net employment growth. One interpretation is that firms that downsize reduce employment among more costly workers first, but it may also suggest wage competition among growing businesses. Further, given the workforce structure, large firms have a higher tax rate than small firms. This suggests that large firms spend, on average, more on non-wage compensation and hence have a more extensive social policy regarding early retirement packages than smaller firms. Firm size has, however, no direct link to early retirement.

The rest of this paper is organized as follows. The next section reviews the relevant literature. Section 2 describes the institutional setting with a focus on the construction of the social costs of labor and the firm's cost of early retirement. Section 3 presents the empirical model. The data are presented in Section 4. The empirical results are presented in Section 5. A concluding discussion is found in Section 6.

[5] There may be a sorting of workers into firms depending on workers' retirement preferences and the type of early retirement coverage (or policy) a firm offers (see Gruber & Madrian, 1995). Generally, this problem may be smaller in Sweden than elsewhere since Sweden has almost universal collective agreement coverage and a high degree of portability of pension entitlements between companies within the same labor market sector (see Section 2 below). However, there may be important differences across industries. By including controls for industry in the analysis, we control for such variation.

2 Earlier literature

Firms' decisions to hire and lay off, or offer early retirement, to older workers have received little attention until recently. In Lazear's (1979) well-known theory a long-term implicit contract is set up between the employee and her employer. It is assumed that monitoring the worker's performance is costly and that there is a compulsory age of retirement when the worker has to leave the firm. To avoid the worker shirking the theory implies deferred earnings relative to the worker's productivity so that workers are underpaid relative to their productivity while they are young and overpaid while old. This would explain why older workers are more costly than young ones (creating incentives for the employer for encouraging early retirement). Feldstein (1976, 1978), Hutchens (1999) and Topel (1984) are often mentioned as pioneers in introducing the concept of the employer's influence on the individual employee's retirement decision. In these papers the individual acts as in a traditional labor-supply model, and decides whether to retire based on the attributes of the alternatives. To some extent the employer can determine the alternatives' attributes by providing early retirement offers. Some scholars (cf. Acemoglu & Angrist, 2001; Behaghel et al., 2005; Hakola & Uusitalo, 2005) have studied how redundancy costs or hiring cutbacks targeted at specific groups (e.g., older workers) affect labor demand, and found evidence that changes in costs or regulations for certain kinds of labor affect the demand for such employees.

The view that retirement is a voluntary choice of the individual may hence give the wrong policy implications. Using cross-national survey data Dorn and Sousa-Poza (2005, 2007) found in close connection to the Hutchens (1999) model that generous early retirement provisions not only make voluntary early retirement more attractive for individuals (pull factor), but also make employers encourage more employees to retire early (push factor). Firms seem to use early retirement as a way to reduce the workforce during economic downturns and as a means to circumvent employment protection legislation.

Coile and Levine (2007) examined how local unemployment affects retirement. They found that retirements only increase in response to an economic downturn when workers become eligible for social security, suggesting that retirement benefits might function as a sort of unemployment insurance. Wadensjö and Sjögren (2000) concluded that early retirement packages became increasingly popular in Sweden during the economic bust of the 1990s, especially in the central government sector. In line with Coile and Levine (2007), Hallberg (2008) found for Sweden – and mainly in the public sector – that aggregated industry employment was negatively related to the probability of early retirement, but also that the replacement level in pension income immediately after early retirement was higher during declining and expanding industry employment than during stable employment. The interpretation could be that employers and employees agree on special early retirement pensions and that these are used in order to persuade older employees to quit voluntarily, but also that they function as rewards for older employees. Also using Swedish data, Eklöf and Hallberg (2010) and Bolin et al. (2008) indicated that special early retirement pensions issued as an extra financial incentive for accepting early retirement could be of substantial importance: the early retirement probability in the age groups 60–64 would be

(depending on the year) reduced by 14–28 percent for males and 7–18 percent for females if no early retirement pensions were issued.

Few papers have looked directly at how non-wage costs may affect early retirement decisions. Plá et al. (2010) found for Spain that a payroll tax reduction had positive effects on wages and on the employment of the targeted group; newly hired workers older than 45 and younger than 30 gained a permanent contract but previously were unemployed. Another type of study has investigated employer-provided insurance (see, e.g., Gruber & Madrian, 1995). Employers may be able to provide the employee with less costly insurance than they could purchase themselves; the employers may even subsidize premiums for their workers, and this is lost in the case of retirement. Johnson et al. (2003) included health insurance costs and benefits in individuals' retirement decision and found as expected that insurance costs significantly reduced retirement rates for older workers.

3 Institutional setting

Sweden is a highly unionized labor market; in 2008 about 71 percent of wage earners were union members (Kjellberg, 2010). Apart from the social insurances in the universal public system, unions and employers' federations have for a long time had collective agreements that grant employees additional social insurances that give supplementary coverage. In 2008 about 90 percent of wage earners were covered by a collective agreement (Kjellberg, 2010). This means that the employer takes additional responsibility for the employee – independently of her union membership status – beyond what is determined in law. These insurances and fees levied on the employer are linked to some form of (normally permanent) employment or employment income earned within a specific labor market sector. Usually the additional social insurance rights are portable in the case of employer switches within the same sector. Some insurance rights (like those for occupational pensions) are also in many cases, but not always, transferable across sectors. Generated pension rights are kept in the case of lost employment. However, the costs for these rights end up unevenly with the last employer, which can explain the low mobility in the Swedish labor market among older workers because of high costs to hire them (cf., e.g., Wadensjö & Sjögren Lindquist, 2005).

The additional insurances exist in a number of areas such as old-age pension, parental leave, unemployment and sickness. Usually they top up the universal national insurance systems with extra insurance above the social insurance cap, but the organization and the type of coverage vary with the type of agreement. In the case of occupational pensions they often give additional possibilities for early retirement with little or no utilization of the national old-age pension before the normal retirement age; see Wadensjö and Lindquist (2005) for details of these insurances.

The social cost is the sum of a) the employment (payroll) tax as regulated by law and b) the collective fees as stipulated by the collective agreements. The employment tax includes contributions to the national social insurance system (e.g., public pensions). The collective fees determined by the collective agreements between unions and employers' federations are administered outside the public system by organizations determined by the parties. Companies affiliated to a collective agreement must pay the negotiated fees and premiums.

The collective agreements, and thus the set-up of fees, vary by labor market sector and sometimes also by industry.[6] As we study the private sector we will concentrate this description on white- and blue-collar workers in the private sector.

[6] There are four main areas for the collective agreements covering four large groups of employees: privately employed white-collar workers, privately employed blue-collar workers, central government employees and local government employees.

Table *1* gives an overview of the current employer fees separately for white-collar workers and blue-collar workers in 2010. The companies that are affiliated to the Confederation of Swedish Enterprise (*Svenskt Näringsliv*), which is the leading organization for private firms in Sweden, pays fees for, e.g., negotiated sickness insurances (AGS), occupational pensions (SAF-LO for blue-collar workers and ITP for white-collar workers), negotiated life insurances (TGL) and work injury insurances (TFA). In addition small fees are paid to the Job Security Council (*Trygghetsrådet*), which helps employees to find new jobs in the case of dismissal due to redundancy, but also plays an important roll in bridging jobs by supplying income replacement. Similar arrangements exist in the public sector. The table shows that while the employment (payroll) tax amount to 31.42 percent for both groups, the fees for the negotiated insurances are substantially higher for white-collar workers than for blue-collar workers, 15.31 percent and 5.11 percent, respectively. To a large extent, this is a result of the different agreements for the two groups of workers mentioned above, in particular regarding the different occupational pension schemes the two groups have.

The occupational pension for blue-collar workers has traditionally given a relatively small supplementary coverage compared with other occupational pension schemes. Their occupational pension scheme gave no or very little extra coverage above 7.5 income base amounts (IB), about 43,000 euro in 2010, which is the social insurance cap above which the national pension system does not cover.[7] The reason is probably that only a small fraction of this group actually had incomes above the cap that needed coverage. The fees that cover the occupational pension for this group (the SAF-LO scheme) are therefore relatively small. Between 1996 and 2007, i.e., during the time period studied in the empirical part of this paper, the SAF-LO scheme meant that blue-collar workers had a fully defined contribution (DC) scheme with a proportional pension premium of 3.5 percent (independently of the social insurance cap).[8]

White-collar workers, on the other hand, are on average better paid and the scheme for this group (ITP) has therefore had a stronger emphasis on supplementary income coverage above the cap (which the national system does not cover). Before 2006, i.e., during the time period studied in this paper, the occupational pension scheme for white-collar workers (henceforth ITP2) was a defined benefit (DB) scheme with a small DC part (ITPK). With the new treaty (hereafter ITP1) that was signed in 2006 the system for white-collar workers will over time turn into a fully DC system. The transition period is long, however.[9]

[7] To be precise, the national pension system and the occupational pensions are indexed by the income base amount (IB), which is tied to the income level. 1 IB was 51,100 SEK (roughly 5,800 euro) in 2010. Other parts of the national social insurance system are indexed by the price level through the base amount (BA, 1 BA was 42,400 SEK in 2010, approximately 4,800 euro). When the IB was introduced in 2001, the IB was almost equal to the BA. Since then incomes have grown faster than prices. Before 2001 pensions were indexed by the BA.

[8] One can note that, over time, this group will be better insured above 7.5 IB. Starting from 2008 the pension premium in the SAF-LO scheme will gradually increase so that the premium in 2012 will be 4.5 percent below 7.5 IB and 30 percent above 7.5 IB. It will remain a DC scheme.

[9] ITP1 is a fully DC pension scheme and applies to white-collar workers born in 1979 or later. For ITP1 the charge is 4.5 percent below the cap and 30 percent above the cap. The old occupational pension system ITP2, which was a fully DB pension scheme, will continue to apply but only to white-collar workers born in 1978 or earlier. Today (2010) ITP1 and ITP2 are in effect in parallel and will be so for a long transition period. As ITP1 only applies to a small share of workers it has had very

The development of the average pension premium for ITP2 during 1998 to 2010 is shown in Figure 1, expressed as percentages of the total annual pensionable wage sum. We note a substantial increase in the average premium from 9 percent to almost 14 percent in 2006, and a slight reduction after that. One reason for the development might be that the workforce during this time was ageing, mainly due to the baby boomers of the 1940s. To a greater extent than before those nearing normal retirement age consisted of well-paid workers with more substantial pension entitlements in the ITP2 system.

The total premium for the DB pension scheme ITP2 has several components: a *proportional premium*, a *fixed-amount premium* and an *individual premium*.[10] The *individual premium* finances a person's DB pension in the ITP2 scheme[11] and is almost always the largest component of the total premium. Since it concerns a DB pension, it is the benefit that the employee is entitled to that defines the size of the individual premium. The benefit is determined from the worker's pensionable income, number of service years and age of retirement.[12] The worker's age can have a large impact on the fee since it affects the number of years that are left until retirement when the benefit has to be fully paid. Also, since ITP2 renders supplementary coverage above the cap, income segments above the cap generate larger fees than income segments below the cap. However, importantly, the amount of the benefit that is pre-paid also defines the size of the individual premium. This means that nominal wage decreases (increases) late in the career may result in substantial cuts (boosts) in the premium. There is a ceiling on how high the individual premium for ITP2 can be. The remainder of the premium above the ceiling is financed collectively through a common premium (the so-called compensation premium).

The ITP pension benefits may be used for early retirement, before the national pension is claimed. In the case of a lower retirement age than 65 the ITP2 premium is higher since early retirement benefits also have to be financed (in such cases the premium ceiling is also raised).

The premiums may vary substantially depending on how the early retirement plan is set up. First, a lower retirement age can be part of the employment contract, i.e., it is *planned early retirement*. In this case the agreement is usually made several years ahead, but this is not a necessary condition. With a longer time left before retirement takes effect the cost can be spread over more years and the premium will therefore be lower per year. The ITP plan specifies in this case rather exact conditions and replacement levels, including actuarial reductions due to early retirement. Second, a lower retirement age can be accomplished through a special early retirement package ('*avgångspension*'), which can be used in the case of *unintended early retirement*. This can occur in the case of redundancy, skill swapping, etc. The exact terms of such contracts can vary between individuals and over time. The nature of the special early retirement package will of course determine the cost for the employer. Rules that

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little impact yet. A company may choose to switch to ITP1 for all its employees, but then special approval is needed.

[10] The *proportional premium* covers the complementary sick pension, premium relief (in case an employee is work absent because of sickness or parental leave), ITPK (the DC part of the pension) and a compensation premium. The *fixed-amount premium* finances TGL (the life insurances), presently 50 SEK per month (about 5 euro).

[11] Besides the old-age pension it also covers the family pension.

[12] In addition the fee also depends on actuarial assumptions regarding, e.g., life expectancy.

determine how the end-payment of pension premiums (the outstanding premiums due to the unintended early retirement) are shared across the collective of employers can make this early retirement option relatively inexpensive. If the employee is 62 or older then the end-payment is levied on the whole group of employers through a collective fee and not charged to the specific employer alone. If the early retirement is made before the age of 62 then the employer alone is responsible for the end-payment of pension premiums.

In Sweden the general principle is to exempt pension saving inflow from tax. Instead the returns on these savings are taxed as well as the pension income that the employee later receives as a retiree (see, e.g., Järliden Bergström, et al., 2010; SOU, 2004). Therefore, savings for a private pension plan made by the employer on behalf of the employee and direct pension payments are subject to favorable tax deduction rules.[13] This means that no legislated employment fees are paid for these labor costs. Instead, these labor costs (direct pension payments and pension premiums) are subject to a reduced payroll tax of 24.26 percent (instead of 31.42 percent).

Hence, from a pure accounting aspect of the firm, a solution involving an early retirement pension can be very favorable compared with a continued employment. Compared with wage payment, the early retirement pension is usually lower, involves a lower payroll tax and is by definition not linked to collective fees. In addition, the end-payment of the premiums – which partly cover the early retirement pension – is in some cases spread over all the employers collectively, i.e., if the early retirement took place at the age of 62 or later. The full cost is hence not borne by the employer alone.

To sum up, the non-wage cost depends (via the occupational pension plan) on the worker's wage, work history (age), the time to retirement and prepaid premiums. In particular, non-wage costs may increase dramatically for wage segments above the social security cap. Therefore, the firm's social cost is – normally – progressive in both the wage and the age composition of the workforce.

[13] This is determined by the supplementary tax rule for companies. One can note that the limits are indexed by the BA. Pension premiums are deductible up to 35 percent of the pensionable income (with a ceiling of 10 BA). One-time payments (end-payments) of premiums that cover, e.g., a special (early) retirement package, are fully tax-deductible conditional that the replacement level in the (early) retirement package is below certain limits. For ages below 65, up to 80 percent of the individual's previous earnings is tax deductible for incomes below the cap of 7.5 BA. The rate is 70 percent for incomes between 7.5 BA and 20 BA, and 40 percent between 20 and 30 BA, while pension costs above 40 BA of the individual's previous earnings are not deductible. At age 65 and above, 20 percent of the individual's previous earnings are deductible for incomes below 7.5 BA, but otherwise the same rates apply as those below age 65.

4 Econometric model

The point of departure is to formulate the individual retirement probability as a function of the share of non-wage costs of the wage sum, i.e., the tax rate, at the firm level. Henceforth, we will denote the tax rate by *TR*. By formulating the retirement probability on the individual level, and not on the aggregate firm level, we are able to hold characteristics of both the individual and the firm constant.

The primary question is related to the *TR*: how much more likely is a worker to retire if employed by a company with a high *TR* compared with a worker employed in a company with a low *TR*, all else being equal? The *TR* is of particular interest as it is a natural policy variable. We hypothesize that a firm with a high *TR* has more incentive to encourage early retirement than a firm with a low *TR*, all else being equal. Given that the employer has an influence on the individual's choice of whether to retire early we expect the *TR* to be positively associated with the retirement probability. Although the employer may be less well informed about each individual worker's exact contribution to the total non-wage cost bill, it is not unreasonable to assume that the employer knows that certain groups of employees are more costly in terms of non-wage costs than others.

To fix ideas we formulate the following regression model:

$$y_{ijt} = \alpha + \lambda TR_{j,t-1} + \boldsymbol{\pi}\mathbf{V}_{j,t-1} + \boldsymbol{\beta}\mathbf{X}_{ijt} + \varepsilon_{ijt} \qquad (1)$$

where y_{ijt} is 1 if the individual *i* employed in company *j* in period *t* retires and 0 otherwise, $TR_{j,t-1}$ is the tax rate in firm *j* in *t*-1, $\mathbf{V}_{j,t\text{-}1}$ is a vector of controls for firm *j* in *t*-1, $\mathbf{X}_{ijt}$ is a vector of controls that describe the characteristics of the individual and ε_{ijt} is an error component.[14] Because we study the transition from work to retirement model (1) applies if $y_{i,j,t-1} = 0$. Our parameter of interest is λ, which is expected to be positive.

As control variables we consider variables that determine the dependent variable and are correlated with the *TR*. Obviously retirement preferences and possibilities to withdraw retirement income vary over age but also over other dimensions. Natural individual-level controls therefore include dummies for age, year, gender and education level. We also control for the type of occupational pension plan (white-collar or blue-collar worker) and industry. Pension entitlements and thus the set-up of fees vary according to the pension scheme, labor market sector and sometimes also the industry. Furthermore, the value of seniority may be very different across industries.

Pension-qualifying income is an important variable for determining future pension entitlements in a DB system. We include a flexible specification of individual pension-qualifying income (where we allow for separate

[14] Note that, to avoid a direct relation between the right-hand side variables and the observed retirement outcome, all the firm-specific controls, including the *TR*, are lagged one period.

coefficients in 10 different income segments). Unfortunately we cannot observe everything that affects pension-qualifying income; in particular, we lack information on work history (however, age may be a good proxy for that). We are also short of information on pre-paid premiums, which may partly determine the size of the individual premium, and information that tells us whether the individual has a lower retirement age as part of the employment contract.

The *TR* is correlated with the whole wage structure in the firm. Since high wages may signal that workers are valuable to the firm and/or that there is a shortage of labor, there might be a positive association between the general wage level and individual retirement. Well-paid workers usually have better jobs with greater flexibility, more enjoyable work tasks and opportunities to transfer to new and more enjoyable positions. The general wage level in a firm may proxy for such characteristics and incentives that these jobs offer for retiring. We control for the general wage level in the firm by including the log of the wage sum per worker. Henceforth we will denote this variable by *w*.[15]

Profitability shocks might affect firms' employment decisions, wage policies and non-wage compensation packages and thus correlate with the *TR*. It is difficult to postulate the direction of these shocks on (voluntary) old-age retirement because firms may compensate for profitability shocks by altering the hiring rate of the young or middle-aged, i.e., there may be reverse causality. The measure of profitability we include is defined as the log of value-added divided by the total labor costs.

With fast technological changes older workers may become redundant because their skills are not updated and the local retirement policies, and thus the *TR*, may be affected as a result of this. We control for investments using gross investments in machinery and equipment per employee (denoted *INV*) as a proxy for technological change.

A company that is expanding might make different trade-offs concerning employment and early retirement compared with contracting firms. We include the annual net employment growth in the regression. It is expected to be negatively correlated with early retirement. If firms act in a cost-minimizing way and reduce employment among the costly workers first, then we expect the net employment growth to be positively correlated with the *TR*. However, a positive correlation may also suggest wage competition among growing businesses.

Big firms may have more options than smaller firms to offer favorable early retirement packages but also greater possibilities for labor hoarding during recessions. Bigger firms should therefore have a higher *TR*, but it is uncertain how retirement depends on firm size.

4.1 Identification issues

Below we discuss two methodological issues involving the interpretation and the identification of our model.

The first issue concerns the interpretation of the parameter of interest (λ). One must be aware that λ represents both a push and a pull factor for retirement. One cannot identify whether retirement was encouraged because workers cost relatively more (push factor) or whether retirement

[15] We use the log in order to reduce the influence of high values.

occurred because individuals had relatively generous early retirement provisions (pull factor). This is a fundamental problem given the data and method used. However, we argue that a substantial part of what we measure in λ is attributed to the push factor and, therefore, that the policy implication of this study mainly concerns the added labor cost of older workers incurred by non-wage costs vis-à-vis younger counterparts. First, we use firm-level variation in the labor cost to explain individual-level behavior in retirement and, in addition, include many controls at the individual level that proxy for several aspects of the pull factor (in particular age, pension-qualifying income, education, etc.). Second, neither the employer nor the employee can, normally, opt out of the collectively agreed pension scheme and negotiate on other terms that are less costly for the employer.[16] This implies that the non-wage cost profile for a particular employee is largely given. To adjust such costs the firm can alter the workforce mix but not the costs for an individual worker employed by the firm. Third, it is reasonable to assume that the financing design of the occupational pensions given by the collective agreements is irrelevant to the retirement preferences of the employee. However, the financing design is most likely to be of significant importance for the employer's incentive to encourage early retirement.

The second issue concerns potential endogeneity bias. There may be a methodological problem since unobserved retirement propensities ('distaste for work') may well be negatively correlated with the earnings levels and, thus, with the *TR* (Coile & Gruber, 2000; Krueger & Pischke, 1992, among others). As mentioned, well-paid employees usually have better jobs with greater flexibility and more enjoyable work tasks. They are presumably a selective group of workers who have greater exchange from work in general. Some of this will spill over to the wage level and the *TR* at the company level. This means that, if a high wage signals a strong preference for work, we expect to see a negative correlation between *TR* and *w* on the one hand and the error term in (1) on the other hand. This will lead to a downward bias in the OLS estimate of λ.[17]

Another reason for suspecting correlation between the error term and the *TR* is that individual *i*:s contribution to the *TR* is not fully captured by the individual-level controls that we are equipped with (as mentioned, we lack information on work history, pre-paid premiums and employment contract terms regarding planned or unplanned early retirement). This may result in an omitted variable bias in the OLS estimate of λ.

In order to eliminate, or reduce, the endogeneity bias we employ an instrumental variable estimator and treat the *TR* and *w* as endogenous. As exogenous variation we use measures of the composition of the workforce in the firm, represented by the shares of workers in a given age group, education level and gender. We denote the instrument set by $\mathbf{Z}_{jt-1}$. An

[16] The *TR* is set by tax legislation and collective agreements. A firm may voluntarily pay additional non-wage compensation (outside the regular collective agreements) for extra pension coverage but this is rare and would normally concern senior managers and executives. This would increase the non-wage costs and thus, if anything, strengthen our argument.

[17] We would thus expect the bias to be negative. Since non-wage costs are more sensitive to high wages than the wage sum we would expect the OLS estimate of *TR* (=*S/W*) to be more biased than that of *w*. However, this depends on the explicit functional relationship between the individual wages in the firm, the *TR* and the wage sum. As described in Section 2 this relationship is highly non-linear, since it relates to, e.g., the social security cap and other institutional details, besides the characteristics of the workforce (wage, work history, expected time to retirement and pre-paid premiums).

instrument in $\mathbf{Z}_{jt-1}$ is thus n^a_{jt-1}/N_{jt-1} , a=1,…,A, where n^a_{jt-1} denotes the number of workers of category a (a given combination of age, education level and gender) and $N_{j,t-1}$ is the total employment in firm j at $t-1$.

For the instruments to be valid we must require that the unobserved individual retirement propensity (conditional on all covariates) is independent of the composition of the workforce in the firm. We argue that the instruments are valid given the Swedish context and the controls included in the model. As pointed out by Gruber and Madrian (1995), among others, workers may select into firms depending on their preferences for (early) retirement pension coverage, i.e., their wish to take early retirement.[18] However, in Sweden it is very unusual for firms to provide their employees with firm-specific pension programs (some minor exceptions may exist for small groups of executives; see footnote 16). This type of selection is therefore small in Sweden, particularly within the same industry, since the collective pension schemes are universal across large parts of the labor market sector and portable, i.e., not linked to a specific employer (see the description of the institutional setting above). Therefore, once the industry is conditioned on the selection into firms is of minor importance.

We will thus assume that the workforce composition measures are valid instruments *within* a specific industry. Between-industry variation in the age, gender and education composition may still be a result of selection.

The workforce is divide into six age groups (ages 16–34, 35–44, 45–49, 50–54, 55–59 and 60+), two education levels (tertiary education level and less than tertiary education level) and gender: in total 24 categories. The share of employees with unknown education is a separate category (altogether resulting in 25 variables that describe the workforce composition). A methodological concern is that many exogenous variables might jointly produce a high partial correlation while in fact only a few of them provide relevant exogenous variation causing an identification problem. There is also an obvious risk that there is multicollinearity in the shares, so using all these variables might be problematic (naturally one must be dropped to avoid perfect correlation). Preliminary analysis shows, as expected, that the relevant variation for explaining the *TR* (and the wage level *w*) mainly comes from the shares of employees with tertiary education (by age and gender): workers with higher education are more important for explaining the *TR* (and the wage level *w*) than the share with less education. Within each age and gender group, tertiary education probably proxies well the groups that are relatively well paid, have relatively large wage segments above the social security cap and thus contribute more to the pension costs of the company. In our baseline estimation we will thus use only the workforce shares that describe the relative elements of tertiary education (by age and gender) as instruments: in total 12 variables.

[18] In the US context it is common with firm-specific pension programs that grant the employee future pension income, which the employee forfeits in the case that she switches to another firm (see Johnson et al., 2003, for a study of retirement and company-specific health insurances). The workforce composition at the firm level may then be a function of unobserved retirement preferences, since the workers select the employer depending on the firm-specific pension scheme and their preferences for retiring early.

The model is fitted as a linear model using two-stage least squares (2SLS). The interpretation of a particular estimated coefficient is therefore the marginal change in the retirement probability. To ensure consistent variance estimates, robust (heteroskedastic-consistent) variance estimates are used.[19]

[19] The estimation was performed by *ivreg2* in Stata 11. Estimation of the model as the instrumental variable probit model (performed by *ivprobit* in Stata 11) gave very similar inference.

5 Data and estimation sample

The data are register data drawn from the database LISA (Longitudinal Integration Database for Health Insurance and Labour Market Studies), Statistics Sweden, and consist of three different parts:

A) Micro data on individuals containing a 3 percent longitudinal sample of individuals aged 50+ for the period 1990–2005.[20] These data contain detailed register information on an annual basis on a wide range of items, such as incomes, demographic situation, education and labor market affiliation.

B) Firm-level data that contain business statistics for companies including all private firms during 1997 until 2005 that had at least one employee. To facilitate the analysis we use data from companies with at least 16 employees. The economic data and data cleaning are sometimes problematic in the case of small businesses. Also, in very small firms the employment decisions are presumably quite special since each employee is central for the business in some sense. Furthermore, job security legislation implemented in 2001 allows companies with ten employees or fewer to exempt two of their employees from the seniority rules of the Swedish job security legislation. The interesting variables include value-added, wage sum, social costs, investments and the number of employees in the firm.[21] *Value-added* is defined in business statistics as production value minus costs for purchased goods and services, excluding salaries and social contributions. *Wage sum* (*W*) includes salaries and other compensation. *Social costs* (*S*) include social contributions (compulsory social security contributions, payroll taxes and other charges) and pension costs (pension payments, pension provisions and pension premiums, etc.).[22, 23] The tax-rate measure for a firm is hence $TR=S/W$.

A few observations (264 out of about 158,000 cases) had social costs that exceeded the wage sum. We have assumed that this is due to some fault in the data and have removed such observations from the analysis.

C) Firm-level data on the composition of the workforce, available between 1990 and 2005. These consist of the number of employees each year by age (in one-year age groups), education level and gender. From these data we create the instrument set mentioned above consisting of the workers in a given category as a share of the firm's total workforce. As mentioned in the methodological section we will use only the workforce shares that describe the relative elements of tertiary education (by six age groups and gender) as instruments in our baseline estimation. In the sensitivity analysis we will, however, experiment with alternative sets of instruments.

[20] In the first year a 3 percent random sample was drawn from all individuals aged 50+. In every following year, 3 percent of individuals entering the target population (50-year-olds and immigrants aged 50+) were added to the panel, and exits (emigrants or deceased) were removed.
[21] One may argue that it would be relevant to examine the role of the economic characteristics at the establishment level instead of at the firm level as employment decisions may be taken at the establishment level. Firms are often more of an administrative unit than a physical one, since firms may change owner, merge or split up. However, the interesting economic variables are only available on the firm level and not on the establishment level.
[22] Other personnel costs (training costs, health care, received grants and allowances for staff and other personnel costs) are not included in the labor costs.
[23] As mentioned, only aggregated social costs data exist at the firm level.

The risk group for retirement in period *t* consists of those working (as defined below) in period *t*-1. From the longitudinal micro data set on individuals aged 50+ (data set A) we thus sample on working status. With these individual–year cases we merge the data from firms and the data on the composition of the workforce (data sets B and C, respectively). The identifier that links an individual to a firm in a given year is based on the person's main employer, i.e., from which the main part of the employment income arises, in November each year.

With annual income data there are some issues relating to the definition of work and retirement. A reduction of work may be partial, possibly combined with work at another company or with partial retirement incomes. We will simplify the analysis by classifying the retirement status from the type of income that dominates. (However, we will return to this assumption in the sensitivity analysis.) The observations are thus classified by highest income from either (a) working, (b) 'voluntary' retirement (public old-age pension, partial retirement, occupational pension and private pension) or (c) other. A prerequisite for working is that earnings amount to at least 1 BA annually (approximately 4800 euros in 2010) and that the working income exceeds all other income categories.[24]

We will only examine the probability of going from (a) work to (b) voluntary retirement. We will assume that joblessness via health-related routes and unemployment is involuntary and therefore omit these types of transitions from our study.

The annual frequency of the data creates some uncertainty as to when retirement occurs. Therefore, the year in which the individual actually goes from work to another type of income source is difficult to classify since it consists of mixes of different types of incomes. We solve this by not using the transition year (i.e., period *t*) in the analysis when classifying work/retirement status. Instead, we analyze the outcome in *t*+1, conditional on being defined as a worker in period *t*-1.

In sum, our micro data will consist of independent three-period panel observations, in which the individual is classified as 'working' in the first period. In the third period the same individual is either classified as 'voluntarily retired' or not.

As a measure of education level we create three dummies: *compulsory education*, *high school education* and *tertiary education*. As a measure of the industry we include 57 dummies that represent the 2-digit level of the SNI code (the Swedish Standard Industrial Classification). We control for the firm size by including dummies for small, medium, big and huge firms (employment sizes 16–70, 71–300, 301–2,000 and 2,001+, respectively).[25] Two important variable imputations were necessary, and a detailed description of how these were performed is found in the appendix. First, the pension-qualifying income may be an important control. For each period we compute pension-qualifying income as the average of the person's last five years of earnings. Second, the type of occupational pension plan a person is affiliated to has to be estimated. Basically the pension plans follow the

[24] The 'other' category consists of benefits from (i) long-term sickness and disability or (ii) unemployment. If the earnings are below 1 BA, and none of the incomes for states (b), (c), (i) or (ii) dominate, the observation is labeled 'inconclusive.' Into this last category fall mainly those with active business incomes or capital incomes.

[25] These limits were chosen to correspond roughly to the firm-size quartiles for firms with at least 16 employees (note that we restrict the study to firms with at least 16 employees). The 25th, 50th and 75th percentile values were 68, 299 and 1,460 employees.

labor market affiliation, for which we have codes in the data. However, these codes lump all privately employed people into one group. Luckily, for retired people we observe the correct plan since we see their source of pensions. Hence, we could estimate the plan for privately employees using a logit and then predict the plan for those not yet retired. The resulting dichotomous variable, *iwhite*, indicates white-collar workers.

5.1 Descriptive statistics

The median *TR* is 43.1 percent in the estimation sample, see Figure 2, which is about what could be expected given the institutional setting. The spread in the estimation sample is, however, quite large. While the 1st percentile is 31.4 percent, the 99th percentile is 75.8 percent. The relatively thick right tail is presumably due to the particular sample used, i.e., firms with more than 15 employees that have at least one worker aged 50+. In Table I in the appendix we report descriptive statistics of the estimation sample. One can note that the firm-level shares of workers with tertiary education are small, especially for older age groups.

6 Empirical results

6.1 Retirement probability estimates

The baseline estimates are reported in Table 2. The 2SLS estimates, in column 1, indicate that the *TR* has a positive effect on early retirement. The effect is significantly different from zero at the 1 percent level.[26] Thus, separation from the firm at an older age through old-age retirement is positively linked with the higher non-wage costs in relation to the wage sum, given the worker's age, gender, education level, pensionable income and characteristics of the firm. According to the point estimate, a 1 percentage point increase in the *TR* increases the retirement probability by 0.14 percentage points. This corresponds to an increase of about 6 percent in the total retirement probability (the average retirement probability is about 2.45 percent in the estimation sample; see Table I in the appendix).[27] This must be judged as quite a large effect.

The naïve OLS regression when all the regressors are assumed exogenous (see col. 2 of the same table) show a much reduced slope estimate of the *TR* (it is about one-sixth of the 2SLS estimate). The coefficient is, nevertheless, statistically different from zero at 1 percent. Further, the 95-percent confidence interval of the OLS estimate of the *TR* is not included in the 95-percent confidence interval of the 2SLS estimate. Thus, there seems to be an endogeneity bias present and our 2SLS estimate differs statistically from the naïve OLS estimate.[28]

[26] The standard error of *TR* is 0.0508 when using a robust (unclustered) variance estimator (p-value 0.005). By instead using standard errors adjusted for intra-group (cluster) correlation on firm level the standard error is slightly higher, 0.0592 (p-value 0.015).

[27] The within-firm standard deviation in the *TR* is 4.75 percent; see Table I in the appendix. Hence, a 1 percentage point change can be viewed as a rather conservative variation and certainly within the inner part of the distribution of *TR*.

[28] Furthermore, Wald tests of exogeneity show that we can reject that the *TR* and *w* are exogenous. One concern is that profitability may be endogenous. However, a test of the suspect orthogonality condition (the C-statistic, or 'difference of the Hansen–Sargan statistic') gives no indication that this is the case. The results are available upon request. As for instrument relevance in the first-stage regressions, the excluded variables have strong relevance as indicated by the high first-stage F-statistics of the excluded instruments shown at the bottom of Table 2, cols. 3 and 4. The LM test statistic reported at the bottom of col. 1 is a test for underidentification (Kleibergen–Paap) and this indicates clearly that the model is not underidentified. (Under the null hypothesis the equation is underidentified, so a rejection of the null indicates that the model is identified. The statistic is distributed as chi-squared with, in our case, 12-2+1=11 degrees of freedom. The critical value on the 0.1 percent level is 31.264.) The Hansen test, which is reported at the bottom of col. 1, is a test of overidentifying restrictions. The joint null hypothesis is that the instruments are valid instruments, i.e., uncorrelated with the error term, and that the excluded instruments are correctly excluded from the estimated equation. In our case Hansen's J-statistic is very low and cannot be rejected on the 50 percent level, indicating that the workforce shares (**Z**) seem to be valid instruments. Hence, the Hansen test gives no indication that the chosen method is inappropriate. However, note, as always, that the Hansen test result should be treated with caution; it is no guarantee that instruments are truly valid since the test is not robust to model specification error in general. Using standard errors adjusted for intra-group (cluster) correlation on firm level do not alter these conclusions; the first-stage F-statistics of the excluded instruments is 18.81 and 36.19 for *TR* and *w*, respectively, the LM test statistic is 37.51 (p-value 0.0001), and Hansen's J statistic is 8.3220 (p-value 0.5974).

As for some of the other covariates, there are some interesting interpretations to be made. There is no support that the log of the wage sum per worker is associated with the likelihood of early retirement (the OLS estimate is however positive). One interpretation might be that high wages cause an income effect that counteracts the initial negative effect that we expected. Interestingly, profitability is not correlated with retirement. As expected, investment in machinery and equipment per worker is positively associated with retirement. However, the relationship is weak and only significantly different from zero at the 5 percent level. There are no statistically significant differences in retirement probability depending on the size of the firm as measured by the dummy variables small, medium (reference), big and huge firm. However, as expected, early retirement and net employment growth are negatively associated. The correlation is strongly significant but seems to be of relatively small economic importance. An employment growth of 5 percent reduces early retirement by 0.024 percentage points (1 percent).[29]

A number of sensitivity checks were performed that generally support our main findings. These are presented in the appendix. In sum, the 2SLS estimate seems relatively stable for using alternative sets of instruments, whether we control for individual characteristics, and the choice of sample, but relatively sensitive to the definition of retirement (partial or full), and the inclusion of some of the control variables: the general wage level, year, industry and age.

6.2 The first-stage regression results

The first-stage regression estimates for the *TR* also lend themselves to exciting analysis. The results in Table 2, col. 3, suggest that larger firms, given the workforce composition, have on average a higher *TR* than smaller firms. Big and huge firms, as they are labeled here (with 300–2,000 employees and more than 2,000 employees, respectively) have about a 2 percentage point higher *TR* than medium-sized firms (71–299 employees) and about a 3 percentage point higher *TR* than small firms (16–70 employees). Bigger firms evidently spend more on non-wage costs (measured as a share of the wage sum) possibly via supplying better early retirement packages to their employees or because they are labor hoarding to bridge economic downturns. The *TR* is also positively correlated with the net employment growth. Firms that reduce employment also reduce their 'tax rate.' This supports an interpretation that firms adjust the mix of their workforce by reducing employment among the more costly workers first. It might also indicate wage competition among growing businesses.[30]

The first-stage regression estimates also confirm our expectation that it is foremost high shares of older, highly educated workers in the workforce that are linked to substantial tax rates. As can be seen, there is quite a large variation in the *TR* across gender and age group for tertiary education. A 1 percentage point increase in the share of men aged 45–54 who have tertiary education corresponds to about a 0.5 percentage point higher *TR*, while an equal increase in the share of men aged 55–59 is

[29] We also find that the retirement probability increases with the pension-qualifying income. One interpretation of the increased retirement probability above the social security cap (of 7.5 BA) could be that these incomes are linked to a high individual premium. However, other interpretations such as varying pension preferences or unobserved private pension savings are possible.

[30] The individual-level control variables mostly show the expected sign. One exception is the positive correlation between *TR* and female. One explanation might be a gender interaction with industry that we have not accounted for.

associated with about a 1 percentage point *TR* increase, compared with the reference group (employees with other education). A corresponding increase in the share of female workers aged 55–59 with tertiary education leads to a 0.7 percentage point higher *TR*.

Changes in the share of older well-educated workers above 60 are either neutral (men) or negative (women) with respect to the *TR*. Explanations for this might be reduced earnings (due to gradual retirement) or selection of who remains in the workforce (notice that the shares aged 60+ are small).[31]

[31] Alternative estimates (not shown) indicate that compositional effects on the *TR* with respect to workers with less than tertiary education are much smaller but, interestingly, there is also a positive age gradient for these groups. For men, the coefficients range from 0.022 to 0.143 for age groups 35–44 and 55–59, respectively, both relative to men aged 16–34 with less than tertiary education.

7 Concluding discussion

From the viewpoint of rising old-age dependency ratios and possibilities to finance pensions and old-age care, it is often argued that older workers need to delay retirement and extend their time in the workforce. Therefore, policies around the world are aimed at increasing the individual's economic incentives to postpone retirement. However, as recent studies suggest, the retirement outcome is not only the individual's choice. The employer and the employee decide mutually on retirement. It is therefore relevant to analyze what makes older workers attractive to the employer. One factor is productivity; empirical evidence – although it is far from a consensus – suggests diminishing productivity over the life cycle, at least beyond the age of 50 (however, the true picture is presumably very heterogeneous across a number of dimensions). Another factor is the labor cost; it is often suggested that older workers cost much more than younger workers in wages and non-wage costs. One major reason why non-wage compensation is high is because older workers are linked to substantial payments towards pension premiums etc. via their occupational pensions. In Sweden these payments are made by the employer as collective fees determined by the rules in the collective agreements, which cover the vast majority of the labor force. Although older workers' high labor costs are an often-used tool in the popular debate, few studies have tried to link such costs to actual retirements.

In this paper we describe the nature of the non-wage costs arising from the collective agreements (which are mainly pension costs) and estimate a statistical relationship between the individual's retirement choice and the employer's labor costs using a large longitudinal employer–employee matched data set that covers the whole private sector in Sweden. We exploit the variation in non-wage costs arising from the fact that, for parts of the workforce, the collective fees are progressive with respect to the employee's age and wage. We attempt to identify the causal implication of the firm's 'tax rate' (non-wage costs as a share of the wage sum) for the retirement choice by assuming that the workforce composition at the firm level is exogenous to the unobserved retirement propensity of the individual.

The instrument variable estimator results show that the probability of retiring early is positively related to the firm's tax rate. The effect must be judged as quite large; a 1 percentage point reduction in the tax rate reduces the retirement probability by 0.14 percentage points or 6 percent. We find furthermore, given the labor cost structure of the firm and other covariates, that there is no additional correlation between shocks in profitability (value-added divided by labor costs) and early retirement.

The regression results indicate, moreover, that big firms spend more on non-wage compensation than small firms do, given the workforce composition. One way that this could arise is that larger companies have a more extensive social policy regarding early retirement packages than smaller firms. Further, the non-wage costs, measured as a share of the wage sum, are as anticipated also positively linked to net employment growth. This supports an interpretation that firms adjust the mix of their workforce by reducing employment among the more costly workers first. It may also suggest wage competition among growing companies.

The results in this paper support that retirement does not seem to be a voluntary one-sided decision by the employee, since firms with a large tax rate on labor, all else being equal, have more retirements. We argue that the policy implications of this effect should mainly concern the incentives of the employer to encourage early retirement. In particular, it concerns the generosity in the rules for early withdrawal, and the low cost for employers of granting early retirement to workers nearing normal retirement, but perhaps mainly the financing design of defined benefit pensions in the occupational pension schemes. In the newly reformed national pension scheme there are several factors that encourage delayed retirement but, obviously, for important parts of the labor force, these may be counteracted by existing occupational pensions. The biggest effects of employer-paid pension premiums are expected for older workers with wages above the social security cap. This is particularly problematic since these workers are often highly educated and have jobs and health that make them particularly suitable for an extended career.

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Tables and figures

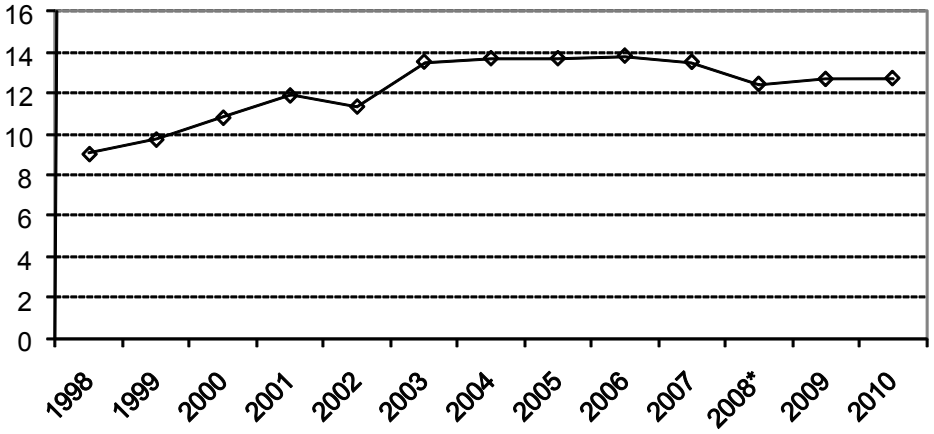

Figure 1. ITP2 premium 1998–2010, expressed as percentages of the total annual pensionable wage sum for all the insured in ITP2 (white-collar workers)

Source: Alecta. The premium for ITPK is included but the negotiated life insurance (TGL) is not. *) In 2008 there was a temporary relief in fees due to a 40 percent reduction in premiums for old-age pension and family pension and a 50 percent reduction in the premium for the risk insurance. This resulted in a net premium of 8.18 percent. The reductions were given to companies that signed a pension agreement with Collectum before December 14, 2007.

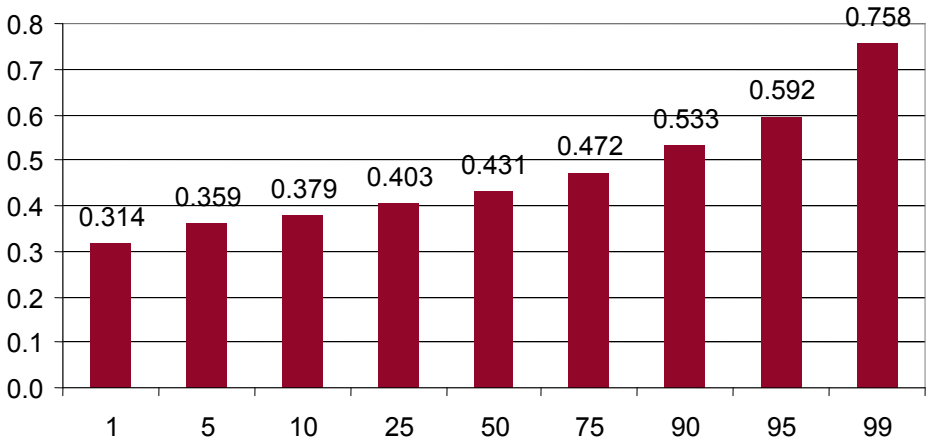

Figure 2. The tax rate (non-wage costs per wage sum) in the estimation sample, by percentile

(author's own calculations)

Table 1. Social costs in 2010, percentage of annual earnings

	Blue-collar workers	*White-collar workers*
Employment fees, legislated		
Old-age pension fee	10.21	10.21
Survivor's pension fee	1.70	1.70
Sickness insurance fee	5.95	5.95
Work injury fee	0.68	0.68
Parental leave insurance fee	2.20	2.20
Labor market fee	4.65	4.65
General wage fee	6.03	6.03
Total, legislated fees	31.42	31.42
Negotiated insurances		
Life insurances (TGL)	0.30	0.20^
Negotiated sickness insurances (AGS)	0.38	
Suppl. labor market adjustment insurance	0.30*	
Suppl. work injury insurance (TFA)	0.01	0.01
Work security council (*Trygghetsrådet*)		0.30^^
Occupational pension		
Pension premium, SAF-LO (blue-collar workers)	3.10**	
Insurance for pension premium exemption, SAF-LO	0.12	
Pension premium, ITP (white-collar workers)		12.0***
Special payroll tax (*särskild löneskatt*)	0.9	2.8
Total fee levied on the employer	*36.53*	*46.73*

Note. Source: The Confederation of Swedish Enterprise, http://www.ekonomifakta.se. *) For companies with a local collective agreement that are not members of the Confederation of Swedish Enterprise the premium is 0.60 percent. **) This is the average fee; in 2010 the fee was 3.1 percent on wage segments below 7.5 IB and 18 percent on wage segments above 7.5 IB (IB is the income base amount, roughly 5800 euro in 2010). Since 2008, the premiums are gradually increasing over time. In 2012 the premiums will be 4.5 percent below 7.5 IB and 30 percent above 7.5 IB. ***) This is a forecast for 2010. This premium is a mix of the new DC scheme (ITP1) and the old DB scheme (ITP2). The ITP1 premium is 4.5 percent below 7.5 IB and 30 percent above 7.5 IB. The ITP2 premium includes the individual premium, the proportional premium and the constant premium, as described in the text. ^) A fixed-amount fee (of generally 50 SEK per month), which is appr. 0.2 percent of the wage sum of white-collar workers. ^^) For companies with a local collective agreement that are not members of the Confederation of Swedish Enterprise the premium is 0.7 percent.

Table 2. Retirement probability: 2SLS estimates and first-stage specifications

	2SLS	*OLS*	*OLS, first-stage*	*OLS, first-stage*
	Retirement	Retirement	*TR(t-1)*	*w(t-1)*
TR(t-1)	0.1437**	0.0224**		
	[0.0441, 0.2433]	[0.0061, 0.0387]		
w(t-1)	-0.0288	0.0094**		
	[-0.0599, 0.0022]	[0.0027, 0.0162]		
Profitability (t-1)	-0.0024	-0.0025	-0.0051***	-0.0220***
INV (t-1)/10^4	0.1125*	0.0575	0.0304	1.5442***
Employment growth (t-1)	-0.0049**	-0.0039**	0.0203***	0.0672***
Small firm (t-1)	0.0013	0.0004	-0.0114***	-0.0066**
Big firm (t-1)	0.0018	0.0033*	0.0179***	0.0188***
Huge firm (t-1)	0.002	0.0053**	0.0257***	-0.0013
I(0<=pred. qual. wage<2)	0.0051	0.0048	-0.0082*	-0.0178
I(2<=pred. qual. wage<4)				
I(4<=pred. qual. wage<6)	0.0017	0.0008	0.0009	0.0241***
I(6<=pred. qual. wage<7.5)	0.0056**	0.0037*	0.0031**	0.0569***
I(7.5<=pred. qual. wage<8)	0.0141***	0.0116***	0.0033*	0.0715***
I(8<=pred. qual. wage<10)	0.0130***	0.0101***	0.0055***	0.0871***
I(10<=pred. qual.	0.0097**	0.0068*	0.0076***	0.0938***
I(12<=pred. qual.	0.0147**	0.0119**	0.0105***	0.1041***
I(14<=pred. qual.	0.0216***	0.0188**	0.0143***	0.1154***
I(16<=pred. qual. wage)	0.0349***	0.0317***	0.0148***	0.1181***
Iwhite	-0.0015	-0.0018	-0.0001	0.0078***
Female	-0.0007	-0.0005	0.0041***	0.0014
Compulsory education	-0.0056***	-0.0055***	-0.0023***	-0.0097***
Tertiary education	0.001	0.0002	-0.0089***	-0.0120***
Workforce shares (**Z**):				
Men 16–34 tertiary (t-1)			-0.0162	0.6175***
Men 35–44 tertiary (t-1)			0.1568***	1.2082***
Men 45–49 tertiary (t-1)			0.5160***	1.3229***
Men 50–54 tertiary (t-1)			0.5610***	0.7291***
Men 55–59 tertiary (t-1)			1.0381***	1.0723***
Men 60+ tertiary (t-1)			0.0556	0.7058***
Women 16–34 tertiary (t-1)			-0.0036	0.4851***
Women 35–44 tertiary (t-1)			0.0146	0.5974***
Women 45–49 tertiary (t-1)			0.1754**	-0.4056*
Women 50–54 tertiary (t-1)			0.0273	0.4551**
Women 55–59 tertiary (t-1)			0.7141***	0.7471***
Women 60+ tertiary (t-1)			-0.7471***	1.6372***
Obs.	55,095	55,095	55,095	55,095
Rank	94	94	104	104

LM test (chi2, 11 d.f.)	666.10		
Hansen J-stat (chi2, 10 d.f.)	9.31		
p-value	0.5028		
F-test, **Z** jointly 0 (prob>F)		253.61 (0.000)	402.15 (0.000)
Partial R^2 of **Z**		0.1138	0.1566

Note. Linear probability model. Significance: * p<0.05; ** p<0.01; *** p<0.001. 95-percent confidence intervals of point estimates in brackets. The predicted qualifying wage is in units of BA. I(x) is 1 if x is fulfilled, 0 otherwise. Models are adjusted for robust standard errors. All models include age, year and industry dummies (2-digit level). The endogenous variables in col. 1 are in bold type. The unique instruments are the lagged workforce shares (**Z**), as they appear in the table. For more details see the text.

Appendix

Calculation of pension-qualifying income

This calculation aims to capture the rules in the pension plan for white-collar workers. It also aims at reducing the within-individual variation in earnings for older workers that comes from labor-supply choices on the intensive margin. According to the pension plan for white-collar workers, the pension-qualifying income (*P*) is normally determined by the 'normal contracted wage income' (including fringe benefits), proportionally reduced if the work history is shorter than 30 years. To adjust for a lower annual wage income in the case of sickness absence we add the sickness insurance income (corrected with the relevant replacement rate) to the wage income. However, no adjustment for work history is made since we lack this information. *P* is then computed as the average of the last five years. However, there is a variation in annual work incomes at ages 60 and above that presumably comes from labor-supply choices on the intensive margin. To reduce this variation, we use predicted values of *P* in the estimation instead of the actual *P*. An individual fixed-effects regression is estimated on the subsample of workers in the age group 50–59 with ln(*P*) as the dependent variable and age interacted with education (compulsory, high school and tertiary) and year dummies as explanatory variables. Using this model we predict the income (also adding the predicted individual fixed effect) for all individuals, independently of their current working status until age 70. This means effectively that individuals will be given the same wage growth rate above age 59 as they have at ages 50–59.

Determination of pension plan affiliation in the private sector

In the private sector, which we analyze in this paper, workers are mainly affiliated to two pension schemes: either that for white-collar workers or that for blue-collar workers. Unfortunately, for privately employed people we have no direct information telling us a person's pension scheme affiliation. However, collected occupational pensions are divided into one of five occupational pension schemes (central, local, white-collar, blue-collar and other), which means that we can identify pension plans for the retired. Using this information on collected pensions we can classify privately employed individuals as either white- or blue-collar workers. For the relatively few observations that received pensions from more than one plan we used the plan with the highest withdrawal.

For the subset of privately employed persons that we observe withdrawing occupational pensions later in the panel we estimate a logit on the probability of being a white-collar worker. Then we use this model to make out-of-sample predictions of the pension plan for the sample that is employed in the private sector but not yet retired. The explanatory variables include education, gender, age, income and owner category of the employer. The predicted dichotomous outcome was determined by the predicted probability from the logit and a random uniform draw from [0,1]. The predicted outcome seems quite convincing; in the estimation sample about 70 percent of all the predictions were correct predictions.

Table I. Descriptive statistics, estimation sample (obs.=55,095)

Variable	*Mean*	*Std Dev.*	*Min.*	*Max.*
Retirement dummy	0.0245	0.1546	0.0000	1.0000
TR(t-1)=S(t-1)/W(t-1)	0.4471	0.0824	0.0000	0.9966
Between-firm variation		0.0670		
Within-firm variation		0.0475		
w(t-1)=ln(W(t-1)/N(t-1))	5.5607	0.2356	3.4012	9.5003
Profitability(t-1)	0.2219	0.2726	-2.0251	2.1046
INV(t-1)/10^4	0.0058	0.0168	-0.0199	2.1104
N(t-1)	1,674.8340	3,268.2710	16.0000	2,1842.0000
Employment growth(t-1)	0.0612	0.4683	-3.6418	5.8645
Small firm(t-1); 16-70 empl.	0.2556	0.4362	0.0000	1.0000
Big firm(t-1); 301-2000 empl.	0.2946	0.4559	0.0000	1.0000
Huge firm(t-1); 2000+ empl.	0.2043	0.4032	0.0000	1.0000
I(0<=pred. qual. wage<2)	0.0079	0.0887	0.0000	1.0000
I(4<=pred. qual. wage<6)	0.2442	0.4296	0.0000	1.0000
I(6<=pred. qual. wage<7.5)	0.2820	0.4500	0.0000	1.0000
I(7.5<=pred. qual. wage<8)	0.0694	0.2542	0.0000	1.0000
I(8<=pred. qual. wage<10)	0.1644	0.3706	0.0000	1.0000
I(10<=pred. qual. wage<12)	0.0745	0.2626	0.0000	1.0000
I(12<=pred. qual. wage<14)	0.0376	0.1902	0.0000	1.0000
I(14<=pred. qual. wage<16)	0.0202	0.1406	0.0000	1.0000
I(16<=pred. qual. wage)	0.0302	0.1711	0.0000	1.0000
Iwhite	0.6273	0.4835	0.0000	1.0000
Female	0.3298	0.4701	0.0000	1.0000
Compulsory education	0.3767	0.4846	0.0000	1.0000
Tertiary education	0.1404	0.3474	0.0000	1.0000
Workforce shares:				
Men 16–34 tertiary (t-1)	0.0174	0.0312	0.0000	0.5714
Men 35–44 tertiary (t-1)	0.0176	0.0260	0.0000	0.3462
Men 45–49 tertiary (t-1)	0.0085	0.0140	0.0000	0.3529
Men 50–54 tertiary (t-1)	0.0091	0.0157	0.0000	0.2778
Men 55–59 tertiary (t-1)	0.0072	0.0139	0.0000	0.3333
Men 60+ tertiary (t-1)	0.0041	0.0099	0.0000	0.2778
Women 16–34 tertiary (t-1)	0.0146	0.0256	0.0000	0.3889
Women 35–44 tertiary (t-1)	0.0105	0.0191	0.0000	0.3333
Women 45–49 tertiary (t-1)	0.0044	0.0116	0.0000	0.3125
Women 50–54 tertiary (t-1)	0.0041	0.0118	0.0000	0.3529
Women 55–59 tertiary (t-1)	0.0032	0.0103	0.0000	0.2432
Women 60+ tertiary (t-1)	0.0016	0.0066	0.0000	0.1667
Men 35–44 other (t-1)	0.1469	0.0822	0.0000	0.6364
Men 45–49 other (t-1)	0.0717	0.0476	0.0000	0.5294
Men 50–54 other (t-1)	0.0809	0.0557	0.0000	0.5000
Men 55–59 other (t-1)	0.0732	0.0544	0.0000	0.5179
Men 60+ other (t-1)	0.0439	0.0388	0.0000	0.5333
Women 16–34 other (t-1)	0.0931	0.1009	0.0000	0.9388
Women 35–44 other (t-1)	0.0670	0.0565	0.0000	0.5238
Women 45–49 other (t-1)	0.0334	0.0326	0.0000	0.3333
Women 50–54 other (t-1)	0.0400	0.0408	0.0000	0.4286
Women 55–59 other (t-1)	0.0347	0.0370	0.0000	0.4706
Women 60+ other (t-1)	0.0196	0.0262	0.0000	0.3478
Unknown education (t-1)	0.0037	0.0092	0.0000	0.5000

Note. I(x) is 1 if x is fulfilled, 0 otherwise. Prediction of the pension-qualifying income (pred. qual. wage) is in units of BA. *Tertiary* means at least university education and *other* means other education. *INV* is gross investments in machinery and equipment per employee. Profitability(t-1) is ln((VA(t-1)/(W(t-1)+S(t-1)), *VA* is value-added. Employment growth(t-1) is ln(N(t-1))-ln(N(t-2)), *N* is firm size.

Sensitivity analysis

First, since the individual may gradually retire and the data used are imperfect in determining the exact retirement date, we experimented with both a softer and a harder classification when someone became retired (depending on the share of the total income that was pension income). These regressions, presented in Table II, show that the 2SLS estimate of *TR* is about the same as the baseline estimate if retirement was 'soft' (defined as having at least one-quarter of the total income from pensions). The 2SLS estimate is substantially reduced compared with the baseline estimate and is no longer significantly different from zero if we require that retirement means having at least three-quarters of the income coming from pensions. One interpretation is that the positive effect of the *TR* on early retirements mainly concerns gradual withdrawals from the labor market and less so the definitive retirements.

Second, considering that we may have an identification problem when using many instruments, and also possibly multicollinearity in the instruments, we re-estimated the model using alternative sets of instruments. However, these results, included in Table III, indicate that our specification is robust to alternative sets of instruments.

Third, we experimented with alternative sets of controls. The results are presented in Table IV. We find that the results are relatively unchanged if we remove the firm controls (profitability, investments, company size, employment growth, dummies for small, big and huge firm) and individual controls (qualifying wage, white-collar worker, female, education level). However, if we remove the controls for year, industry or age dummies, then Hansen's J-statistic (test of instrument validity) is rejected. This may indicate problems of instrument validity due to the selection process discussed above. If we mix firms without controlling for the type of industry they belong to, we will not control for the between-industry variation that we think exists in both the workforce composition and the unobserved retirement preferences. However, this is just one possible interpretation since the test result might also be an indication of a general model specification error. Additional sensitivity checks, included in Table V, concern the age restriction of the sample and further variation of the model specification. One may argue that individuals who retire very early (before 55) or very late in life (after 65) are special in some sense. We therefore limit the estimation sample to ages 55–64 (thus dropping about 37 percent who are younger than 55 and 0.7 percent who are older than 64). The age restriction has little effect on our key parameter, compared with the baseline model (col. 1). If we remove individual controls (except for age and year), col. 2, the parameter estimate of *TR* is also relatively unchanged compared with the baseline model (col. 1). To judge the influence of other firm characteristics, cols. 3 and 4 show estimates with *w* removed from the specification. In both cases the 2SLS estimate of *TR* is reduced by about one-third compared with the model in col. 1.

Table II. Retirement probability using alternative definitions

Variable	*Ret1/4*	*Ret1/4*	*Ret3/4*	*Ret3/4*
	2SLS	OLS	2SLS	OLS
TR(t-1)	0.1554*	0.0455***	0.0531	0.0260***
w(t-1)	0.0020	0.0097*	-0.0010	0.0083***
Obs.	55,095	55,095	55,095	55,095
Hansen J-stat (chi2, 10 d.f.)	16.65		9.47	
p-value	0.0826		0.4879	

Note. Significance: * p<0.05; ** p<0.01; *** p<0.001. Models are adjusted for robust standard errors. Endogenous variables (in the case of 2SLS) are in bold. The dependent variable is *Ret1/4* in cols. 1 and 2 and is *Ret3/4* in cols. 3 and 4. *Retx* is 1 if pension income/total income>*x*, 0 otherwise, where *x* is either ¼ or ¾. The model specification is otherwise exactly the same as in Table 2, cols. 1 and 2.

Table III. Using an alternative set of unique instruments, 2SLS estimates

	Full set of instruments	*Only males' shares as instruments*	*Only females' shares as instruments*	*Reduced set of instruments*
Variable	2SLS	2SLS	2SLS	2SLS
TR(t-1)	0.1561***	0.1367**	0.1588**	0.1394**
w(t-1)	-0.0273*	-0.0190	-0.0322*	-0.0245
Obs.	55,095	55,095	55,095	55,095
Rank	102	102	102	102
LM test (chi2, 23 d.f.)	1,253.62	985.69	552.70	977.68
Hansen J-stat	27.73	14.91	12.62	14.55
p-value	0.1848	0.1353	0.2458	0.2041

Note. Significance: * p<0.05; ** p<0.01; *** p<0.001. Models are adjusted for robust standard errors. Endogenous variables are in bold type. The unique instruments (Z) in col. 1 are the full set of instruments, i.e., gender-specific shares of employees in age groups 16–34, 35–44, 45–49, 50–54, 55–59 and 60+, interacted with education group (tertiary education and other), all lagged one period (one is exempted to avoid perfect collinearity). In col. 2, Z includes the shares for male employees only (by age and educational group). In col. 3, Z includes instead the shares for females. In col. 3 only the age groups 16–34, 45–49 and 55–59 are included in Z (by education group and gender). (In cols. 1, 2, 3, 4, there are thus 24, 12, 12, 13 instruments, respectively.) Additional controls in all the models include firm profitability, investments, employment growth, firm size (in level and three dummies), individual pension-qualifying income, Iwhite, individual pension-qualifying income interacted with Iwhite, female, dummies for age, year and industry.

Table IV. Retirement probability, 2SLS estimates, alternative sets of controls

Variable	*2SLS*	*2SLS*	*2SLS*	*2SLS*	*2SLS*	*2SLS*	*2SLS*	*2SLS*	*2SLS*
TR(t-1)	0.0330	0.1015**	0.1322***	0.0843**	0.1119**	0.1330***	0.1565***	0.1582***	0.1502***
w(t-1)	0.0174*	0.0031	-0.0146	0.0122	-0.0033	-0.0277*	-0.0092	-0.0302**	-0.0302*
Firm controls 1									x
Firm controls 2							x	x	x
Individual controls						x		x	x
Year dummies				x	x	x	x	x	x
Industry dummies			x		x	x	x	x	x
Age dummies		X	x	x	x	x	x	x	x
Obs.	55,095	55,095	55,095	55,095	55,095	55,095	55,095	55,095	55,095
Rank	3	19	69	25	75	88	79	92	95
LM test (chi2, 23 d.f.)	1,424.91	1,424.22	1,593.78	1,391.55	1,452.91	1,390.02	1,301.52	1,252.28	1,263.36
Hansen J-stat	153.59	52.07	36.60	42.68	30.26	29.15	26.75	27.20	27.34
p-value	0.0000	0.0003	0.0262	0.0052	0.1122	0.1406	0.2210	0.2036	0.1987

Note. Significance: * p<0.05; ** p<0.01; *** p<0.001. Models are adjusted for robust standard errors. Endogenous variables are in bold. The unique instruments are the full set of instruments, i.e., gender-specific shares of employees in age groups 16–34, 35–44, 45–49, 50–54, 55–59 and 60+, interacted with education group (tertiary education and other), all lagged one period (24 instruments). Firm controls 1 includes controls for lagged profitability, investments and company size. Firm controls 2 includes controls for lagged employment growth and dummies for company size (small, big, huge firm). Individual controls include dummies for qualifying wage in 9 different income segments, white-collar worker, female and education level (tertiary, high school and compulsory).

Table V. Sensitivity checks (only individuals aged 55–64), 2SLS

Variable	*2SLS*	*2SLS*	*2SLS*	*2SLS*
TR(t-1)	0.1655**	0.1814**	0.1158**	0.0977**
w(t-1)	-0.0205	0.0019		
Controls:				
Firm controls 1	X	X	X	
Firm controls 2	X		X	
Individual qual. wage	X		X	X
Gender and sector	X		X	X
Obs.	34,086	36,373	34,086	34,095
Rank	85	76	84	77
LM test	770.21	839.95	2,212.33	2,138.27
Hansen J-stat	23.12	25.80	24.17	24.07
p-value	0.3951	0.2602	0.3946	0.3998

Note. Significance: * p<0.05; ** p<0.01; *** p<0.001. Models are adjusted for robust standard errors. Only individuals aged 55–64. All models include age, year and industry dummies (2-digit level). Endogenous variables are in bold. The unique instruments are the full set of workforce shares, i.e., gender-specific shares of employees in age groups 16–34, 35–44, 45–49, 50–54, 55–59 and 60+, interacted with education group (tertiary education and other), all lagged one period (24 instruments). Firm controls 1 includes lagged profitability, investments, firm size and employment growth. Firm controls 2 includes lagged dummies for company size (small, big, huge firm). See Table 2 and text for more details.

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